Labor Markets and Institutions

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Labor Markets and Institutions: An Overview

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The importance of the labor market is indisputable. The countries' economic outcomes rely to a significant extent on its performance, as production, economic growth, and prices are all intimately linked with it. Moreover, the functioning of the labor market is a key determinant of social welfare. Elements such as unemployment and its duration, job quality, wages and compensations, greatly influence individual well being, and are in turn greatly influenced by the performance of the labor market and its institutions. Thus, labor markets and their response to shocks or changing conditions, the functioning of labor institutions under different scenarios, the ways in which they can be modified to improve their efficiency and the endogeneity of suboptimal institutional arrangements, are some of the very significant issues that are analyzed in the collection of studies included in this book. Surely, it will contribute to a better understanding of this market by proposing new empirical evaluation approaches, and will enlighten the policy discussion by suggesting ways to improve the design of labor regulations and institutions.

Regulations and institutions govern labor markets all over the world. Regulations are more stringent in some markets than in others, but labor markets that are allowed to freely allocate resources with no intervention at all are hard to find. Every day, firms must abide by a labor code when hiring or firing people or when determining work loads, schedules, and other conditions.

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The pervasiveness of labor regulations has led people to question their very existence and wonder what the world would be like without them. How would macro- and microeconomic variables respond to shocks in the absence of labor market regulations? Questions of this nature—together with issues related to the specific effects of regulations on different outcome variables, the optimal design of regulations, and the motivations behind them—drive the authors of all the papers presented in this volume. The recent theoretical and empirical research reported here contributes to a better understanding of the functioning of labor markets and thus offers important evidence for economic debates on themes as varied as growth, inequality, poverty, unemployment, and human capital.

The field of labor economics can be simplified into two basic lines of thought, which offer alternative explanations as to why regulations came to life and how they managed to subsist in the modern world. Economists in the first group, often called distortionists, believe that institutions are born because of pressure from rent seekers—mostly employed workers who benefit from the regulations they push, causing inequality—and that they produce adverse economic effects by driving the market away from its supposedly optimal laissez-faire position. Researchers adhering to this approach regard regulations as the creators of distorted incentives that misguide economic agents' behavior, which results in inefficient resource allocation. The best the society can do is to remove all regulations and let the market function freely, because the best possible outcome will be achieved when each agent looks out for its own benefit.

The second group, frequently called institutionalists, contends that regulations originate as a response to market failures, which lead to suboptimal outcomes. These economists claim that labor institutions pursue an efficiency objective and not just a redistributive one. Properly designed and implemented regulations are believed to play a role in moving the market toward a better equilibrium that Pareto–dominates the laissez-faire position. Contrary to the distortionist view, institutionalists assume that it is not always desirable to remove existing regulations. Deregulatory movements may be fully justifiable in many circumstances, however, including changes in the nature of the market failure or in the way agents and institutions interact.

It is not difficult to think of problems associated with the labor market that could be an impediment to its efficient behavior. The typical arguments used to justify intervention on efficiency grounds are that markets are incomplete and imperfect, cause external effects, and have some characteristics of a public good. All of these failures are easily found in most labor markets, at some point in time. One case of market incompleteness, for example, occurs when private insurers are unable to fully insure the labor income of risk-averse workers. Unemployed workers are then incapable of smoothing consumption, and they suffer a significant welfare loss as a result. Worker allocation among jobs is suboptimal in this case, as workers tend to stay in stressful jobs or accept an offer that is not their best match, just to avoid the income loss during the transition or the uncertainties of an employment change. Employers' behavior may also be affected, in that they will probably fire too often if they do not consider the cost imposed on the redundant worker. This inefficiently functioning laissez-faire market will trigger the design of institutions to improve its behavior. Common regulations and institutions for addressing this specific failure are unemployment insurance, severance payments, job banks, and search agencies.

The classic example of market imperfection is that of a labor market consisting of only one employer that behaves monopsonistically when hiring its employees. This would be the case, for instance, of a one-firm town. This unregulated market generates a wage and a number of hired workers that are both lower than they would have been had a benevolent social planner solved the equilibrium. Consequently, regulations like minimum wages are introduced as a means of forcing the monopsonist to behave in a competitive manner.

Finally, conditions at the workplace, such as comfort and security, are examples of nonrival and nonexcludable goods, and they can therefore be classified as public goods. Workers are inclined to free ride on the effort of others to attain better conditions, thus obtaining less than is socially optimal given the total value assigned to them. Worker organizations, such as unions, serve a purpose in this specific case, too.

All these regulations that are intended to solve a specific market distortion and ultimately help people often end up doing neither. Generous unemployment benefits may provide a good safety net for those who lose their jobs, but they also imply costs in the form of higher taxes or lower wages to finance the system and distorted search incentives for the unemployed. High severance payments discourage firms not only from firing, but also from hiring workers. High minimum wages in a non-monopsonistic market improve the welfare of some workers, but they lower overall employment and raise unemployment among the young and unskilled.

Thus, although the regulatory framework can bring about benefits, success is not guaranteed. Many elements play a role in determining their final outcome, including the type and severity of the market failure, the design and implementation of the specific regulation, the degree of compliance, and interaction with agents and other regulations. Even when a regulation is correctly designed for the distortion in the market at a given time and then effectively implemented, changing conditions may render that regulation useless or even counterproductive at a later time, and it may not apply at all under different circumstances or in a different country. It is therefore essential to design adaptable regulations that can change with evolving conditions.

Regulations are not bad per se and need not be abolished altogether. Some should be eliminated, however, as they do not serve their purpose, either because they were badly designed or implemented or because conditions have changed, making the regulations unfruitful or even perverse. A modern view of the regulatory role thus combines both the institutionalist and distortionist perspectives. It starts from the premise that many labor markets today ache from market failures of varying degrees and natures that inhibit them from functioning well without proper regulation. It recognizes that many of these regulations are badly designed and require major makeovers or even elimination, as they hurt more than they help. Moreover, the selection of specific regulations and their design should be closely tailored to the precise problem observed in the market, since one size does not fit all.

This proposition raises several questions. How did the labor market regulations currently in place come into being? What are the effects of today's regulations? What should optimal labor regulations look like? The papers compiled in this book represent steps forward in answering these questions. We address them one at a time below.

1. The Origins of Current Labor Market Regulations

Interventions are justified on efficiency grounds when they respond to some market failure that inhibits the market from arriving at a competitive equilibrium. Thus, if we expect the regulations to be driven by concerns for efficiency, we should expect their design to follow a careful study of the market, its flaws, and options for solving them. This is typically not the case. More often than not, regulations are promoted by interest groups, whose influence depends on a country's social and political structure.[1] Similarly, well-intentioned policymakers frequently dictate rules to help specific groups of people, only to end up hurting them or others.

The stylized model presented by Giuseppe Bertola (in this volume) highlights additional mechanisms that motivate collective interventions aimed at altering laissez-faire wage and employment outcomes. These are tested both in a competitive labor market and in a market that fails to achieve its optimum because of mobility costs and limited access to financial markets. The key insight is that workers do not hold a proportion of all production factors as the representative agent. Consequently, workers maximize a welfare function, which leads to a situation that departs from the competitive market outcome. In this case, the portion of total production that does not accrue to them in the form of wages is ignored. Therefore, workers may push for a reduction in employment, since it has a negligible negative effect at the point at which the wage equals nonemployment opportunities, but a first-order beneficial effect on the still-employed workers' welfare.

So, regulations are not necessarily implemented to maximize production, employment, or welfare. They often obey specific interests and may result in outcomes that are even worse than those produced by the imperfect market in the first place. Implementing them is not free of cost; it requires substantial information acquisition and interpretation. Monitoring is not cheap, either. As Bertola points out, countries will therefore create different kinds of regulations to confront the same market distortion, in order to best accommodate their information processing, monitoring, and enforcement capacities. Moreover, the same regulations will have different effects depending on the characteristics of the economy in which they are applied.

2. The Effects of Today's Labor Market Regulations

Independently of the motivation behind regulations, most of the relevant literature tries to estimate their effect on specific variables of interest, such as measures of aggregate and disaggregate labor market outcome or macroeconomic variables and trends. Not surprisingly, the estimated effects of different regulations, in different countries at different times, are varied. Common sense indicates that changing circumstances make each country's regulations more or less suitable

1. See Saint-Paul (2000).

to pursue their objectives. Bertola presents simple descriptive analyses for countries in the Organization for Economic Cooperation and Development (OECD) and in Latin America that show that the effects of given regulations do vary in a changing world.[2] For instance, the cross-country evidence has established that the same institutions were associated with very different outcomes in the 1960s and 1980s. In general, the evidence presented by Bertola for OECD countries shows that job security provisions are detrimental to job creation. High layoff costs reduce both hiring and firing, thereby lowering turnover and, with it, employment volatility. The economic literature is not conclusive regarding the overall effect on average employment, since it depends on the specifics of each economy.

Nevertheless, cross-country comparisons do not say much about the effects of regulations because such effects are regularly endogenous. As Agell (2002) puts it, we do not know "what comes first, the chicken or the egg." Most unemployment protection regulations are present in high-unemployment countries—as a consequence, but also as a cause of the unemployment (see Bertola, in this volume). This endogeneity complicates cross-country comparisons because the real effects of the regulatory framework must be disentangled from the structural features that led to them.

Additional problems arise in such comparisons because the effects of regulations vary with other variables normally excluded from the analysis, such as the type of market failure, the intensity and nature of shocks that hit the economies, and the initial conditions of the market. This last variable is shown to be important by Juan J. Dolado, Marcel Jansen, and Juan F. Jimeno (in this volume), who model state-contingent effects of partial reforms on various labor market outcomes based on a search equilibrium model. What they find corroborates the intuitive idea that the effects will vary according to the circumstances and market conditions.

For instance, a reduction in layoff costs targeted to less productive workers in sclerotic labor markets may reduce low-skilled workers' unemployment without affecting the unemployment rate of high-skilled workers, and it may increase the wage and welfare of both categories. In tight labor markets, however, such a policy often increases unemployment among low-productivity workers and has little effect on the unemployment rate of high-productivity workers.

2. Bertola, Blau, and Kahn (2002) and Blanchard and Wolfers (2000) arrive at similar conclusions.

The welfare of low-productivity workers typically falls, while the welfare of the more productive ones increases. Obviously, the policy is much less politically feasible in this second scenario than in the first.

According to Heckman and Pagés (2004) the aforementioned problems—endogeneity and variation in effects depending on initial conditions—explain why the evidence for the OECD is not conclusive on the effects of regulations. They claim that research on Latin American countries should be more fruitful in that endogeneity is less of an issue, since most reforms followed big political regime changes and thus were mostly exogenous. Their findings are open to controversy, however. This, Bertola says, is due to the low quality of the data on Latin American labor markets, which resemble the data available for the OECD fifteen years ago. Another serious problem involves defining how to measure the regulatory framework. Should one use an aggregate index of regulations, or some degree of disaggregation? How should one capture the fact that the same regulation is more stringent in one country than in other? For example, an accurate analysis must consider not only the existence of minimum wages, but also their level and enforcement. Finally, one must capture both the aggregate effect of a regulation and its distribution among workers.

Claudio Montenegro and Carmen Pagés (in this volume) address the issue of the distributive impact of labor regulation. They use the large variation in labor market regulations experienced in Chile since the 1960s to analyze the impact of the regulations on the distribution of employment across age, gender, and skill level. They consider the effects of the total costs of dismissing workers, including advance notice, severance payments for each tenure level, and the probability that a firm's economic difficulties serve as justification for termination. They also consider the effects of the minimum wage on employment distribution. Their detailed analysis confirms that young, low-skilled workers and women are the most hurt, although an increase in the minimum wage seems to benefit women.

Nevertheless, it is not only the aggregate and distributional effects of regulation that matter, but the way they may interact with other variables is also a concern to many policymakers. Pierre-Richard Agénor (in this volume) makes an appealing point in describing—analytically and empirically—the potential tradeoff between institutional changes oriented at alleviating poverty and those intended to reduce unemployment, such as payroll tax cuts for unskilled labor and reductions in minimum wages or severance costs.

The chapters by César Calderón and Alberto Chong and by César Calderón, Alberto Chong and Rodrigo Valdés (in this volume) further contribute, in a number of ways, to the literature on the effects of the regulatory framework of the labor market. First, they focus on two macroeconomic variables that accurately signal the total effect of regulations: growth and inequality. If a regulation's aim is to correct an inefficiently achieved equilibrium in an imperfect labor market, it should enhance efficiency and thus produce a growth effect. If the regulation has a distributive rationale, then it should result in changes in overall inequality. Moreover, these two variables not only capture the initial motivation for the regulation, but also reflect its aggregate effects quite concisely.

A second important contribution is the way the authors treat regulations. They incorporate two measures—one that reveals the amount of regulation (their de jure or "thickness-of-the-code-book" measure) and one that reflects the stringency of the regulation and the degree to which it is enforced (their de facto measure). Additionally, regulations are not added up; rather, each is introduced individually, as they may and will have different effects on the outcome variables. They find that some regulations deter growth (for example, minimum wages and trade unions), but not all of them produce a significant negative effect. Moreover, the effects are extremely small, requiring drastic regulatory changes to produce a modest increase in growth. The authors also find that inequality is negatively affected by some regulations, specifically by minimum wages. Other labor regulations, such as unionization, maternal leave, and government employment plans, seem to improve income distribution.

The conclusions of these two chapters are intuitively appealing, in that de facto regulation dominates de jure legislation when explaining labor market outcomes. Also, adding regulations up in a single index conceals information, because the effects of each legislative statute vary. Finally, in most cases, the effects found are very small.

Another alternative to the traditional approach of assessing directly policy measures is to analyze quantitatively the actual performance of the market. For that purpose, they could either use micro or macroeconomic data to build indicators to look whether the market is farther from its competitive equilibrium than markets in other countries or to compare the behavior of the market returning to the equilibrium after being disturbed.

Gilles Saint-Paul (in this volume) builds indicators of wage rents, as proxy for labor market competition, to evaluate the impact of recent labor reforms in European countries. On analyzing the evolution of within-industry wage differentials and welfare differences between employed and unemployed workers, he finds that contrary to popular belief, most European countries are neither more nor less competitive than in earlier periods.

A technique designed to determine the relative flexibility of labor markets is chosen and applied, using macroeconomic data from a sample of OECD and emerging markets, in Elías Albagli, Pablo García and Jorge Restrepo (in this volume) and, using plant-level data from five Latin American countries, in Ricardo Caballero, Eduardo Engel and Alejandro Micco (in this volume).

When evaluating flexibility, these two chapters introduce alternatives to the traditional approach of directly measuring labor regulation (typically de jure regulation) in the style of the work done at the OECD and Heckman and Pagés (2004). Instead, they measure labor market flexibility by the way the market is functioning. Albagli, García, and Restrepo do so at the macroeconomic level, while Caballero, Engel, and Micco focus on microeconomic plant-level data.

Measuring de facto rigidity in labor markets requires first distinguishing its source. What is perceived as labor market rigidity in the raw data may actually be the economy's response to a sequence of negative shocks over a given period of time. In addition, the response of macroeconomic variables will be different in the presence of labor market rigidity. For instance, the more rigid the labor market is, the longer unemployment will last. What is called for is identifying and disentangling the set of structural shocks driving the economy, which can be done by estimating a structural vector autoregression (SVAR). Albagli, García, and Restrepo use this econometric technique to identify four structural shocks, based on an open economy model, with a supply side and wage bargaining à la Blanchard and Summers (1986). They use this methodology to compute a direct measure of labor market rigidity—namely, unemployment persistence—for a heterogeneous sample of countries.[3] They assess each country's performance by its responses to identified structural shocks.

This paper's contribution over previous work on this line of research is, first, to extend the model to small open economies; this is a natural way of proceeding, since many of the countries in the authors' sample are small and open. Second, their index is comparable

3. Balmaseda, Dolado, and López-Salido (2000), Dolado and Jimeno (1997), and Fabiani and Rodriguez-Palenzuela (2001) are examples of the same approach for closed economies.

across emerging and OECD economies. Their index of the half-life of unemployment—after the economy is hit by a shock—depends exclusively on the rigidity coefficient in the wage-bargaining equation of the model. In contrast, some of the rigidity indices found in the related literature depend not only on this rigidity parameter, but also on other structural parameters. This does not represent a serious problem if the economies being ranked are similar, but it can lead to misinterpretation of results when the sample includes countries with heterogeneous levels of development and openness. The authors find that Chile, Hong Kong, Korea, and the United States rank among the most flexible labor markets, while Colombia, Germany, Spain, and Sweden stand among the most rigid.

Caballero, Engel, and Micco, in turn, follow a different approach to assessing labor market flexibility. Based on earlier work by Caballero and Engel (1993) and Caballero, Engel, and Haltiwanger (1997), they measure and compare microeconomic flexibility by estimating the speed at which establishments close the gap between labor productivity and the marginal cost of labor.[4] Their methodology is derived from an adjustment hazard model in which a change in the employee headcount at any plant is a probabilistic function of the gap between desired and actual employment. The model allows for nonlinearities (lumpy adjustments) and state-dependent responses. This strategy had never before been applied to Latin American countries.

The authors estimate that Brazil, Chile, and Colombia are more flexible than Mexico and Venezuela. A detailed analysis of the Chilean case identifies signs of a downward trend in microeconomic flexibility after the 1998 crisis. This conclusion points in the same direction as Heckman and Pagés (2004).

3. Optimal Labor Regulations

Labor regulations and institutions undoubtedly affect the dynamics of important variables of the labor market. The impact varies depending on circumstances and design. Any perverse effect of a regulation does not necessarily imply that no regulation is needed, but rather that the specific regulation has not been designed or implemented efficiently to address a specific market failure at a specific

4. Caballero, Engel, and Haltiwanger (1997) reproduce the aggregate dynamics of the U.S. labor market using data from a large set of individual firms.

point in time. In this sense, we lack analysis on what optimal regulations should look like.

Identifying the market failure and its effects is first and foremost, in order to evaluate alternative regulations and define policies that best solve the specific problem. The regulation's implementation and design depend crucially on the market conditions under which it will be applied. The ongoing research of Blanchard and Tirole (2003) in optimal labor market institution design is probably one of the most recent efforts of modern economics in this field. It is particularly relevant because it presents a concrete analysis of a specific market failure, its effects, and the detailed design of regulations that solve the problem. It also recognizes how other institutions and interactions between agents and the market modify the way in which the regulation should be designed.

Based on that joint research, Olivier Blanchard proposes a system in which unemployment insurance and employment protection, in the form of layoff taxes, coexist, and in which their specific design is molded by complications that may arise from the interaction in the market between firms and workers. The identified failure is the workers' limited access to financial markets. Thus, he works through a detailed design of unemployment insurance intended to diminish the negative utility effect on dismissed workers and to make firms internalize layoff costs, trying to have the appropriate incentives to avoid misbehavior. Therefore, he examines the effects of labor regulations on hiring, firing, and job search decisions in markets where workers are not fully insured against changes in income.

Blanchard considers elements such as how the insurance amount, periodicity, and administration affect firms' firing and hiring decisions and workers' job search and acceptance decisions. He also introduces ex post wage bargaining power and worker heterogeneity into the analysis to build concrete proposals on how the design should adjust to them. For example, he proposes limiting payments if the fact that workers can collect income while unemployed diminishes search incentives. While one size does not fit all, certain elements of markets can be analyzed and translated into specifically designed regulations that work efficiently to solve an identified market failure. The chapter works through several of them, but in the end each country's authority must consider all the possible complications and interactions to arrive at an optimal regulatory design.

In the specific case presented by Blanchard, the problem arises from the inability to insure workers' income in bad states of nature, such as unemployment spells, during which income drops drastically and unexpectedly. Consequently, the probability of falling into a bad state is a crucial element for the design of the optimal regulation. When this probability is high, the usefulness and necessity of unemployment protection increases from the workers' perspective. This happens, in particular, when markets are not complete and risks are not covered, meaning that consumption cannot be insulated from income shocks.[5] It is therefore critical to study the evolution of workers' income during their life cycle. Cristóbal Huneeus and Andrea Repetto (in this volume) make a remarkable empirical contribution to the knowledge of the earnings process for the specific case of Chile. They find that, as in other countries, adverse income shocks as a result of unemployment spells are more likely to occur as workers become older, but they are less likely to happen at all in Chile than in the United States. Their empirical analysis also confirms that government transfers in Chile have little effect on consumers' ability to compensate persistent shifts in their earnings. They show that the Chilean distribution of income is highly persistent, which is explained by the underlying variability of the earnings process. One conclusion is that it is much harder for Chilean than U.S. consumers to smooth consumption and move to a higher income quintile. As Huneeus and Repetto put it, the welfare consequences of income uncertainty are high in a developing economy, where the public welfare system is small and consumers cannot share risks or are liquidity constrained. These findings should be taken into account in the design of regulations in Chile and in the analysis of cross-country differences in institutional arrangements.

Another element that interacts with labor market imperfections and regulations and should be considered when assessing and designing labor market policies is the long–run trend in the demand for skilled workers. In Chile, as in other emerging economies, the demand for skills has grown more than proportionally, in association with international trade, as confirmed by Olga Fuentes and Simon Gilchrist (in this volume).[6] Fuentes and Gilchrist's analysis covers

5. Dynarski and Gruber (1997) show that even in a developed economy like the United States, income risks related to unemployment are not pooled, so families' consumption is not perfectly insured. Smoothing consumption could also be in the interest of entrepreneurs, since it could contribute to smoothing business cycles.

6. Previous work on the subject reports increasing relative demand for skilled workers, which translates into increased returns for education and rising wage dispersion. Studies of developing economies include Meza (1999) and Cragg and Epelbaum (1994) for Mexico; Robbins (1999) and Attanansio, Goldberg, and Pavcnik (2003) for Colombia; and Robbins (1994) and Pavcnik (2002) for Chile.

the period between 1979 and 1995. The authors disaggregate their data by trade orientation: export-oriented, import-competing, or nontradable sectors. They then examine labor composition and wage premiums between skilled and unskilled workers and also estimate the relative demand for skilled workers. Their results suggest the existence of skill-biased technological change. This long, deep process is related to characteristics of the production function that, in the presence of labor market imperfections, could affect income distribution and heighten the vulnerability of low-skilled workers.

4. Some Final Remarks

The papers compiled in this book are useful in several ways. They explain theoretically why existing labor regulations, though not optimal, were originally put in place. The major conclusion is that market failures are at the heart of labor regulations. Given the existence of those failures, some of the papers suggest appropriate regulations for handling them (for instance, layoff taxes and unemployment insurance), together with their optimal or efficient design in order to generate the appropriate incentives for both firms and workers. Other papers focus on how initial conditions affect the final outcome, how different outcome variables and types of workers are affected, and how each type of regulation may have a different effect on the same outcome variable when applied in a different context, time, or place.

Not all regulations are properly designed to achieve the most efficient outcome given the specific market conditions and type of failure to be addressed. Many regulations are not intended to improve efficiency, and those that are often do so ineffectively, either because the regulations are not implemented correctly or because they were designed without an adequate analysis of market interactions. For example, several design failures are identified in the specific case of unemployment protection regulation. The final outcome of any regulation thus depends on many factors, including initial conditions, as evidenced by the modeling of state-contingent effects of partial reforms.

The book also provides a wide-ranging empirical analysis of labor markets and their regulatory framework. The effects of regulations are not easily accounted for. Data problems and endogeneity are just some of the obstacles to obtaining a robust estimate of the effects.

Furthermore, the importance of enforcement versus legislation is highlighted, as is considering each regulation independently, as individual regulations have varying effects on different outcome variables, such as growth and inequality. In addition, regulations also have an impact on the composition of unemployed workers with regard to skill level, age and gender.

Given the difficulties of measuring the impact of specific regulations, several chapters of the book offer alternative methods for assessing the degree of labor market flexibility by looking directly at the adjustment processes. Two papers provide a cross-section view of the adjustment process at the macro- and microeconomic level, respectively, while another presents a time-series analysis of the recent evolution of competitiveness in European countries.

Designing better regulations, particularly unemployment insurance, requires detailed knowledge of the earnings dynamics of the respective country. In the case of Chile, the empirical work done on this issue helps determine how much workers would be willing to give up to stabilize their income. It is also confirmed that the demand for skills in Chile has grown more than proportionally in association with international trade, similarly to what has been reported in other emerging economies. This long–run trend interacts with labor market imperfections and should be taken into account when assessing and designing labor market policies and regulations.

Finally, the very relevant issues that are analyzed in the collection of studies in this book will definitely contribute to the understanding of the labor market, and will enlighten the discussion with new suggestions of ways to assess it empirically, and proposals to design better labor regulations and institutions.

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Distribution, Efficiency, and Labor Market Regulation: in Theory, in OECD Countries, and in Latin America

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In a perfect world, labor markets would not be regulated. Competition, rather than union contracts or legislation, would determine wages at a level consistent with full employment. Only technological obstacles would prevent labor reallocation toward the most productive jobs and occupations. No subsidies would be paid to unemployed workers, and employment would not be taxed.

The world we live in is not perfect, however, and the market for labor services is more heavily regulated than most other markets. The intensity and character of institutional interference with laissez-faire determination of employment and wages exhibit some variation over time, but they vary sharply across countries. Relating labor market outcomes to institutional indicators—from payroll taxes financing unemployment and social security benefits, to the minimum wage, collective bargaining provisions, and job security legislation—has made it possible to assess empirically their implications for aggregate and disaggregated employment and unemployment rates, for wage inequality, and for macroeconomic trends and cycles. The literature mostly treats labor market institutions as exogenous determinants of labor market outcomes. Less attention has been directed at explaining and interpreting the motivation behind institutional arrangements in the labor market. Efforts in the latter direction are all the more necessary when researchers study labor market institutions in a heterogeneous group of countries, as in the case of the

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contributions collected in Heckman and Pagés (2004). The chapters of that book, as well as the editors' introduction, provide interesting analyses of individual Latin American countries and comparisons of samples of Latin American countries and member countries of the Organization for Economic Cooperation and Development (OECD). The studies make it clear that comparisons across diverse countries are useful, but difficult, and they are not easily conducive to sharp insights. When the structure of the problems facing the countries (and markets and individuals within countries) is very heterogeneous in the sample considered, it is necessary in principle and hard in practice to disentangle the implications of structural characteristics from those of institutions and policies.

To understand why policy aims at distorting labor market outcomes, it is important to consider the effects of institutions on distribution and allocation in a second-best environment. Policies that would be clearly inefficient in a hypothetical representative-agent/perfect-markets situation can appeal to workers' representatives because they make it possible for their constituents to earn a larger share of aggregate welfare. The pros and cons of labor market institutions thus depend on deep structural and political features of different economies. The implementation of legislation and other collective action reflects not only the relative efficiency of market and policy mechanisms from a hypothetical representative agent's perspective, but also whether and how financial and other market imperfections prevent workers from internalizing profitability losses (see Bertola, 2004). The effectiveness of labor market regulation in shifting welfare toward workers depends on the extent to which employers may flexibly adjust on other margins, which in turn reflects an economy's degree of openness to international trade and factor flows, among other factors (see Bertola and Boeri, 2002). Moreover, the extent to which labor interests influence institutional arrangements reflects political factors, which in turn depends on a country's social and political structure.[1]

This paper offers a simple formalization of some relevant interaction channels and discusses their applicability to OECD and Latin American countries. Section 1 outlines a simple framework for labor market analysis, which offers a stylized representation of realistic sources of production and welfare inefficiencies. Section 2 discusses how institutional interference with laissez-faire in labor markets bears

1. See Saint-Paul (2000) for a wide-ranging analysis of the political economic channel, with special emphasis on redistribution across categories of workers.

on labor allocation. The formal framework used to illustrate the relevant insights is a simple explicit version of models analyzed by Bertola (2004b); an appendix reports mathematical derivations.

Section 3 discusses how detailed evidence from OECD countries can be interpreted from the resulting theoretical perspective. Section 4 offers an illustrative empirical exercise on simple available indicators of aggregate labor market institutions and outcomes in both OECD and Latin American countries. The exercise shows that basic insights from the similarly limited OECD evidence that was available in the early 1990s are applicable to Latin American countries. Section 5 concludes with an outline of new challenges and opportunities in empirical analysis of countries that regulate their labor markets heavily, but that differ in important respects from OECD members and lack comparable statistics regarding disaggregated labor market outcomes.

1. A Simple Perspective on Labor Markets

The following discussion is centered on a basic tool of labor economics, namely, the negative relation between employment and wage levels implied by a standard labor demand function. The appendix specifies a constant-elasticity functional form for that relation, with multiplicative shifts in the level of productivity and wages associated with each employment level. Specifically, the market features both good jobs, which are characterized by relatively high productivity, and bad ones, with relatively low productivity. Workers may also not be employed. For simplicity, non-employed workers' income-equivalent flow of utility is a given constant in the model discussed in the appendix and illustrated by the figures below. The qualitatively similar implications of upward-sloping labor supply are briefly discussed below.

1.1 Perfect Markets

The labor market model has only two endogenous variables: the overall level of employment and the fraction of that employment allocated to high-productivity jobs. Consider first how these are determined when efficient marginal conditions maximize total production in a static setting. As shown in figure 1 and derived formally in the appendix, the wage earned by employed workers depends on the fraction of high-productivity employment, which in equilibrium is such as to equalize wages across all employment opportunities if mobility among them is costless. The overall level of employment must be such as to equalize that single wage to the income-equivalent utility flow of individuals who are not employed. These individuals are out of the labor force in the perfect-market situation discussed here, but they may be involuntarily unemployed in some of the institutional configurations discussed below.

This employment allocation maximizes production, inclusive of equivalent non-employment income. It thus also maximizes the income of the "representative" individual who owns labor and other factors of production in the same proportions as the aggregate economy. It does not, however, maximize the income of workers. In the baseline case, in which the non-employment opportunities of workers are represented by a given constant, if the marginal worker is indifferent to employment, then all workers are, and their welfare is left unchanged by the fact that some of them are employed.

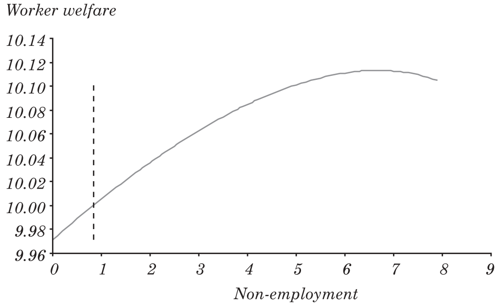
1.2 Underemployment

Since the average product of labor is larger than its marginal product, suppliers of labor benefit from restricting supply, like any monopolist. Workers as a group enjoy higher total income and utility if the level of employment is lower than that which equalizes wages and non-employment opportunities, because the labor demand relation implies that all wages will then be higher than non-employment income.

Figure 1. Labor Allocation across High- and Low-productivity Jobs[a]

a. The crossing point is identified by equation (2) in the appendix. The parameters are as follows: $A_g = 2$; $A_b = 1$; $\beta = 0.33$; and $W_u = 1$.

Figure 2. Worker Welfare as a Function of Non-employed Labor[a]



a. Equation (4) in the appendix. The vertical line identifies the competitive allocation ($W_g = W_b = W_u$). The parameters are as follows: $A_g = 2$; $A_b = 1$; $\beta = 0.33$; $W_u = 1$; and $\alpha = 0.1$.

Of course, that lower level of employment can still be allocated efficiently across jobs. In that case, as shown in figure 2, worker welfare increases with non-employment as the latter becomes larger than in laissez-faire, and it continues to increase until further decreases in employment outweigh the positive wage-bill effect of higher wage rates.

1.3 Worker Reallocation Costs

Consider next the implications of labor demand shocks. Labor productivity at each job may, with some probability, switch from good to bad or from bad to good. If mobility is costless, it should ensure that employment levels are so much higher at the good sites that the downward-sloping productivity (and wage) of labor is equalized across all jobs. If mobility is costly, however, the marginal productivity at good jobs must be higher than that at bad ones, because only a positive spread between the two can offset mobility costs. If the costs of mobility are paid by risk-averse workers rather than by a well-diversified representative agent, then the wage instability generated by costly arbitrage across employment opportunities has important implications for worker behavior and worker welfare.

The appendix formally analyzes a situation in which the labor income of a worker who is relocating from a bad to a good job falls short of the good wage by a fixed proportion that represents, for example, time spent seeking and reaching the new employment opportunity. Mobility costs are specified as a fraction of the wages for high-productivity jobs. Since those wages decrease (along downward-sloping labor demand functions) when the labor market delivers more labor to such jobs, more intense mobility reduces the cost of mobility.[2] The simple model also supposes that workers' utility is a strictly concave function of their labor income, net of mobility costs. As in Bertola (2004b), this represents the limited opportunities for workers to access financial markets in order to smooth their consumption in the face of labor income shocks.

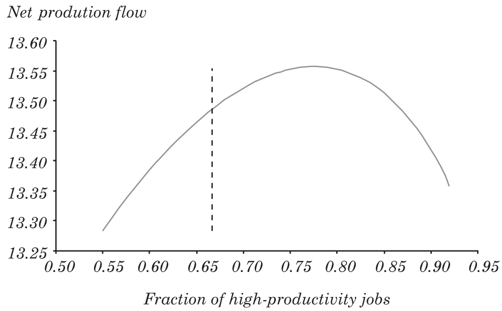
Since mobility toward good jobs is financed out of relatively low consumption flows, its costs in terms of utility are higher than they would be if labor income risk could be insured and pooled at the aggregate level. To spur mobility, future expected wage gains need to be larger when their decreasing marginal utility is smaller relative to that of the moving workers' low consumption. The labor market delivers larger wage differentials by allocating less labor to currently more productive jobs or, equivalently, by reducing the intensity of labor mobility from low- to high-productivity jobs. As fewer units of labor move from low- to high-productivity jobs, the latter become less numerous in steady state, and aggregate production is low in an economy where risk-averse workers finance mobility.

The net production flow of the model in the appendix is plotted in figure 3 as a function of the share of employment in high-productivity jobs, for given total employment. The labor allocation delivered by this simple labor market's decentralized equilibrium is identified by the vertical dashed line. It falls short of maximizing production for two reasons. On the one hand, the intensity of labor reallocation is lower than would be necessary to maximize production flows, because the specification of mobility costs makes them a decreasing function of mobility through external effects. On the other, workers' utility curvature unambiguously implies that the fraction of good employment is reduced by the need to offer larger wage differentials to moving workers. When mobility costs bear on individual workers' consumption, rather than on aggregate resources, decentralized decisions intuitively fail to maximize the latter. As in the case of the total employment level, if workers' consumption coincides with their labor income, then the

2. This may represent market-thickness effects of the type studied by Hosios (1990) and others in markets that clear through search rather than price-taking behavior. In such models, these external effects may imply that mobility is excessive in decentralized equilibrium. The simple framework outlined in the appendix has that implication if mobility costs are specified as a fraction of bad wages.

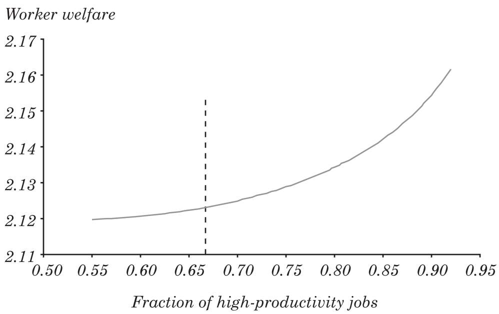
impact of labor allocation on their welfare differs from the impact of that same variable on total production (and on the welfare of individuals who earn non-labor income). The average utility yielded to workers by each unit of employment allocated on the basis of uninsured mobility decisions is plotted in figure 4. It is a monotonically increasing function of the fraction of high-productivity employment, for the simple reason that average wage both increase and become more equal as more workers are employed at firms with a strong labor demand.

Figure 3. Aggregate Production as a Function of High-productivity Employment[a]



a. Equation (10) in the appendix. Aggregate production is net of mobility costs. The parameters are as follows: $A_g = 2$; $A_b = 1$; $\beta = 0.33$; $W_u = 1$; $\alpha = 0.1$; $p = 0.2$; and $X = 0.3$. The vertical line identifies the allocation implied by condition 9 in the appendix.

Figure 4. Average Welfare of Employed Workers as a Function of High-productivity Employment[a]



a. Equation (12) in the appendix. The parameters are as follows: $A_g = 2$; $A_b = 1$; $\beta = 0.33$; $W_u = 1$; $\alpha = 0.1$; $p = 0.2$; and $X = 0.3$. The vertical line identifies the laissez-faire allocation implied by condition 9 in the appendix.

In summary, the simple model highlights related, but conceptually different interactions among the welfare of workers, the economy's productivity, and the allocation of labor across high-wage, low-wage, and non-employed states. First, as workers disregard the portion of the economy's total production that does not accrue to them in the form of wages, their welfare is not maximized by the laissez-faire employment level determined by decentralized decisions. Reducing employment has a negligible negative effect at the point at which the wage is equal to non-employment opportunities, but the higher wage enjoyed by all the workers who remain employed has a first-order beneficial effect on collective workers' welfare. This is obviously true when all employed workers are strictly indifferent between employment and non-employment in laissez-faire, and it is also true when an upward-sloping labor supply schedule lets employed workers enjoy some surplus. Second, reallocation of labor from low-productivity to high-productivity jobs improves the economy's total productivity. However, when it is driven by the mobility decisions of risk-averse workers whose consumption directly bears the cost of mobility, it is not as beneficial to their welfare and, in decentralized equilibrium, falls short of the intensity that would maximize total production net of mobility costs.

2. Labor Market Institutions and Income Distribution

The stylized model above highlights qualitative mechanisms that arguably motivate most collective interventions aimed at altering laissez-faire wage and employment outcomes. This section discusses that possibility and explores whether and how collectively decided policy measures and structural mechanisms interact to determine labor market outcomes.

2.1 Overall Employment

Workers' preferred employment level differs from that of a hypothetical representative individual interested in maximizing the economy's total production flow over and above the workers' non-employment opportunities. If workers only earn labor income, they do not mind reducing the portion of output that accrues to other factors of production. Hence, they prefer wages to be higher than those delivered by competitive interactions, and they may gladly accept the reduction in employment that delivers that outcome. If the wage is increased above that which equates aggregate supply and demand, workers who have no stake in profits or rents are collectively better off. Lower employment is a matter of indifference at the margin in laissez-faire. As wages become discretely higher than non-employment welfare, the simple sum of workers' utilities continues to increase until the lower welfare of workers who fall back on the outside opportunity more than compensates the higher wage earned by the workers who remain employed.[3]

This perspective can rationalize legal, or otherwise collectively set, minimum wages. If individual workers were allowed to bid for employment, any situation in which some workers remain unemployed who would rather be working would unravel to the competitive outcome. A legal prohibition to do so, as implied by mandated minimum wages or by the administrative extension of collective wage agreements to all workers, supports an outcome that is in the workers' collective interest because it eliminates each individual worker's incentive to bid for another's job. Of course, the unemployed would ex post prefer to be employed, especially if (as in the situation illustrated above) individual utility functions are strictly concave. However, a low-employment outcome can be agreeable to all workers if non-employed workers can partake of the higher average productivity and wages of employed workers. Some such redistribution occurs through intrafamily transfers, as in the case when the unemployed are sons and daughters of the employed. Outcomes with a larger wage bill and lower employment (and profits) may also be supported by a system of payroll taxation that funds either non-employment subsidies—such as pensions, unemployment benefits, and other welfare transfers—or public-sector employment opportunities at favorable wage-effort ratios (Algan, Cahuc, and Zylberberg, 2002). The purpose and effect of all such policies are similar to those of an explicit wage floor: while the latter prohibits workers from bidding down other workers' wages, income support eliminates the need to bid for employment.

A variety of collective policies can thus insert a wedge between labor demand labor supply, thereby reducing employment. Contractual or legislative lower bounds on wages result in open unemployment, as do tax-and-subsidy schemes in a smaller or less effectively employed

3. This point is reached sooner when labor supply is steeper. As discussed by Bertola, Blau, and Kahn (2002b), this can explain why labor market regulation tends to cause smaller employment losses for worker groups with rigid labor supply (such as prime-age males).

labor force. In different countries and periods, a range of policies are implemented in divergent ways and more or less incisively (see the discussion in the next section), and they can be interpreted in terms of distributional conflicts taking place under structural constraints, represented by the downward slope of the labor demand function. Not only the ease of union organization and the ability to enforce collectively agreed wages on all workers, but also the ability to use non-market instruments to redistribute the wage bill from high-wage employed workers to unemployed workers affect the desirability and feasibility of the low-employment outcome that approximates the maximization problem illustrated in figure 2.

2.2 Labor Reallocation and Income Stability

Institutions also interfere with the dynamics of labor reallocation in the face of labor demand shocks. Recall that the simple model outlined above does not allow workers to shelter their consumption from labor demand fluctuations by accessing financial markets. Then, smaller wage differentials and easier mobility not only improve workers' welfare through a standard consumption-smoothing channel, but also better aligns individual mobility incentives to aggregate rates of transformation, which tends to improve productive efficiency as indexed by the proportion of high-productivity employment in the model. Addressing this imperfection would thus be in the interest of both employers and workers.

In reality, workers' consumption and their mobility investments can be financed by contingent financial securities, by self-insurance through asset accumulation and decumulation, and by private labor contracts with employer-financed training or severance pay provisions. All such instruments fall short, however, of implementing the smooth consumption paths and efficiency-based reallocation and retraining decisions that would characterize a labor market with perfect financial market access. Empirically, earnings and consumption data tend to track each other quite closely at the individual level, especially at the low end of their distributions.[4]

When private financial and labor market contracts cannot shelter workers' consumption from idiosyncratic labor demand shocks and ensure that labor reallocation takes place efficiently, collective

4. See Attanasio and Davis (1996); Cutler and Katz (1992); Blundell and Preston (1998); Blundell, Pistaferri, and Preston (2002).

interventions can try to achieve both goals. In the stylized framework outlined above, improvement on the laissez-faire outcome entails taxing the payroll of high-productivity jobs, subsidizing that of low-productivity jobs, and reducing the workers' cost of moving from the latter to the former. Intuitively, taxing high wage realizations that have relatively low marginal utility and subsidizing the consumption of workers who earn low wages makes sense from an ex ante insurance point of view. Moreover, since equalization of take-home pay would remove workers' incentives to move toward high-productivity jobs, a policy package meant to mimic a first-best allocation also needs to finance mobility out of aggregate resources, subsidizing mobility as needed to ensure that additional production is valued on the risk-neutral basis appropriate for idiosyncratic shocks.

Many real-life policies can be interpreted from this perspective. Progressive taxation can smooth workers' income and consumption, thereby offsetting the implications of missing insurance markets (Varian, 1980). Centralized contracts which specify a compressed wage structure across heterogeneous regions or sectors can be rationalized by risk aversion on the part of immobile workers (Agell and Lommerud, 1992). Unemployment insurance schemes, in turn, can fund job losers' search for high-productivity jobs and ease their exit from low-productivity, low-consumption jobs while also making them more reluctant to take similar ones. Active labor market policies offer training and job-search assistance to displaced workers, as well as job subsidies to low-earners; they can, in principle, address the efficiency implications of uninsured workers' reluctance to undertake forward-looking investment decisions. Employment protection legislation also tends to have qualitatively similar effects in an equilibrium environment. By making it costly to dismiss workers, such legislation induces labor hoarding at low-productivity firms, and the wedge between wages and labor's marginal productivity at such firms is similar to that introduced by an explicit low-wage subsidy. Concern with future firing difficulties discourages hiring at the model's high-productivity firms, so employment protection tends to depress wages below marginal productivity there, just like a payroll tax would. Employment protection measures can also ease reallocation, as mobility subsidies would, to the extent that they mandate payments to workers who are laid off through no fault of their own, induce such payments in equilibrium, or encourage employers to react to labor demand shocks by reorganizing their workforce internally rather than through market mechanisms.

Using such policies to address the shortcomings of laissez-faire allocations is desirable in principle. In practice, the collective administration of tax subsidies and severance schemes is imperfect and certainly not costless, such that these mechanisms may be more or less advisable in different circumstances.

Unemployment Insurance...

The details of policy implementation are important in models that explicitly account for the informational asymmetries that prevent markets from supporting insurance contracts and prevent collective policies from achieving perfect efficiency. Much depends on equilibrium interactions and what causes unemployment in the first place. When unemployed workers are tempted to exert low search effort, a declining pattern of benefits can induce them to intensify their search initially. This efficiently reduces the duration of unemployment, even though the relatively high initial level of benefits affords the same overall insurance as that implied by a lower constant level (Shavell and Weiss, 1979). However, the different search behavior of unemployed workers influences the equilibrium distribution of wage offers: declining benefits can lead to the inefficient rejection of low wage offers by unemployed workers receiving high initial benefits, even as their stronger search effort increases their rate of matching (Albrecht and Vroman, 2003); and high initial benefits can reduce the efforts of currently employed workers to retain their jobs (Wang and Williamson, 1996). The benefits and costs of unemployment insurance systems depend on the character of information problems and market interactions, while the balance between them depends on the efficiency of policy implementation.

...and Employment Protection

Employment protection legislation requires that the termination of individual employees be motivated and that workers be given reasonable notice or financial compensation in lieu of notice. It grants workers a right to appeal against termination, sometimes stipulating reinstatement with back pay when the appeal is successful. Since such rights may not be lawfully overridden by contractual provisions, employment protection legislation interferes with individual contractual freedom as regards dynamic aspects of employment protection, just like administrative extension of collective agreements similarly interferes with individual wage-contracting freedom. Furthermore, legislation often mandates administrative procedures, involving formal negotiations with workers' organizations and local or national authorities, when large employers wish to proceed to collective dismissals or plant closures. These and other aspects of labor law aim at addressing informational problems and ascertaining whether dismissals are fair: in countries with stringent employment protection legislation, the letter of the law may allow employers to fire incompetent or lazy workers, but firms are required to prove—through regrettably costly court procedures—that termination is justified. The administrative review of collective dismissals is generally aimed at ascertaining that employers have properly considered ways to adjust internally and encouraging them to compensate workers for the social costs of dismissals—costs that financial markets may fail to internalize properly to firms' dynamic profit maximization problems. While it would be quite naive to expect government interventions to provide free of cost the same insurance that markets find it impossible to provide, it would also be naive to presume that properly designed policies cannot go some way toward resolving the relevant imperfections (Bertola, 2004).

2.3 Insurance and Production Efficiency

Institutional intervention in the labor market certainly entails costly information collection and performance monitoring, together with possible deadweight inefficiencies, but at the same time it may improve the efficiency of market allocations in an imperfect world. Not only markets, but also collective policies can present difficulties for implementing appropriate state-contingent transfers both to improve workers' welfare and increase aggregate production (and profits). The relative merits of different policies vis-à-vis markets depend on structural features. A society that can process information more efficiently at the aggregate level than in market interactions is likely to feature active policy interventions based on tax-and-subsidy packages or the direct management of labor reallocation costs. Societies with limited administrative capabilities might tend to privilege simple regulatory policies and, as in the case of employment protection legislation, mandate employers (who are presumably better informed and better insured than their employees) with avoiding or financing labor reallocation.[5]

5. Experience-rated contributions to reallocation or unemployment funds, of the type advocated by Blanchard and Tirole (2003), may combine elements of both. In reality, workers' marginal utility and employers' marginal productivity are not directly observable, and they are much more heterogeneous than in the simple two-state model used here for illustration purposes. Hence, labor courts and designers of bureaucratic schemes face difficult problems.

In general, one would expect to see relatively limited policy interference in economies where it causes small beneficial effects and large deadweight losses because, for example, a very elastic structure of economic interactions gives ample scope for individuals to escape regulation and taxation. However, the welfare effects of policy are not the same for workers who cannot shelter their consumption from income fluctuations and for individuals who can access perfect financial markets. As shown in figure 4, workers benefit from an increased allocation of labor to high-productivity jobs, but mobility per se does not improve workers' welfare, which is only a function of the overall level and stability of wages. While an efficient allocation of labor increases profits, it does not benefit uninsured workers when it is achieved by making wages more flexible (and consumption more volatile). Hence, workers and other agents differ in how they are affected by any given policy's impact on productive efficiency.

To illustrate this point, I present a policy that, in the context of the uninsured labor reallocation model of section 1, imposes a proportional payroll tax on all employment relationships at good firms, pays a proportional subsidy to all workers employed by bad firms, and uses a portion of excess good-wage tax revenues (which exceed low-wage subsidies, since bad jobs are less numerous and pay lower wages than good jobs) to pay a subsidy to all workers who are changing jobs. This policy configuration leaves employers free of turnover costs, but the implications for marginal productivities can be similar to those of employment protection legislation, as mentioned.

By correcting the misallocation introduced in laissez-faire by workers' inability to access financial markets, redistribution and mobility subsidies can improve both workers' welfare and the economy's overall production flow, provided the policy implementation is efficient (that is, more efficient than private financial market interactions). Figure 5 plots worker welfare and overall production as a function of the tax-and-subsidy rate, when 20 percent of net tax revenues is lost to administrative and deadweight costs. Although policy implementation is costly, efficiency gains can be obtained in the second-best situation we are considering. The welfare of workers whose consumption is smoothed by the tax-and-subsidy component of the policy increases relative to its laissez-faire level, normalized at unity in the figure. The policy also has a positive effect on the economy's overall production flow, since it subsidizes mobility and increases the fraction of high-productivity employment.

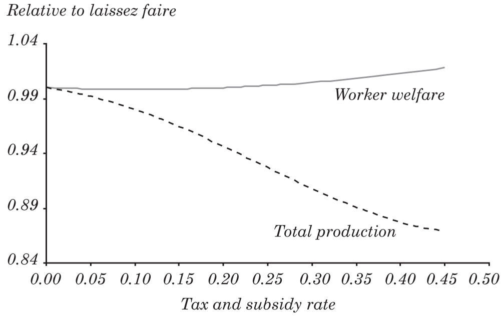
Figure 5. Total Net Production Flow and Worker Welfare as a Function of Labor Taxes and Subsidies, with Mobility Subsidies[a]

a. Large proportion of resulting revenues from taxes on high wages is used to subsidize low wages. Total net production flow is from equation (10) and worker welfare from equation (12) in the appendix; $\lambda = 80$ percent of the remaining revenue finances mobility subsidies; other parameters are as in the previous figures.

Figure 5 shows that the welfare of workers increases as the policy becomes more incisive, because they benefit from higher, more stable consumption. It continues to increase even as net production flattens out, through the insurance-based channel. Figure 6 illustrates a very inefficient policy that sets to zero the fraction of payroll tax revenue devoted to mobility subsidies. Worker welfare is still an increasing function of the tax-and-subsidy rate (which smoothes consumption), but the net production flow declines dramatically as equalization of take-home pay removes workers' incentives to move toward more productive jobs. Workers' insurance gains lead them to prefer even very wasteful arrangements of this type to a laissez-faire that burdens them (and not society) with mobility costs and, as shown in figure 7, leads to a sharp decline of "profits" (or income flows accruing to factors of production other than labor) as incentives to mobility are removed. This can decrease total production if the supply of non-labor factors of production is elastic (for example, if capital can flow into and out of the economy considered).

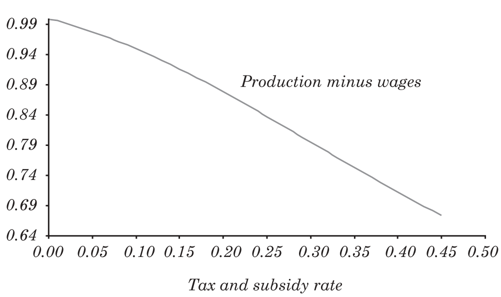
In summary, just as workers who have no stake in aggregate production over and above their wages do not appreciate high employment that depresses wages toward their outside opportunity, so workers without access to financial instruments favor policies that stabilize their labor income. They support such policies even when they have negative effects on aggregate production, because they reduce labor mobility along with the wage and consumption differentials that drive it. Labor market institutions may thus reduce overall employment when workers have limited access to ownership of non-labor factors of production and consumption-smoothing opportunities and when they have some weight in political decisionmaking processes. Under the same circumstances, wage compression may reduce workers' incentives to reallocate or retrain at the same time that dismissal restrictions with costly administrative authorization procedures prevent labor reallocation, rather than finance and encourage it.

Figure 6. Total Net Production Flow and Worker Welfare as a Function of Labor Taxes and Subsidies, without Mobility Subsidies[a]



a. Large proportion of resulting revenues from taxes on high wages is used to subsidize low wages. The parameters are the same as in figure 5, except that $\lambda = 0$ (that is, high-wage tax revenue that is not used to subsidize low wages is wasted).

Figure 7. Gross Production Flow Minus Gross Wage Flow[a]



a. The parameters are the same as in figure 5.

2.4 Why Are Institutions So Different across Countries?

The question of why institutions differ across countries is difficult to answer precisely, but it is easy to answer in broad generality. As illustrated by the theoretical framework discussed above, the benefits and costs of collective interventions in the labor market depend on structural features (such as the externality arising from the specification of mobility costs above), the ease of individual financial market access (as indexed by the degree of utility curvature in the simple model above), and the efficiency of collectively implemented policies.

The costs and benefits of the relevant policies and institutions also differ across individuals. Incomplete financial markets imply that aggregate efficiency considerations are not conveyed to all individuals by appropriate prices, and the resulting incentives to introduce distortions are heterogeneous across individuals with different productivity, non-employment opportunities, or mobility costs. Minimum wages, for example, certainly reduce the employment opportunities of low-productivity individuals at the same time that they increase the average wage and reduce the dispersion of wages among higher-productivity workers. Such distributional effects within the population of workers —rather than between workers and capitalists—interact importantly with a variety of economic and social characteristics of the relevant population. Heterogeneous labor productivity may depend on exogenous characteristics, such as age and gender, but it may also reflect policies and individual choices regarding education and training. The impact of lower employment on consumption and welfare for low-wage workers depends, first, on whether such workers belong to families who also gain from the better wages of higher-productivity workers and, second, on their non-employment opportunities.

The need to address market failures, the ability of collective schemes to do so, and the political weight of those interested in obtaining more protection against consumption shocks at the expense of productive efficiency all differ across economies and historical periods. This can, in principle, offer useful insight into the rationale of institutional settings and the desirability of reforms. Distributional concerns within an economy subject to uninsurable reallocation shocks can explain why labor market institutions often aim at providing job and wage security in the face of uninsurable labor demand shocks. Empirically, the relation between the relevant institutions and outcomes is complex. There is no clear evidence that regulated labor markets lead to inefficiency or that open economies tend to deregulate (Agell, 2002). The effects of institutions depend on the same economic features that may lead to their adoption. It is thus empirically difficult to disentangle endogenous interactions among structural and institutional features from exogenous differences across countries. Labor market institutions, employment prospects, and access to financial markets all interact endogenously, and it is not easy to ascertain the direction of causality in these interactions since institutions' economic effects also bear on their desirability. For example, the proportion of long-term unemployment is high in countries with high job security and generous unemployment benefits. Models with exogenous institutions explain small flows in and out of unemployment as the endogenous result of optimizing choices by workers (who will not search as hard when benefits are high) and employers (who will not hire and fire as much when firing costs are high). It is also possible, however, to interpret the evidence in terms of institutions' desirability, since workers will be inclined to favor stringent employment protection and generous employment benefits when job loss entails long periods of unemployment with little job-finding prospects. The development of financial markets may also be partly endogenous to demand for financial services, which can be choked by extensive regulation and redistribution.

Such positive feedback between the motivation behind policies and their effects implies that small differences in truly exogenous features of different economies can lead to widely divergent outcomes. For example, cultural and ethnic homogeneity may ease the implementation of collective policies, while an efficient legal system may improve market interactions. The structure of product markets also matters, in that real wages are determined by producers' markups over labor costs and by union markups over non-employment welfare (Nicoletti and others, 2001). While such features are not easy to measure and need not be fully exogenous to economic interactions, progress has been made in bringing such a perspective to bear on labor market institutions. Djankov and others (2003) and Bertola and Koeniger (2004) offer different perspectives on the relation between indicators of judicial efficiency, labor market institutions and outcomes, and financial market features. The following sections outline simple empirical facts that, while stopping short of answering the question above, offer suggestive indications of meaningful covariation across aspects of labor market regulation and labor market outcomes.

3. Some Evidence from Industrialized Countries

A vast literature studies unemployment and other aspects of labor market experience in light of labor market institutions, emphasizing the contrast between the United States (and other Anglo-Saxon countries), on the one hand, and European (especially Continental European) countries, on the other. The experiences of these two groups of OECD member countries have largely mirrored each other over the last few decades: in the 1960s through much of the 1970s, the unemployment rate of typical European countries was much smaller than its U.S. counterpart, but by the late 1980s a virtually uninterrupted trend increase had raised European unemployment rates above U.S. rates by a large multiple. The literature seeking explanations for this reversal of fortunes focuses primarily on labor market institutions, such as high levels of union coverage and generous social insurance benefits. Since cross-country differences in such respects were largely the same in the 1960s, 1970s, and 1980s, the literature also focuses on restrictive monetary policy in Europe and other macroeconomic shocks that have been found to explain a large portion of diverging unemployment experiences, especially when interacted with institutional features. Public employment patterns and demographic factors (such as the rapidly decreasing youth population) also play a potentially important role.[6]

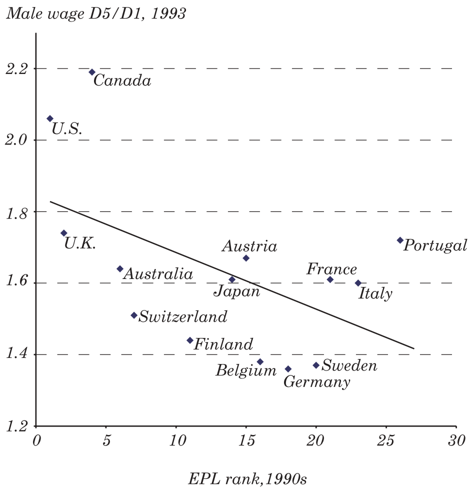
The relation between labor market institutions and the inequality and instability of labor incomes is particularly relevant from this perspective. Empirical work on such aspects can exploit the wide variation in the stringency of employment protection legislation across OECD countries. Only some employment protection features are readily measured quantitatively, such as the number of months' notice required for individual and collective dismissals. Others are more difficult to measure precisely, as in the case of the labor courts' willingness to entertain appeals by fired workers or judges' interpretation of the notion of just cause for termination. When available employment protection indicators are positively correlated with each other, however, it is possible to form qualitatively unambiguous cross-country rankings of

6. For a review of the issues and empirical results, see Bertola, Blau, and Kahn (2002a) and the references therein, especially Nickell and Layard (1999); Blanchard and Wolfers (2000); Nickell and others (2003); and Algan, Cahuc, and Zylberberg (2002). See Bertola (2001) for a review of simple economic insights and empirical indicators.

employment protection legislation and then to relate these rankings to qualitative indicators of labor market performance, in light of theoretical implications. The evidence reviewed by Bertola (1999) suggests that more stringent employment protection legislation is indeed associated with more stable aggregate employment paths. The remainder of this section briefly reviews some simple evidence on the relation between labor market institutions and labor income distribution and stability in OECD countries. The relevant empirical findings are illustrated by simple bivariate graphs. Since many aspects of labor market regulation are highly collinear across countries, these graphs offer remarkably clear insights, which can be confirmed on a multivariate basis.[7]

Figure 8 plots wage inequality against the OECD employment protection index. In line with the simple theoretical perspective outlined above, wages are compressed in the same markets where employment protection legislation is most stringent. Quantitative firing restrictions, in fact, could hardly be binding if wage fluctuations were completely unrestrained: wages would fall in response to the labor

Figure 8. Employment Protection and Wage Dispersion[a]



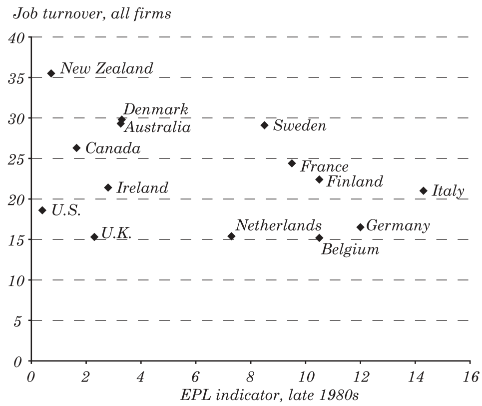
Source: Author's calculations, based on data from OECD.
a. The horizontal axis is the employment protection ranking based on 1990s indicators (the lower the value, the more stringent the protection); the vertical axis is the ratio of median wages to the tenth percentile of the wage distribution for full-time male workers.

7. See Bertola, Blau, and Kahn (2002b) and the references therein, and Bertola (2004a), which focuses particularly on the changing impact and configuration of labor market institutions in OECD countries.

demand shocks that employment protection legislation is meant to protect workers against, making stable employment profitable or inducing voluntary quits. Hence, limiting the freedom offered to employers and workers in setting wages gives force to quantity constraints. Moreover, to the extent that severance payments reduce workers' mobility costs and are larger when they are mandated by legislation than when they are left to imperfect private contracts, it is not surprising to find that more stringent employment protection legislation is associated with smaller equilibrium wage differentials. To the extent that job security provisions explicitly require, or implicitly encourage, payments from the firing firm to departing employees, more stringent employment protection legislation implies that mobility costs are at least partly borne by firms, rather than by workers, and this should be associated with smaller wage differentials in the presence of voluntary mobility across jobs. Figure 9 shows that indicators of the intensity of labor reallocation across firms are only mildly related to employment protection rigidity indicators. This negative evidence has been the subject of extensive investigation (see, for example, Bertola, 1999). The data are very noisy, which may explain the insignificant relationship between the two variables. The bivariate relationship displayed in figure 9, however, is not driven by obvious missing covariates. For example, the cyclical state of the economy has very small effects within each country, so the timing of turnover measurement cannot account for the lack of a stronger relationship between employment protection legislation and turnover. The evidence does not readily support a simple view of employment protection legislation as a rigidity factor. This may, perhaps, be taken to indicate that in terms of the simple framework above, payments to laid-off workers foster the financing of mobility by financially constrained workers.

In practice, rigid labor market configurations appear quite effective in sheltering workers from idiosyncratic labor-income fluctuations. The OECD index of the stringency of employment protection legislation is strongly associated with average tenure lengths (see figure 10). It is also positively associated with wage stability indicators (figure 11). The latter piece of evidence only refers to the few countries for which time-series stability indicators are available, but it is particularly relevant to the theoretical perspective outlined above. In heavily regulated labor markets, workers who are employed tend to remain employed, and their wages tend to remain stable over time. Stability of labor income for such workers is valuable in protecting their consumption (and that of their families) from fluctuations.

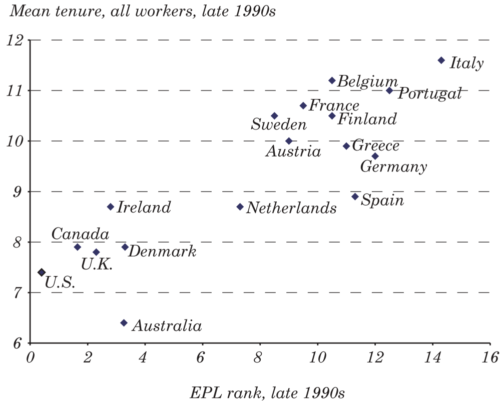
Figure 9. Employment Protection and Job Turnover[a]



Source: Author's calculations, based on data from OECD.
a. The horizontal axis is the fixed-term and regular employment protection legislation ranking; the vertical axis is job turnover, measured as absolute employment increases and declines across all firms, normalized by employment level.

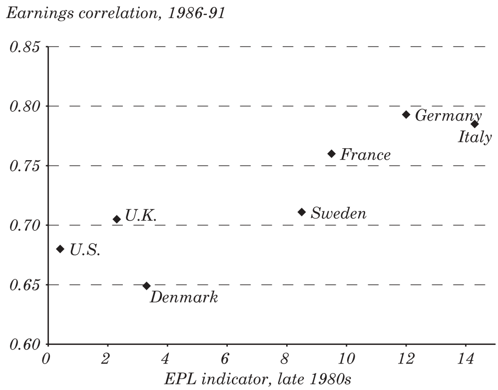
In the absence of suitable smoothing instruments, heterogeneous welfare losses from labor demand instability may rationalize frequently expressed concerns with increasing wage inequality and labor market insecurity in the United States, the United Kingdom, and other relatively unregulated labor markets. The simple evidence of figure 11 can also be interpreted in terms of the welfare effects of labor income inequality and instability at the individual level. As in Bénabou and Ok (2001), inequality can be associated with higher welfare for risk-averse individuals if mobility is intense and if the transition probabilities to higher and lower income (and consumption) levels are nonlinear, so as to give good prospects of upward mobility. This condition is almost satisfied in the United States, where individuals need not resent inequality very much. In other OECD countries, however, workers' prospects of upward mobility appear much more limited—and even when currently poor workers may look forward to higher future income, financial markets tend to prevent consumption smoothing. Data and information are scarce on households' (as opposed to firms') financial market access. Bertola and Koeniger (2004) show that less-developed consumer credit (as determined by countries' historically determined judicial efficiency) makes stable labor incomes more attractive from the perspective of welfare theory, and it is empirically associated with stringent employment protection legislation and wage compression in available cross-country data. To the extent that wage inequality is

Figure 10. Employment Protection Legislation and Job Tenure[a]



Source: Author's calculations, based on data from OECD.
a. The horizontal axis is the overall employment protection rank in the late 1990s; the vertical axis is the mean tenure across jobs existing in 1995.

Figure 11. Employment Protection Legislation and Earnings Fluctuations[a]

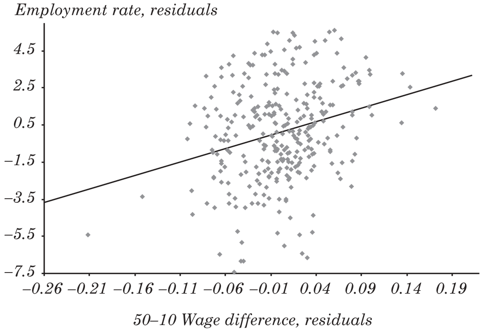


Source: Author's calculations, based on data from OECD.
a. The horizontal axis is the overall employment protection indicator in the late 1980s; the vertical axis is wage stability, measured as the correlation of earnings over five years after 1986 for full-time employees.

endogenous to labor market institutions, it is not surprising that it is very limited in countries with low earnings mobility and tight borrowing constraints.

While labor reallocation across jobs is only mildly related to the stringency of employment protection legislation in OECD countries (figure 9), flows between employment and unemployment are much smaller in economies with a high degree of employment protection (see Blanchard and Portugal, 2001). Economies that provide equal and stable incomes to employed workers often feature other institutions (like collective bargaining) that tend to increase average wages, make low-productivity workers difficult to employ, and generate a large and stagnant stock of unemployed workers. The data provide indications of meaningful trade-offs between employment rates and wage equalization. Figure 12 shows that higher wage inequality is significantly associated with higher employment rates, after controlling for country effects (which may capture institutional and structural features that change only slowly over time, if at all, within each country) and time effects (which may offer a stylized summary measure of the common technological or trade-related forces that tended to increase the differentiation and turbulence of labor demand in industrialized countries over the 1970–2000 period).[8] To the extent that overall employment is lowered by high collectively bargained wages, unemployment and other forms of non-employment are concentrated at the beginning and end of individual working careers, as well as among women.[9]

Figure 12. Residuals from Regression on Year and Country Dummies[a]



Source: Author's calculations, based on data from OECD.
a. The horizontal axis is the difference between the median wage and the tenth percentile of the overall wage distribution; the vertical axis is the employment rate of the working-age population.

8. Bertola, Blau, and Kahn (2002a) and Bertola (2004a) offer a more detailed discussion of such phenomena, as well as the role of labor market institutions in mediating the impact of structural shocks on wage and employment patterns.

9. Bertola, Blau, and Kahn (2002b) show that institutions motivated by rent extraction in labor's favor naturally tend to induce stronger wage increases and steeper employment declines, for elastically supplied labor.

4. Latin American Data and Evidence

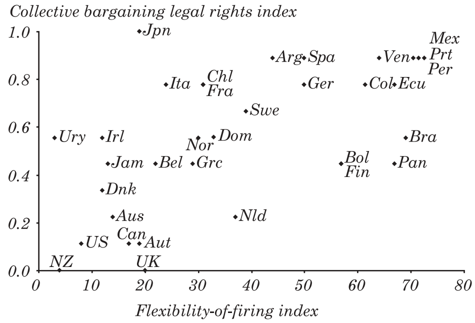
The literature is just beginning to apply the insights reviewed above to OECD countries. The cross-country set of labor market institutions and outcomes assembled by Djankov and others (2003) spans a very wide range of countries, and the authors argue that historically determined judicial and legal systems may be viewed as an exogenous source of institutional variability across countries. For Latin America, the contributions summarized and discussed by Heckman and Pagés (2004) offer mostly country-specific studies of time-series experiences, which emphasize the role and impact of labor market reforms and the use of microeconomic information. Reforms, however, are infrequent and seldom radical, and they do not appear to produce consistent patterns of effects.

Research is hampered by the relative scarcity of comparable cross-country information on the relevant static, dynamic, and distributional aspects. Unemployment and labor force participation (or employment) rates are available in the World Bank WDI database yearly only from 1980 to 2000 for most (but not all) countries. Some yearly data are missing, perhaps indicating changes of country-specific definitions. Not all countries are large enough to represent an independent observation, and not all countries in the WDI database are also present in institutional-indicator databases. The relevant information is available for a set of twenty industrialized and sixteen Latin American countries.

The available information on Latin American institutions is also less plentiful and precise than that compiled and made available for OECD countries. The Heckman and Pagés (2004) study of Latin American labor laws is based on institutional information obtained from surveys of country officials, aggregated in an index aimed at summarizing in terms of wage labor costs the impact not only of social security and other tax-and-subsidy provisions, but also of firing restrictions, on the basis of U.S. turnover rates. This indicator is somewhat sparse and cannot properly assess employment protection legislation's impact on employment dynamics. However, when consistently computed for both OECD and Latin American countries, it does indicate that the latter tend to be more regulated. A similar impression is conveyed by the purely institutional indicators compiled by the World Bank's Rapid Response Unit, which are displayed in figure 13 along with a simple index of collective-bargaining rights that may capture workers' ability to implement low-employment outcomes.[10]

These simple data deliver equally simple messages. First, institutional interference with employers' freedom to dismiss workers covaries positively with constraints on individual contractual freedom across both OECD and Latin American countries—and not only among the former. This is consistent with the notion that both may be motivated by underlying country-specific economic and political concerns with imperfect (especially from the workers' point of view) laissez-faire outcomes. Second, Latin American countries (with the exception of Uruguay and Jamaica) cluster in the high-regulation quadrant of the figure,

Figure 13. Covariation across Labor Market Institutions[a]



Source: Author's calculations, based on data from World Bank, Rapid Response website (rru.worldbank.org/DoingBusiness); and International Institute for Corporate Governance, 2003 labor database (iicg.som.yale.edu/data/datasets/labor_dataset_4_01_03.xls).

a. The horizontal axis is the World Bank's flexibility-of-firing index from 0 (most flexible") to 100 (least flexible); the vertical axis is the International Institute for Corporate Governance's index of collective-bargaining legal rights, measured as the "normalized sum of (i) labor union power [employer duty to bargain with unions, extension of collective contracts to third parties, law allows closed shops, (ii) right to unionization in the constitution, and (iii) right to collective bargaining in the constitution" (index_col_barg1 in the data set). The countries included in the figure are as follows, with the abbreviated code given in parentheses: Argentina (Arg), Australia (Aus), Austria (Aut), Belgium (Bel), Bolivia (Bol), Brazil (Bra), Canada (Can), Chile (Chl), Colombia (Col), Costa Rica (CRi), Denmark (Dnk), Dominican Republic (Dom), Ecuador (Ecu), Finland (Fin), France (Fra), Germany (Ger), Greece (Grc), Honduras (Hnd), Ireland (Irl), Italy (Ita), Jamaica (Jam), Japan (Jpn), Mexico (Mex), Netherlands (Nld), New Zealand (NZ), Nicaragua (Nic), Norway (Nor), Panama (Pan), Peru (Per), Portugal (Prt), Spain (Spa), Sweden (Swe), United Kingdom (UK), United States (US), and Uruguay (Ury), Venezuela (Ven). Other countries omitted for lack of data or economic significance in unweighted regressions.

10. The index is from the 2003 Labor Database maintained by the Yale School of Management's International Institute for Corporate Governance (available at iicg.som.yale.edu/data/datasets/labor_dataset_4_01_03.xls). Experimentation with the index of unionization rights gives rather different results; in particular, it yields no association with low-employment outcomes. This may indicate that the quality of the data is poor or that unions foster efficiency in the face of both market and government imperfections, along the lines of Checchi and Lucifora (2002).

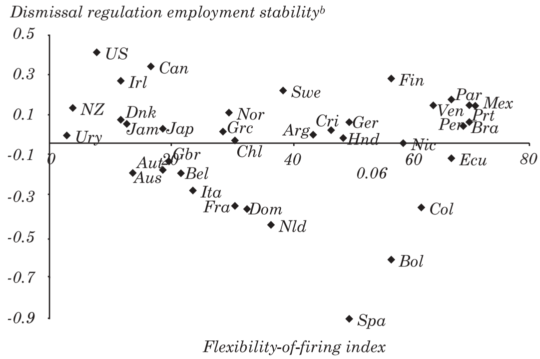
in the company of such OECD countries as Germany and Spain, while Anglo-Saxon countries are found in the opposite quadrant. Finally, the covariation between these conceptually different indicators is not perfect, and this can be exploited, in principle, to disentangle their different theoretical effects.

Empirical work on such data may provide indications of the data's quality and the fit of theoretical implications. Over and above the quality of the data, however, empirical investigation is constrained by their scarcity. Unlike the simple bivariate diagrams of the previous section, which illustrate empirical relations whose robustness could be assessed in light of multivariate results in other work, those that follow illustrate methodological issues rather than offer empirically sound results.

To see whether labor market outcomes across these countries are consistent with available institutional information, consider first the theoretical association between dismissal costs and employment stability. Figure 14 reports the cross-country association between the available index of dismissal regulation and a simple measure of employment stability (namely, the deviation of country-specific coefficients from their cross-country average in a regression of changes of employment rates on the growth rate of output per worker). This coefficient should be more positive when employment reacts strongly to changes in productivity and demand, that is, when labor hoarding is less pervasive. The regression also includes country-specific intercepts. These may, to some extent, control for the underlying productivity and employment rate trends. The results, however, can certainly be polluted by regime changes in such trends, as may occur on transition out of agriculture and on changes in demographic employment patterns across genders and age groups. Employment does react more strongly to production changes in countries with more rigid employment relationships, but the relation is far from clear-cut. A linear regression of employment-rate changes on GDP growth rates and on the interaction of GDP growth with the index of employment protection yields the expected negative sign for the latter's coefficient, but the t statistic is only 0.67. Different cyclicality of wages in different countries could be an important source of bias and noise in the estimated relationship. The evidence from OECD countries indicates that the cyclical behavior of wages is very shallow in all countries. The evidence for Latin American countries may be different, but comparable wage data are not as readily available.

The assumption that economic structures and shock intensities are similar when assessing the implications of institutions is clearly less appropriate in for Latin America than for the OECD, where all countries experienced broadly similar fiscal and monetary policies and energy cost shocks. An extensive study of OECD countries could collect observable shock indicators of the type considered by Blanchard and Wolfers (2000). In the absence of such information, however, one can infer the typical strength of labor demand and other shocks in the relevant countries from the volatility of production. That volatility is, itself, affected by movements in both labor demand labor supply, or more generally by the cyclical character of wage fluctuations (which, as mentioned, are not easily measured for OECD countries). Volatility is also affected by the stringency of constraints on employer's ability to hire and fire labor: if employment patterns are smoothed by firing restrictions, then total production should be less variable, and production per worker more variable, than would otherwise be the case in the face of similar driving processes. For a given value of the firing rules index in figure 15, and with presumably similar incentives to engage in labor hoarding behavior, Latin American countries tend to display higher volatility of production. Those countries' labor demand thus appears to be less stable than that of OECD countries—and their employment dynamics, while more pronounced than those of more industrialized and less regulated labor markets, are substantially smoothed by binding institutional restrictions.

Figure 14. Employment Protection and Employment Stability[a]

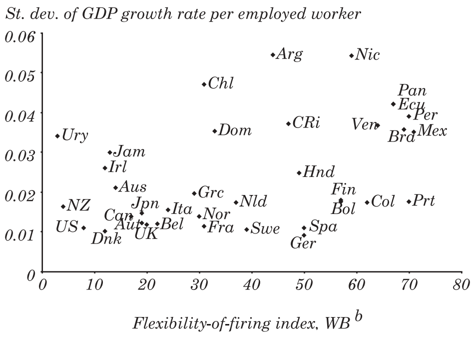


Source: Author's calculations, based on data from World Bank, Rapid Response website (rru.worldbank.org/DoingBusiness) and *World Development Indicators*.

a. The horizontal axis is the World Bank's flexibility-of-firing index from 0 (most flexible) to 100 (least flexible); the vertical axis is the deviation from the overall average of the country-specific slope coefficient in a regression of employment rate changes on real output growth rates, with country-specific intercepts (1980–2000, although the sample period stops earlier for some countries). See the notes to figure 13 for a definition of the country codes.

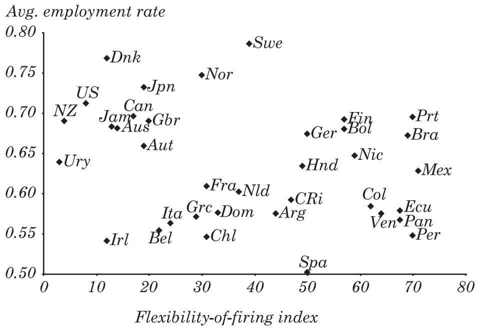
b. The horizontal line depicts the average effect of worker output growth rate on employment changes.

Figure 15. Firing Rules and Volatility of Production per Worker[a]



a. The horizontal axis is the World Bank's flexibility-of-firing index from 0 (most flexible") to 100 (least flexible); (see previous figures); the vertical axis is the standard deviation of the GDP growth rate per employed worker (computed for 1980–2000 based on data from World Bank, *World Development Indicators*). See the notes to figure 13 for a definition of the country codes.

Figure 16. Firing Rules and Average Employment Rates[a]



a. The horizontal axis is the World Bank's flexibility-of-firing index from 0 (most flexible") to 100 (least flexible); (see previous figures); the vertical axis is the employment rate (computed for 1980–2000 based on data from World Bank, *World Development Indicators*). See the notes to figure 13 for a definition of the country codes.

Turning next to the impact of regulation on employment and unemployment rates, it is instructive to first consider the bivariate association of dismissal restrictions and average employment rates in the 1980–2000 period (the shorter period of data availability). Figure 16 displays a mild negative association: employment rates are higher in countries (such as the Anglo-Saxon members of the OECD) that do not regulate dismissals extensively than in countries (such as Latin American ones) where employment protection legislation is stringent.

This association does not necessarily imply that rigid labor market relations damage employment. It is possible, in principle, to control for the structure and evolution of the labor force, for other observable development indicators, and for unobservable country-specific and time-varying institutional indicators (see Bertola, Blau, and Kahn, 2002b). In practice, however, suitable data are not available for Latin American and other non-OECD countries. Nevertheless, available indicators of firing costs and collective bargaining are imperfectly correlated, as shown in figure 13. This offers a useful opportunity to disentangle the effects of institutions such as employment protection, which are meant to protect workers against uninsurable shocks, from those of institutions such as collective bargaining rights, which are aimed at decreasing employment and increasing wages.[11]

To this end, table 1 reports the results of regressing unemployment and employment rates on the two institutional indicators displayed in figure 13, both for the whole sample and allowing Latin American coefficients to differ from OECD ones. The indicators are constant within each country, and the fit of these equations is poor.[12] In the pooled regressions reported in the first and fourth columns of the table, however, collective bargaining rights tend to increase unemployment and decrease employment rates, which is consistent with theoretical predictions. Theory has no clear prediction regarding the impact of firing restrictions on overall employment or unemployment. In regressions that control for wage-setting institutions, firing restrictions tend to be insignificantly (and negatively) associated with unemployment and insignificantly (and positively) associated with employment, which is contrary to the impression conveyed by the bivariate plot of figure 16. This simple evidence offers some support to the theoretical distinction between labor market policies affecting wage setting versus mobility. It casts doubt on the view that all labor market regulation has adverse employment and welfare effects, and it may suggest useful methodological venues for further data gathering and econometric work.

The evidence is rough, however, and far from clear-cut. The second and fifth columns report regressions with a simple Latin American dummy, which turns out to be more economically and statistically significant than the institutional indicators. When the latter are interacted

11. Of course, unemployment insurance and payroll taxation with nonemployment subsidies also have related effects, but better information is required to disentangle them.

12. Specifications with year dummies, which may capture employment developments that are common across countries within the period, yield similar estimates for the coefficients of interest.

with the Latin American dummy, the Latin American coefficients are significantly different as a group from the OECD coefficients in the case of unemployment, but the differences of individual coefficients are large, insignificant, and unbelievable. If these regressions are taken at face value, collective bargaining rights would tend to increase unemployment less in Latin America than in the OECD; firing restrictions would have no impact on OECD unemployment but a somewhat negative impact on Latin American unemployment; and the different unemployment outcomes would be accounted for by other features specific to Latin America. This evidence fails to support simplistic views of labor market institutions, but it clearly does not offer precise insights. Given the lack of appropriate controls and the poor quality of institutional information, it is impossible to tell whether similar institutions have structurally different implications in different countries.

Table 1. Regressions of Unemployment and Employment Rates on Institutional Indicators[a]

	Unemployment rate			*Employment rate*		
Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
Flexibility of firing[b]	–0.019	–0.041	0.015	–0.0004	–0.0001	0.0002
	(0.45)	(0.89)	(0.25)	(0.54)	(0.07)	(0.14)
Interaction with			–0.097			–0.0003
Latin America dummy			(1.19)			(0.25)
Collective bargaining	1.694	1.135	1.344	–0.075	–0.065	–0.056
rights[c]	(0.53)	(0.34)	(0.37)	(1.35)	(1.17)	(0.80)
Interaction with			–10.135			–0.106
Latin America dummy			(1.52)			(0.96)
Latin America dummy		2.474	13.067		–0.040	0.043
		(1.11)	(2.54)		(1.36)	(0.77)
Constant	8.543	8.794	7.137	0.694	0.690	0.679
	(6.66)	(6.63)	(5.56)	(34.39)	(32.88)	(25.17)
R^2	0.01	0.06	0.17	0.12	0.17	0.19
Test of coefficient equality across Latin America and OECD (p value)			0.052			0.618

Source: Author's calculations, based on data from World Bank, *World Development Indicators* and Rapid Response website (rru.worldbank.org/DoingBusiness); and International Institute for Corporate Governance, 2003 labor database (iicg.som.yale.edu/data/datasets/labor_dataset_4_01_03.xls).
a. The dependent variable is the unemployment rate (in percent) in regressions 1 through 3 and the employment rate (as a fraction of the total labor force) in regressions 4 through 6. Unreported results for specifications with year dummies are very similar. The sample covers 593 country*year observations, clustered by country. Robust t statistics in parentheses.
b. Flexibility of firing is indexed from 0 (most flexible) to 100 (least flexible).
c. Collective bargaining rights are indexed from 0 (weakest) to 1 (strongest).

5. The State and Future of Latin American Labor Market Research

The quality of the data—and thus of the empirical results—on labor market institutions and outcomes currently available for Latin America is perhaps comparable to that available for OECD countries fifteen years ago (see, for example, Bertola, 1990). The aggregate data broadly conform to basic theoretical implications based on the very simple indicators and regressions examined in section 4 above. Collective bargaining may well tend to decrease employment and increase unemployment, because lower employment and higher wages are attractive for workers aiming to improve their own (and not a representative individual's) welfare. The effects of more general forms of labor market regulation are less clear-cut, and the evidence does not refute the theoretical prediction that employment protection legislation stabilizes employment without decreasing its average level below that implied by the wage-setting process.

Little information is available for assessing more subtle channels of interaction, but the available evidence appears to be consistent with the insights gained by the literature's analysis of OECD data. For example, Kugler's (2004) difference-in-differences analysis of a Colombian employment protection reform finds that its effect on unemployment was small (a reduction of 0.15 percent) and of doubtful significance. And the cross-national evidence analyzed by Heckman and Pagés (2004) finds little evidence of a meaningful association between their employment protection indicator and unemployment outcomes.[13] Both of these studies report important effects of labor market regulation on the stability of employment relations and on the composition of employment across demographic groups. These disaggregate findings are consistent with the OECD evidence discussed in section 3 above, and they are readily interpreted in terms of the distributional implications emphasized above and the desirability of labor income stability and disemployment (especially of secondary labor force groups, as in Bertola, Blau, and Kahn, 2002b).

The comparable labor-income inequality and stability data that would make it possible to extend the discussion of section 3 to

13. Earlier versions of both Kugler's and Heckman and Pagés's research reported larger unemployment effects, which were reduced (respectively) by more careful computation of steady-state relations and by the inclusion of Chile in the sample.

Latin American countries are not yet available. When extrapolating theoretical and empirical work from OECD countries to the Latin American reality, it is important to account for differences between these two groups of countries in terms of within-country inequality of resources and levels of market development, on the one hand, and the time-series instability of macroeconomic and institutional conditions, on the other.

Inequality is a serious matter in Latin American and other industrializing countries. The Gini coefficients reported in the World Bank's *World Development Indicators* database range from 24.7 (Denmark, 1992) to 40.8 (United States, 1997) among OECD countries, and from 36.4 (Jamaica, 1996) to 60.7 (Brazil, 1998) among Latin American ones. The simple average of the twenty OECD Gini coefficients is 30.9; that of the twenty-three Latin American Ginis is 49.4. If, as argued above, labor market institutions often trade production efficiency and profits for the protection of workers with limited access to financial markets, then conflicts of interests between capitalists and workers may well be more relevant in Latin America than conflicts between employed and unemployed workers (and other categories), as studied by Saint-Paul (2000) in an industrialized context. Furthermore, mobility in Latin America is probably more limited by a strong class structure and limited education and training opportunities than in most of the developed countries depicted in figure 11. Thus, if labor market institutions are endogenous to political processes in which workers have some weight, then institutional interference should be more intrusive in Latin America than in OECD countries. Furthermore, to the extent that income stabilization for workers with limited or no access to international financial markets destabilizes other agents' (capital) income, capital flows in that region should be highly volatile (Bertola and Drazen, 1994).

As regards the dynamics of macroeconomic and institutional developments, the relevance of forward-looking considerations in labor market behavior and the strong reform tensions in Latin American labor markets imply that measurement issues should, in principle, be addressed by using some measure of expected (rather than current) institutional rules in constructing forward-looking indicators. An example is the simple, but potentially insightful measurement approach of Heckman and Pagés (2004), in contrast with the purely cross-sectional indicators used above for illustrative purposes. The relevant empirical issues are complex because suitable reform instruments are difficult to identify in a setting where not only outcomes, but also institutions may be viewed as endogenous variables. Still, detailed country-specific studies of the type collected and discussed by Heckman and Pagés (2004) offer intriguing indications of sensible covariation between reform features and economic circumstances, and they suggest that the starkness of Latin American developments should offer fertile grounds for interpretation of similar, if slower, developments in European countries. For example, Cassoni, Allen, and Labadie (2004) report that Uruguay experienced a dramatic regime change from a so-called command economy to unionization and from near-autarky to Mercosur integration. The large increase in the economy's openness, which presumably resulted in more elastic labor demand, had theoretical and empirical implications for labor market interactions that appear similar to and much more dramatic than those relevant to industrialized countries experiencing globalization and European integration: the 49 percent real wage decrease in the decade after 1973 in Uruguay certainly dwarfs wage-moderation trends in Europe in the 1990s.

Inequality and reforms interact importantly. If labor market institutions are a partial substitute for inefficient financial contract enforcement (Bertola and Koeniger, 2004), and if increased flexibility in the labor market makes limited access to consumption smoothing all the more painful for workers, then labor market liberalization will face heavy resistance in industrialized countries with poor financial markets. Labor and financial market reforms should therefore be packaged together, as was the case in the United Kingdom in the 1980s (Koeniger, forthcoming). As both the redistributive political appeal and the efficiency costs of labor market regulation are enhanced by Latin America's inequality and instability, much more dramatic reforms may occur there than in OECD countries. From this perspective, Latin American countries offer a rich set of reform experiences: several have pioneered the use of notional benefit accounts, which may indeed target the financial market failures emphasized above as possible rationales for observed collective interference with labor market outcomes. If the rate of return on notional benefit accounts suitably reflects that of investment opportunities available at the level of the aggregate economy but unavailable to individual workers, they can ease the liquidity-constraint problems studied by Bertola and Koeniger (2004), at least if withdrawals are allowed in the relevant contingencies. The dramatic variability of macroeconomic and institutional dynamics in most Latin American countries offers empirical opportunities to gain further insights into theoretical mechanisms. The equally dramatic heterogeneity of personal circumstances in less developed countries makes it important to take into account the intended benefits of institutional interference with the workings of labor markets when discussing their actual shortcomings.

Appendix

Let the relationship between the marginal worker's revenue product and the wage be

$$A_i L_i^{-\beta} = W_i \, , \tag{A1}$$

that is, labor demand has constant elasticity $0 < \beta < 1$ and is subject to multiplicative exogenous shocks. The level of that shock, A, and the employment, L, and wage, W, are all indexed by i. They may refer to a specific firm, sector, or region within an aggregate economy. To make the points of interest, it is sufficient to consider only two possible values of that index: $i = g$, for good employment opportunities and $i = b$ for bad ones. The wage-equivalent income of workers who are not employed in either kind of job also plays a role in what follows; let it be denoted W_u and, for simplicity, let it be the same for all individuals in a population of total size $\bar{L}$.

The number of nonemployed labor units are denoted U, while P represents the proportion of employment at high-productivity sites. This yields

$$L_g = \left(\bar{L} - U\right)P \text{ and } L_b = \left(\bar{L} - U\right)\left(1 - P\right) .$$

Equality of marginal productivity at the two sites is obtained if

$$P = A_g^{1/\beta} / \left(A_g^{1/\beta} + A_b^{1/\beta}\right) , \tag{A2}$$

which, since $A_g > A_b$, implies that the fraction of employment in jobs with higher-than-average productivity is intuitively larger than the fraction in jobs with lower-than-average productivity. The overall level of employment must equate wages to the outside opportunity, W_u. The condition,

$$A_b \left[\left(1 - P\right)\left(\bar{L} - U\right)\right]^{-\beta} = A_g \left[P\left(\bar{L} - U\right)\right]^{-\beta} = W_u \, , \tag{A3}$$

is solved by

$$U = \bar{L} - \left(A_g^{1/\beta} + A_b^{1/\beta}\right) W_u^{-1/\beta} \, .$$

If this is positive, then overall employment falls short of total available labor, $\bar{L}$. Otherwise, all individuals are employed, and their income equals or exceeds W_u.

Let the utility accruing to workers be a power $0 < \alpha < 1$ of their labor income, and consider a simple-minded measure of labor's collective welfare—namely, the sum total of all workers' utility:

$$\bar{V} = P(\bar{L} - U)\left(W_g\right)^{\alpha} + \left(1 - P\right)(\bar{L} - U)\left(W_b\right)^{\alpha} + \left(W_u\right)^{\alpha} U \ .$$

When $W_g = W_b = W_u$ as in equation (3), workers' welfare is simply given by $(W_u)^{\alpha}\bar{L}$: all workers are indifferent between working or simply enjoying their nonemployment opportunities. Lower employment increases aggregate worker welfare by increasing wages along sloped demand curves in the form of equation (1). If the proportions of employment allocated to the two types of jobs are kept fixed at the expressions in equation (2), then

$$W_g = W_b = \left(A_g^{1/\beta} + A_b^{1/\beta}\right)^{\beta}\left(\bar{L} - U\right)^{-\beta} \ ,$$

and thus

$$\bar{V} = (\bar{L} - U)^{1-\beta\alpha}\left(A_g^{1/\beta} + A_b^{1/\beta}\right)^{\beta\alpha} + \left(W_u\right)^{\alpha} U \tag{A4}$$

is increasing in U at the point where $W_g = W_b = W_u$: as long as $\beta\alpha < 1$, workers' welfare is maximized when the ratio of wages to nonemployment income equals $(1 - \alpha\beta)^{-1/\alpha}$. In the $\alpha = 1$ case of risk neutrality, this is a familiar markup in the form $(1 - \beta)^{-1}$. The markup can be shown to be smaller when $\alpha < 1$ and utility is not only increasing, but also strictly concave in labor income. Intuitively, higher unemployment increases both the mean and the variance of workers' incomes, and this is less attractive when the utility function is more concave.

Let p denote the probability of a shock that causes productivity to fall from A_g to A_b, and also of the opposite transition. While workers who are not relocating earn W_b or W_g, workers who move from a bad to a good job earn XW_g, with $X \leq 1$. Again supposing that workers' utility flows are a concave function of their labor income, consider the undiscounted expected values of utility accruing over an infinite horizon to workers holding each type of job. If V_g denotes that value from the point of view of a worker who holds a good job and has no reason to move, then

$$V_g = \left(W_g\right)^{\alpha} + \left[\left(1 - p\right)V_g + pV_b\right] , \tag{A5}$$

since the job may remain good with probability $1 - p$ but may also turn bad. Symmetrically, a worker holding a bad job and remaining there can hope that a positive labor demand shock will be realized, which occurs with probability p. Thus,

$$V_b = \left(W_b\right)^\alpha + \left[pV_g + \left(1-p\right)V_b\right]. \tag{A6}$$

Subtracting equation (6) from equation (5) yields

$$2p\left(V_g - V_b\right) = \left(W_g\right)^\alpha - \left(W_b\right)^\alpha . \tag{A7}$$

A worker holding a bad job, however, can move to a good job. The value of doing so is

$$\left(XW_g\right)^\alpha + \left[\left(1-p\right)V_g + pV_b\right].$$

If workers are individually indifferent to mobility, it must be the case that this is the same V_b in equation (6), which implies that

$$\left(W_b\right)^\alpha - \left(XW_g\right)^\alpha = \left[\left(1-2p\right)\left(V_g - V_b\right)\right].$$

This equation and equation (7) yield a relationship between the wages paid by the two types of jobs, which differ in terms of both their current productivity and possible future developments:

$$\left(W_g\right)^\alpha = \left(W_b\right)^\alpha + \left[\left(W_g\right)^\alpha - \left(XW_g\right)^\alpha\right]2p . \tag{A8}$$

Solving for the wage configuration that makes mobility optimal for workers, labor allocation must be such as to yield

$$\frac{W_g}{W_b} = \left[1 - \left(1 - X^\alpha\right)2p\right]^{-1/\alpha} .$$

In equilibrium, the proportional wage premium paid by good jobs is positive, since $0 < 1 - (1 - X^\alpha)2p < 1$ for $\{2p, X, \alpha\} \in (0, 1)$, and decreasing in X, since mobility becomes costless (and equalizes wages) as X approaches unity. This premium is also increasing in p, the probability of a change in labor demand conditions: as labor demand becomes less stable, larger wage premiums are needed to compensate mobility investments by workers who have an option to stay put and hope for an improvement of their current job's conditions. As p approaches one-half, the proportional wage premium approaches $1/X$.

In the $p = 0.5$ case, in which the future outlook is the same at all jobs, wages must be such as to compensate workers for their mobility costs within the same period when mobility occurs.

Conversely, if $p < 0.5$, in which case the labor income of workers who change jobs is lower than W_b, mobility is still optimal, but only because workers can look forward to persistently high wages once they reach a good job. The relevance of forward-looking considerations implies that their utility's degree of concavity plays an important role in determining equilibrium wage differentials and the intensity of mobility. In fact, the equilibrium ratio of good to bad wages is decreasing in α as long as $p < 0.5$: it is lowest at the upper boundary of the $\alpha \in (0, 1]$ range, and it increases as α declines toward zero, making the utility function increasingly inelastic.

With P again representing the proportion of employment at good firms, in equilibrium it must be the case that

$$\frac{A_g P^{-\beta}}{A_b (1-P)^{-\beta}} = \frac{W_g}{W_b} \ ,$$

so

$$P = \frac{\left[\left(W_g / W_b\right) A_b\right]^{-1/\beta}}{\left(A_g\right)^{-1/\beta} + \left[\left(W_g / W_b\right) A_b\right]^{-1/\beta}} \ . \tag{A9}$$

As a fraction, p, of high-productivity jobs experiences a negative shock, and the same fraction of bad jobs experiences a positive one, $p(2P - 1)$ units of labor are relocated each period. Each moving worker earns only a fraction, X, of the good wage to which he or she is moving. Consequently, in the aggregate, mobility dissipates $(1 - X)W_g = (1 - X)A_g P^{-\beta}$ units of output in flow terms. Total production net of such costs is

$$\left[\frac{A_g}{1-\beta}(P)^{1-\beta} + \frac{A_b}{1-\beta}(1-P)^{1-\beta} - p(2P-1)(1-X)A_g(P)^{-\beta}\right]\left(\bar{L} - U\right)^{1-\beta} \tag{A10}$$

The labor allocation delivered by this simple labor market's decentralized equilibrium does not maximize this expression. That is, it does not satisfy the first-order condition,

$$A_g\left(P\right)^{-\beta}-A_b\left(1-P\right)^{-\beta}-\left(1-X\right)2pA_g\left(P\right)^{-\beta}$$
$$+p\left(1-X\right)A_g\beta P^{-\beta}\left(2-\frac{1}{P}\right)=0\ . \tag{A11}$$

The first three terms of condition (A11) add up to zero in equilibrium only if mobility decisions are taken on a risk-neutral basis ($\alpha = 1$); otherwise, they exceed zero, since mobility by uninsured workers calls for larger wage differentials. Even under risk neutrality, the last positive term would call for even higher P and even lower wage differentials, because the proportional specification implies that external effects reduce the cost of mobility as mobility becomes more intense and reduces good wages.

The average utility generated for workers by each unit of employment is

$$\bar{V}=A_g^{\alpha}\left(P\right)^{-\beta\alpha}-A_b^{\alpha}\left(1-P\right)^{-\beta\alpha}\ . \tag{A12}$$

To simplify the notation, let the tax-and-subsidy rate be the same fraction, τ of pre-tax wages (and labor marginal productivities). The after-tax mobility condition for uninsured workers is then

$$\left[\left(1-\tau\right)W_g\right]^{\alpha}=\left[\left(1+\sigma\right)W_b\right]^{\alpha}+$$
$$2p\left\{\left[\left(1-\tau\right)W_g\right]^{\alpha}-\left[X\left(1-\tau\right)W_g\left(1+\nu\right)\right]^{\alpha}\right\}, \tag{A13}$$

where the proportional mobility subsidy, ν, must obey the policy's budget constraint: $\nu X(1-\tau)W_g$ is paid to each of the $p(2P-1)$ units of labor reallocated in a typical period, and the revenue of payroll taxes net of low-wage subsidies is $[\tau A_g(P)^{1-\beta}-\tau A_b(1-P)^{1-\beta}]$, but only a fraction, λ, of this is available to finance mobility. The shortfall of λ below unity represents administration costs and the deadweight costs of distorted economic behavior. Hence,

$$\nu=\frac{\tau A_g\left(P\right)^{1-\beta}-\tau A_b\left(1-P\right)^{1-\beta}}{p\left(2P-1\right)X\left(1-\tau\right)W_g}\lambda$$

is the mobility subsidy rate. If equation (13) is rewritten inserting this expression and expressions for gross wages in terms of P, the equilibrium fraction of high-productivity jobs, then the latter is determined in equilibrium by the condition,

$$(1-2p)\left[(1-\tau)A_g P^{-\beta}\right]^{\alpha} = \left[(1+\tau)A_b(1-P)^{-\beta}\right]^{\alpha}$$

$$-2p\left\{\frac{p(2P-1)X(1-\tau)A_g P^{-\beta} + \tau\lambda\left[A_g(P)^{1-\beta} - A_b(1-P)^{1-\beta}\right]}{p(2P-1)}\right\}^{\alpha},$$

which can be solved numerically.

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On the Effects of Targeted Employment Policies

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In many European countries, employment policies are often framed as measures aimed at favoring particularly disadvantaged groups in the labor market. These groups are defined in terms of individual characteristics such as age, gender, skill, or unemployment duration, which are thought to be negatively correlated with worker productivity. Differentiated or dual labor market policies with different provisions for high-wage and low-wage jobs are pervasive across the labor regulations of many countries. Thus, the higher incidence of unemployment among low-skilled workers is often used to advocate targeted employment subsidies for this group (see, for instance, Drèze and Malinvaud, 1994), which have been introduced in many countries. Payroll tax rebates for low-skilled workers and the introduction of "atypical" employment contracts (such as part-time, fixed-term, or seasonal contracts), which have low firing costs and are restricted to certain groups of workers, are also very common.[1]

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1. For a description of the nature of temporary contracts in some European Union countries, see Booth, Dolado, and Frank (2002).

Labor Markets and Institutions, edited by Jorge E. Restrepo and Andrea Tokman R., Santiago, Chile.

While using two-tier schemes in reforming the labor market may enhance political support (see Saint-Paul, 2000), the full consequences of allowing for different employment regulations affecting different workers in markets with heterogeneous agents are not yet fully understood. Indeed, most papers analyzing the effects of employment policies do not take into account the targeted nature of some of those policies.

Our intuition on the importance of the targeted nature of some employment policies stems from our previous work on labor markets with two-sided heterogeneity in jobs and workers. In labor markets with heterogeneous workers searching for jobs that may be occupied by workers of different skills, the turnover rate of one type of worker affects the overall labor market tightness and, hence, the flows in and out of unemployment for all workers.[2] We explore these ideas in an earlier paper (Dolado, Jansen, and Jimeno, 2003), in which we consider a matching model with two-sided heterogeneity (complex and simple jobs and low- and high-skilled workers). In the model, high-skilled workers can be mismatched (that is, they can work in simple jobs), and, if they are, they engage in on-the-job search for better jobs.[3] We show that, when job requirements and workers' skills are heterogeneous, differentiated firing costs may reduce equilibrium unemployment. The intuition for this result is that the mismatch of high-skilled workers implies a negative externality of on-the-job seekers on low-skilled workers when both types of workers are equally productive in simple jobs: because mismatched workers have a higher quit rate than low-skilled workers, they make those jobs more unstable and thus reduce the profits of firms opening them, leading to a lower job creation rate.[4] Alternatively, firms opening simple vacancies will be less eager to do so in the presence of on-the-job search, and this will worsen the employment prospects of less productive workers who can only work in simple jobs. To the extent that larger firing costs for workers in complex jobs reduces workers' turnover in this type of job, the size of the negative externality will also decrease,

2. Labor market tightness is defined as the ratio between the number of vacancies and the number of workers searching for a job (namely, the unemployed) in models without on-the-job search and as the ratio between the number of vacancies and the number of unemployed workers plus mismatched workers in models with on-the-job search.

3. Low-skilled workers cannot perform complex jobs.

4. There is also a positive externality on the supply of unskilled vacancies, however, since more workers are looking for those jobs. The negative externality dominates.

leading to a larger supply of simple jobs. Indeed, it is possible to construct examples in which skill-biased technological change in the presence of targeted employment protection legislation may end up reducing skill mismatch by so much that the unemployment rate of both types of workers falls.[5]

In this paper, we extend some of the previous ideas to focus the discussion on the effects of targeted employment policies. Again, we envisage a labor market with worker heterogeneity and imperfect substitution of workers with different skills to perform a single type of job. Workers can be hired under different types of contract subject to different firing costs or employment subsidies. In this setup, as discussed above, workers' flows and job reallocation depend on the overall labor market tightness, so that any measure, whether targeted or not, affects the employment outcomes of all workers. Our aim here is to identify the main factors determining the effects of targeted employment policies in order to establish what is needed for partial reforms to become successful, in terms of both cutting unemployment and increasing welfare. We argue that the employment effects of targeted policies may hinge crucially on the initial state of the labor market. Thus, for instance, a reduction in firing costs targeted to low-productivity workers in sclerotic labor markets, where it is easy to find unemployed workers to fill vacancies, raises labor market tightness by so much that it may end up reducing the low-skilled unemployment rate without affecting the unemployment rate of high-skilled workers, which may marginally rise or fall. At the same time, the wages of both workers increase, because their outside option value rises in a tighter labor market. With an increase in wages and a reduction in unemployment, the welfare of both types of workers rises, yielding support for such a policy.

In tight labor markets, where it is difficult to find candidates to fill vacancies, the abovementioned targeted employment policy often increases the unemployment rate of low-productivity workers, because the increase in job destruction is larger than the increase in

5. Other papers also use search equilibrium models with worker or job heterogeneity to analyze the effects of some employment policies. For instance, Acemoglu (2001) shows that unemployment benefits and minimum wages may raise welfare in a model with good and bad jobs in segmented markets. Albrecht and Vroman (2002) analyze a labor market in which low and high-educated workers can be hired for unskilled jobs while only high-educated workers can perform skilled jobs, albeit without allowing for on-the-job-search as in Dolado, Jansen, and Jimeno (2003).

job creation, while there is little effect on the unemployment rate of high-productivity workers. Indeed, the welfare of low-productivity workers typically falls, while the welfare of high-productivity workers increases. The political support for such a partial reform will depend on the composition of the labor force. If, as is reasonable, there is a large proportion of unskilled workers in the economy, this kind of targeted test will not be politically feasible.

The rest of the paper is organized as follows. Section 1 presents some examples of targeted employment policies, paying particular attention to employment protection legislation. The focus on employment protection legislation is justified because in many countries, not just in Western Europe, provisions such as notice periods, procedures for dismissals, and severance payments often vary across occupations. Moreover, recent reforms to employment protection legislation have implemented targeted reductions of firing costs, in many occasions through the introduction of atypical contracts, yet only for workers with the worse employment outcomes. Section 2 is devoted to the empirical literature on targeted employment policies, including both cross-country and case studies pertaining to specific country experiences, so as to identify the effects of these reforms. Section 3 contains a summary of the theoretical implications derived from a search equilibrium model addressing the effects of partial reforms on various labor-market outcomes, such as unemployment, wages, job reallocation, and welfare. This theoretical discussion highlights two key issues: the main channels through which targeted employment policies may affect the labor market outcomes of all population groups, not only those of the targeted group; and the main factors determining the sign of the overall effect on unemployment and its distribution among population groups. Lastly, section 4 concludes.

1. Examples of Targeted Employment Policies

The targeted nature of employment policies is quite evident in the provision of employment incentives and in recent reforms of employment protection legislation. With regard to the provision of employment incentives, most countries provide financial incentives—such as top-up wages, social security contribution rebates, and tax credits—for hiring workers with some specific characteristics leading to worse employment outcomes. As for employment protection legislation, recent reforms have mainly liberalized atypical employment contracts to allow hiring certain population groups under less strict dismissal regulations.[6]

Employment incentives typically target young workers with low skills. For instance, in Germany employee allowances are aimed at improving the employment prospects of unemployment assistance recipients, with a strong focus on young persons.[7] In Italy, vocational integration schemes aimed at young people in depressed areas offer training allowances that are equally financed by the employer and the provincial employment office. These schemes may be converted into a training-cum-work-contract entailing reduced social security contributions. The conversion of these contracts and also of apprenticeship contracts into permanent ones is encouraged by extending the reduction of social security contributions for an additional twelve months.[8] Tax relief is also offered for job creation, but only when the new employee is under 25 years of age and has been unemployed for a period of at least twenty-four months. France targets young people who are encountering special difficulties in finding work, with a focus on those aged sixteen to twenty-one without a diploma and those under twenty-five who do not have a vocational qualification or who abandoned their studies. These groups are eligible to be hired under so-called orientation contracts, qualification contracts, and employment-initiative contracts, in which the wages may be set below the minimum wage, the employer is exempted from social security contributions, and the state may pay a lump-sum recruitment subsidy. Finally, Spain offers reductions of social security contributions, which can reach 70 percent, for employers that grant permanent contracts to unemployed persons under thirty or over forty-five years of age or to unemployed women registered as jobseekers for at least one year and recruited in occupations and activities in which female workers are underrepresented.

Employment protection legislation, in turn, varies significantly not only across countries, but also within countries, based on firm and worker characteristics such as firm size, the existence of collective agreements, job tenure, and workers' skill and educational levels.[9]

6. Detailed information about these measures in European Union countries can be found in European Commission (2003).

7. The allowance is about twelve euros, on top of the wage received from the employer, for every day the employee worked at least six hours.

8. This rebate is 50 percent in the north of Italy and 100 percent in the south.

9. OECD (1999) presents a detailed and comprehensive description of employment protection legislation in several countries and its variation by worker skills, tenure, the existence of collective agreements, and firm size. For a justification and the implications of variable enforcement of employment protection legislation by firm size, see Boeri and Jimeno (2005).

With regard to skill level, for example, there are two sources of variation in the enforcement of employment protection legislation. First, procedural requirements for dismissals, notice periods, and severance pay provisions for unfair dismissals are usually stricter for white-collar workers. Second, high-skilled workers are not always entitled to be hired under atypical employment contracts with less strict employment protection provisions.

Countries where employment protection provisions are less strict for blue-collar workers include Austria, Belgium, Denmark, France, Greece, and Italy. In all of these but France, the required notice period is shorter for blue-collar workers than for white-collar workers, and in Denmark and Greece, blue-collar workers are entitled to lower severance pay. Severance pay for unjustified dismissal is also lower for blue-collar workers in Belgium and Greece.[10]

Spain provides a paradigmatic case study of partial reforms introducing atypical employment contracts. Faced with an unemployment rate above 20 percent in 1984, the Spanish government tried to implement a significant change in employment protection legislation by liberalizing temporary contracts in two main respects: their use was extended to include hiring employees to regular positions (not just to seasonal or probationary positions); and they entailed much lower severance payments than the regular permanent contracts. As a result of this two-tier reform (permanent contracts retained their previous indemnities for fair and unfair dismissals), the proportion of temporary employees in total salaried employment surged in the second half of the 1980s and stayed above 30 percent (35 percent in 1995) after 1990. A series of countervailing labor market reforms aimed at reducing the reform's incidence were introduced in 1994, 1997, 2001, and 2002, aimed at providing a less stringent employment protection legislation for permanent contracts and considerable restrictions on the use of fixed-term contracts.[11] A new permanent contract for new hires was introduced in 1997. The main novelty was that under this contract, mandatory firing costs for unfair dismissals were lower than those pertaining to the old permanent contracts (thirty-three days of wages per year of seniority with a maximum of twenty-four months' wages versus forty-five days of wages and forty-two months' wages,

10. Institutional details of employment protection legislation in these countries are in OECD (1999). The information in the text refers to the late 1990s.

11. See Dolado, García-Serrano, and Jimeno (2002) for a detailed description of those reforms.

respectively). Only certain population groups were eligible to be hired under the new contract, however—namely, young workers (aged eighteen to twenty-nine), long-term unemployed registered at the public employment office for at least twelve months, unemployed above forty-five years of age, disabled people, and workers whose contracts were transformed from temporary into permanent ones. In the 2001 reform, the government managed to extend the use of the new contracts to young workers between sixteen and thirty years of age, long-term unemployed registered for at least six months, and unemployed women of any age working in sectors where they were underrepresented.

Spain is not the only country that has liberalized atypical employment contracts or reduced firing costs contingent on specific workers characteristics. In 1984, Italy introduced employment promotion contracts (*contratti di formazione e lavoro*) aimed at promoting the hiring and firm-based training of young workers (aged fifteen to twenty-nine). Likewise, fixed-term contracts were first introduced in France in 1979, but their scope was very much reduced by the socialist government in 1982. As of a reform in 1990, fixed-term contracts can be used only for seasonal activities, the replacement of an employee on leave, temporary increases in activity, *and* the facilitation of employment for targeted groups, from the young to the long-term unemployed (Blanchard and Landier, 2002).

Partial labor market reforms have taken place in many Latin American countries, sometimes aimed at decreasing firing costs (Colombia and Peru at the end of the 1980s) an others at raising them (Brazil, Venezuela, Chile, the Dominican Republic, Nicaragua, and Panama).[12] However, the only country that significantly liberalized the use of atypical contracts targeted on certain demographic groups was Argentina, where a reform in 1991 introduced fixed-term contracts and training contracts for young workers. A new reform in 1995 introduced special contracts to promote the employment of population groups facing disadvantages in that respect.

2. Empirical Evidence on the Effects of Targeted Employment Policies

The empirical literature on the labor market effects of labor regulation contains two distinctive streams. First, cross-country studies

12. See IDB (2003).

use quantitative or qualitative indicators representing the effect of those institutions to explain international differences in labor market outcomes, such as employment and unemployment rates.[13] Within this stream of the literature, recent studies look at the interactions between institutions and shocks, as well as the different impacts of institutions on the labor market outcomes of different population groups, such as young people and women.[14] However, this literature often considers targeted employment policies or partial labor market reforms only in the construction of the overall institutional indexes, and not separately as individual institutional features. As stressed in this paper, this treatment can be very restrictive since, for instance, a general reduction of firing costs does not have the same labor market effects as a commensurate reduction in the severance payments of a certain group of workers.

Nevertheless, a few studies estimate the labor market impact of some targeted employment policies, such as temporary contracts, separately from aggregate indexes of employment protection legislation. Among them, Jimeno and Rodriguez-Palenzuela (2002) find that less strict regulation of fixed-term employment contracts tends to reduce youth unemployment rates without any impact on the unemployment rate of prime-age males. In a similar vein, Nunziata and Staffolani (2001) use an unbalanced panel of nine member countries of the Organization for Economic Cooperation and Development (OECD) during the second half of the 1980s and first half of the 1990s to estimate the effects of employment protection legislation. They allow for three types of regulations: employment protection legislation on firing permanent employees, regulations regarding fixed-term employees, and regulations on temporary work agencies. They find that less stringent fixed-term contract regulations have a significant positive impact on temporary and total employment during upturns, with no significant effect on total permanent employment. In the case of young workers (fifteen to twenty-four years of age), less stringent fixed-term contract regulations increase both temporary and permanent employment. With regard to temporary work agencies, they find that less stringent regulations have an incremental effect on

13. See Nickell and Layard (1999).

14. On interactions, see Blanchard and Wolfers (2000). On the differential impact of labor market institutions across population groups, see Bertola, Blau, and Kahn (2002), Jimeno and Rodriguez-Palenzuela (2002), and Neumark and Wascher (2003). On the impact of employment protection legislation on employment adjustment, see Caballero, Engel, and Micco (2003).

temporary employment and total employment, but only during downturns. In the case of young workers, however, less stringent regulations for temporary work agencies raise temporary employment while reducing permanent employment.

The second stream of the literature looks at specific country episodes to evaluate the effects of labor market reforms through the analysis of labor market outcomes before and after the reform (that is, using a differences-in-differences format). Studies of this type include Kugler, Jimeno, and Hernanz (2003) on the 1997 Spanish reform, Blanchard and Landier (2002) on France, and Hopenhayn (2001) on the Argentine reform. In Spain, Kugler, Jimeno, and Hernanz (2003) find that the reduction of firing costs (and payroll taxes) for young and older workers and the long-term unemployed had a positive effect on the employment rate of young workers with hardly any effect on their dismissal rate, whereas it increased both dismissals and hiring among older workers. Blanchard and Landier (2002), in turn, look at transitions between temporary and permanent employment in France. They find an increase in worker turnover since 1983, especially among younger cohorts, for whom the probability of holding a fixed-term job increased, the probability of holding a permanent job decreased, and the probability of becoming unemployed showed no clear trend. For Argentina, Hopenhayn (2001) also finds that the introduction of fixed-term contracts had a very strong impact on labor turnover, inducing an increase in hiring accompanied by some substitution of permanent jobs with temporary jobs.

3. How Targeted Employment Policies Work: Some Theoretical Predictions

As discussed in the introduction, a comprehensive analysis of targeted employment policies ought to start with a consideration of the existence of worker heterogeneity and imperfect substitution among workers of different skills to perform a single type of job.[15] Worker heterogeneity is needed to justify differential treatment by targeted

15. This section is based on a companion paper (Dolado, Jansen, and Jimeno, 2004) in which we present a full-fledged search model to investigate the effects of dual employment protection legislation—namely, the reduction of firing costs for less productive workers—in the spirit of recent partial reforms of employment protection legislation in many countries.

measures. Imperfect substitution is the channel through which measures affecting only one type of worker impinge on the labor market outcomes of nontargeted groups.

In principle, the definition of worker heterogeneity must cover two relevant dimensions. First, different types of workers have different productivity levels. For instance, the distribution of productivity levels of young and low-skilled workers are bound to lie to the left of the distribution of productivity levels of older and high-skilled workers.[16] Second, the arrival rate of shocks to productivity may also vary with workers' characteristics, such as age, educational levels, and skills. Unskilled workers could be more prone to large and negative productivity shocks than skilled workers. In standard search equilibrium models of the labor market, the population group with the lowest productivity and the highest arrival rate of productivity shocks would experience the highest unemployment rate, so that, in principle, there is a rationale for introducing employment policies that target the unskilled group.

Given the distributions of productivity and their evolution over time, firms would find it optimal to hire workers of each type when their productivity is above a certain threshold, which depends on workers' reservation wage, the shapes of the productivity distributions, the arrival rate of productivity shocks, overall labor market tightness, and the skill composition of the unemployed. Similarly, firms would destroy jobs whose productivity is below a certain threshold, which depends on job termination costs, in addition to the same determinants of the hiring threshold. Targeted measures typically aim at lowering the hiring productivity threshold, the firing productivity threshold, or both, for workers with the worse productivity distribution. Thus, employment subsidies for less productive workers or a reduction in their firing costs typically makes it more attractive to hire this type of worker. Hence, the first direct effect of targeted employment policies is on the targeted group's flows in and out of unemployment.

Yet, the change in one particular group's unemployment flows affects overall labor market tightness. Under imperfect substitution of workers, hiring and firing decisions for the nontargeted group depend on overall labor market tightness. A second effect of targeted

16. In formal terms, the distribution for low-skilled workers is stochastically dominated by that of high-skilled workers.

employment policies, therefore, is to change the hiring and firing decisions that affect the nontargeted group. Labor market tightness increases because firms open more job vacancies when their wage costs drop as a result of the reduced firing costs of the low-skilled workers. Wages also rise as a result of workers having a better outside opportunity. The size of this effect, which is different for the two groups of workers under consideration (targeted or not), depends on how wages are determined and on the sensitivity of profits from opening job vacancies to the overall labor market tightness. In any case, overall labor market tightness rises while firing and hiring productivity thresholds change for all workers. Typically, the employment rate of the targeted group would improve relative to that of the nontargeted group. In the literature on active labor market policies, the fact that the relative hiring of the two types of workers changes is called the substitution effect of targeted measures, which is widely discussed in policy analysis.

However, the theoretical analysis of employment policies pays less attention to the identification of the main parameters determining the size and sign of these two effects, the direct effect on the targeted group and the spillovers on the nontargeted group. Search and matching models of the labor market, which are currently the standard toolkit for the analysis of employment polices, typically ignore worker heterogeneity and nonsegmented markets for workers of different skills, features which we argue are key for a comprehensive account of the effects of targeted employment policies. As sketched above, this set up includes many parameter values that determine these effects, and models of this class generally cannot be solved analytically. One has to resort to simulated results obtained from appropriately calibrated models.

To more precisely illustrate the nature of the interactions involved in the analysis of targeted employment policies, we now examine the reduction of firing costs for low-skilled workers when severance payments before the reform are identical for the two types of workers.[17] We first assume that the labor market initially is very sclerotic, so unemployment rates for all workers are high (but much higher for low-skilled workers) and overall labor market tightness is low, namely, it is easy to find candidates for unfilled new vacancies. Lowering the firing cost would increase job reallocation and, therefore, tightness. This leads to a reduction of the unemployment rate of

17. For equations and simulations of this kind of model, see Dolado, Jansen, and Jimeno (2004).

the targeted group, as the direct effect on job creation should dominate the effects of increasing tightness on the expected profits from opening job vacancies. The flow out of unemployment for nontargeted workers is low, as more workers of the targeted group are hired, and the unemployment rate of nontargeted group is thus likely to rise initially. However, given that firms find it easy to locate candidates for their vacancies, job creation will spill over to the nontargeted sector, so the steady-state unemployment rate of high-productivity workers hardly changes or may even fall if the matching process between workers and vacancies is sufficiently efficient. Further, the higher labor market tightness causes an increase in wages for both workers because of their higher reservation wages. With an increase in wages and a reduction in unemployment, the welfare of both types of workers rises, yielding overall support to such a policy.[18]

The process works differently in tight labor markets, where it is difficult to find candidates to fill vacancies. The targeted employment policy described above often increases the unemployment rate of low-productivity workers because the increase in turnover dominates the increase in job creation, with little effect on the unemployment rate of high-productivity workers. More precisely, the reduction in firing costs leads to an increase in layoffs of low-skilled workers, but the increase in efficiency translates into fewer additional jobs because the matching rates are relatively insensitive to changes in labor market tightness when the labor market is already tight. Indeed, the welfare of low-productivity workers typically falls, while the welfare of high-productivity workers increases. The political support for such a partial reform depends on the number of winners and losers, which is determined by the composition of the labor force. Summing up, the intuition behind this differential effect arises from the sensitivity of job creation to the increase in firms' profits from filling jobs with low-productivity workers. Jobs are filled relatively fast in a sclerotic labor market. Any change in the expected profits from hiring a low-productivity worker will therefore translate into a strong increase in job creation and in the number of matches, leading to lower unemployment. By contrast, vacancies

18. As shown by Ljungqvist (2002), the employment effects of firing costs depend crucially on how wages are determined. When firing costs are assumed to reduce the firm's threat point in the initial match, firing costs tend to increase equilibrium unemployment, whereas they tend to increase employment when the worker's relative share of match surplus is assumed to stay constant when severance pay is varied. Mortensen and Pissarides (1999) propose alternative specifications of the bargaining process in which the workers extract rents from firing costs in continuing matches but not in the first match, as in the bonding scheme.

remain unfilled for a long time in a tight labor market. Changes in the profits of filled jobs thus have a smaller effect on job creation than on job destruction, leading to a rise in unemployment. The theoretical predictions of our model also point out that the case of a higher elasticity of the matching rate of workers with respect to labor market tightness—which is an indicator of the speed at which workers find jobs—combined with a higher incidence of productivity shocks on low-skilled workers relative to high-skilled workers strengthens our previous conclusions regarding the differential effects of targeted policies in sclerotic and tight labor markets.

The fact that the effects of targeted employment policies may vary across the two types of workers depending on the initial state of the labor market is obviously very relevant for their analysis from a political-economy perspective. In particular, the targeted group will not always gain from a partial reform. The political feasibility of this type of partial reforms depends crucially on the initial state of the labor market (sclerotic or tight), the composition of the labor force, the relative incidence of productivity shocks across workers, and the efficiency of the matching process in a frictional labor market.

4. Concluding Remarks

One relevant feature of employment policies and labor market reforms is that they are very often targeted at specific demographic groups, normally those with difficulties in finding jobs (for example, youth, women, and the long-term unemployed). Some empirical studies that estimate the effects of this type of policy conclude that the impact on the labor market outcomes for different population groups can vary widely and do not always go in the same direction.

In this paper, we have argued that more analysis of these policies using search equilibrium models of the labor market with worker heterogeneity is needed. We have conjectured that the effects of targeted employment policies may depend on the state of the labor market (the degree of tightness). An interesting outcome of our analysis is that support for partial reforms is likely to be greater in sclerotic labor markets than in tight ones, since the welfare of all workers increases in the former. This issue is relevant to the debate in the literature on the optimal timing of reforms (see, for example, Saint-Paul, 1996). It is often argued that reductions in firing costs should be undertaken during expansions rather than recessions, but Saint-Paul (1996) presents compelling evidence that the opposite happens in practice. To the extent that a sclerotic labor market corresponds to bad times and a tight labor market to good times, the above discussion provides a rationale for that seemingly suboptimal practice in frictional labor markets where it is costly to match workers and vacancies.

It is illustrative to compare the effects of a partial reduction of firing costs for a targeted group, as considered above, to a comprehensive reform that delivers a commensurate reduction of firing costs for all workers (skilled and nonskilled). In the latter case, the direct effect on hiring and firing is smaller for low-skilled workers (since the reduction of their firing costs is smaller) and larger for high-skilled workers. Since, initially, the unemployment rate of low-skilled workers is bound to be higher than the unemployment rate of high-skilled workers, the comprehensive reform is likely to have a smaller impact on flows in and out of unemployment than would a targeted reform. The direct effect on overall labor market tightness is thus lower under the comprehensive reform, and the overall reduction in equilibrium unemployment is also likely to be lower. In any case, our discussion provides clear reasons to believe, first, that the effects of a targeted reform on overall equilibrium unemployment and the incidence of unemployment across population groups are different from those of a commensurate reform and, second, that in both cases, these effects may change depending on the initial state of the labor market. This calls for more theoretical and empirical analysis to achieve a better assessment of the balance between these two types of labor market reforms.

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Who Benefits from Labor Market Regulations? Chile 1960-1998

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The economic literature has devoted considerable attention to studying the impact of labor market regulations on labor market outcomes. However, the issue of whether some sub-groups of workers bear the brunt or enjoy the benefits of such regulations has been much less studied.[1] One notable exception has been the burgeoning literature studying the effect of statutory minimum wages on youth employment. Although this subject remains controversial, many studies have found negative effects of minimum wages on teenagers and young workers.[2] Less attention has been paid to the issue of whether

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1. One reference in this literature is the paper by Bertola, Blau and Kahn (2002) on the effect of unions' involvement in wage setting on the relative employment of youth, women and older individuals.

2. Among the most recent studies, Williams and Mills (2001), Partridge and Partridge (1998), Bazen and Skourias (1997), and Currie and Fallick (1996) find a negative relation between minimum wages and youth employment, while Katz and Krueger (1992), Card, Katz, and Krueger (1994), and Card and Krueger (2000) find no evidence of such an effect.

minimum wages particularly affect women versus men or unskilled versus skilled workers. One exception is the study by Lang and Kahn (1998) for the United States, which finds that a rise in the minimum wage shifts the composition of employment in the eating and drinking sector from adults to teenagers and students. Neumark, Schweitzer, and Wascher (2000) also examine the effect of minimum wages across different individuals by focusing on differential impacts of workers at different points in the wage distribution. They find that although wages of low-wage workers increase, hours worked and employment levels decline, reducing earnings for these workers.

Similarly, very little attention has been paid to the effect that job security provisions may have on particular subgroups of the labor force. Two recent exceptions are the Organization for Economic Cooperation and Development (OECD) (1999) and Bertola, Blau, and Kahn (2002). The OECD (1999) reports negative, but not statistically significant, effects of job security provisions on youth and prime-age females. Bertola, Blau, and Kahn (2002) find evidence that job security provisions increase the employment rates of male prime-age workers relative to the employment rates of male older workers. They also find evidence that job security provisions are associated with higher employment rates for prime-age women relative to women aged fifteen to twenty-four. Instead, they do not find statistically significant effects on youth relative to prime-age employment rates for male workers, or in the distribution of employment across women and men.

In this chapter, we take advantage of the unusual variance in labor market policies in Chile to examine how minimum wages and job security provisions affect different types of workers. We look at the effects of regulations on the distribution of employment by age, and also, by skill, which to our knowledge has not been examined before. To this effect, we use a sample of repeated household surveys spanning the period 1960-1998 and several measures of labor market regulations across time. We make use of cross-section and time series methods to estimate the effect that these policies have on the distribution of employment and on particular subgroups' employment rates. We are able to control for time effects that affect all workers in a similar manner, as well as demographic groups-specific effects of business cycles and labor market institutions. In addition, to assess whether our estimates are reflecting the effect of regulations instead of the effect of some unobservable correlates, we also estimate the effect of labor policy on sectors not covered by regulations. We find large and statistically significant effects on the covered sectors and no effects, or effects going in the opposite direction, on the uncovered sectors.

Our results indicate that labor market regulations are far from neutral. We find that job security provisions and minimum wages reduce the employment rates of youth and the unskilled at the benefit of older and skilled workers. We also find opposite effects of these policies on women's and men's employment shares and rates. Job security provisions tend to benefit men at the expense of women, while the reverse seems to be true for an increase in the minimum wage.

We then explore some explanations for these regularities and, while we cannot fully discriminate among all of them, we are at least able to reject some hypotheses. There is little evidence that these differential effects are driven by differences in labor supply elasticities or wage adjustments across subgroups. Instead, our findings suggest that job security regulations produce unequal shifts in labor demand across groups of workers. Regarding minimum wages, our results tend to fit the predictions of the competitive model for age and skill but not gender. Contrary to our results, the competitive model predicts higher effects of minimum wages for women because they tend to earn lower wages than men.

The rest of the paper is organized as follows. Section 1 reviews the arguments that predict non-neutral effects of regulations. Section 2 describes the evolution of job security and minimum wage regulations in Chile. Section 3 describes the data used in our empirical section. Section 4 describes the methodology implemented to estimate the effects of regulations on the distribution of employment. Section 5 describes our results for both the distribution of employment and the overall effect on employment rates. Finally, section 6 concludes.

1. Why Regulations May Affect Some Workers Differently

There are a number of reasons to suspect that labor market regulations alter the distribution of employment across subgroups. In the next two subsections, we review the theoretical arguments that predict differential effects of job security provisions and minimum wages across workers of different age, skill level, and gender.

1.1 Job Security

Job security provisions are introduced to discourage firms from adjusting their labor force in the face of adverse economic conditions. However, job security provisions also alter hiring decisions. In good times, firms hire fewer workers because they take into account that these workers may have to be laid off in the future, which is costly. Therefore, the overall impact of job security provisions on employment rates is ambiguous because it depends on whether the negative effect on layoffs is offset by the reduction in hiring rates.[3]

Job security provisions will have differential effects across subgroups of workers if changes in legislation bring changes in hiring and layoff rates that have a larger impact on some subpopulations than on others. Lazear (1990) conjectured that an increase in job security might act as a barrier, preventing the entry of young workers into the labor market. This is because job security reduces job creation, and entry rates are especially high among youth. This argument, however, does not consider that the effect of lower job creation rates can be offset by lower job destruction rates—which also tend to be large among youth. Pagés and Montenegro (1999) suggest an argument whereby job security provisions may actually *increase* young workers' layoff rates. Their argument is related to the regularity that, across countries, job security is positively related with a worker's tenure. Mandatory severance payments that increase with tenure change the cost of dismissing workers with short tenures relative to workers with more seniority at the firm. In this context, it is expected that job security concentrates layoffs among youth because, other things being equal, young workers tend to have lower average tenures than older workers. If severance pay increases substantially with tenure, and this effect is important, job security simultaneously reduces entry and increases layoffs among youth, resulting in a lower employment share and lower employment rates for this group of workers. Instead, the share of older workers in employment tends to increase due to their relatively lower layoff rates.

Similar reasoning can be used to predict the effect of job security provisions across gender. To the extent that women experience higher

3. For example see Bertola (1990), Bentolila and Bertola (1990), Bertola (1991), Bentolila and Saint-Paul (1994), Hopenhayn and Rogerson (1993), and Risager and Sorensen (1997), for a theoretical discussion of the effects of job security on employment rates.

rotation and, therefore, have lower average tenure than males at every age, high job security will tend to concentrate layoffs among women. This effect will tend to reduce their employment share relative to men. However, higher turnover rates also imply that stringent job security may be less of an issue when hiring female workers because employers expect them to quit prior to attaining high job security.[4] In this case, employers might be more willing to hire women relative to men, but also more likely to lay them off should bad times arise. The overall effect on female versus male employment rates is undetermined and remains an empirical issue.

It is tempting to extend the former argument to unskilled and skilled workers. If unskilled workers have higher rotation and lower tenures than skilled workers, the same reasoning applies. However, while higher female turnover rates may be motivated by life-cycle decisions exogenous to the employer, such exogeneity is more difficult to claim when explaining the higher rotation of unskilled workers.

The insider-outsider literature provides further arguments for why job security may have a differential effect on the employment rates of different sub-populations.[5] According to this literature, more stringent job security reduces the elasticity of wages to changes in the unemployment rate. When employed workers know their jobs are insured against demand fluctuations, they may be less willing to accept the wage adjustments necessary to reduce unemployment rates. This situation may help to create two kinds of workers: insiders, who hold their jobs and have high wages; and outsiders, who either are unemployed or hold temporary, part-time or fixed-terms jobs without job security.[6] If women, the young, and the unskilled are more likely to be outsiders, then job security (through this wage effect) will bias employment against these groups.

Finally, differences in labor supply elasticity may contribute to differential effects across subpopulations, even if job security brings a uniform change in labor demand across groups. Let us assume that an increase in job security reduces labor demand. If women, the

4. See Pagés and Montenegro (1999) for a more formal development of this argument in the context of a partial equilibrium model.

5. See, for instance, Lindbeck and Snower (1988).

6. The insider-outsider argument requires a strong union fixing wages for new entrants. Otherwise, firms could always pay very low wages at the beginning of the employment relationship to compensate for higher wages in the future. See Bertola (1990) for an analytical study of this issue.

young, and the unskilled have higher labor supply elasticity than the average worker, higher job security would bring a higher decline in employment for these workers than for other groups with a lower elasticity of labor supply.[7]

In summary, the arguments put forth in this section suggest that youth, and possibly women and the unskilled, bear the brunt of job security regulations.

1.2 Minimum Wages

The effect of minimum wages on employment remains a controversial topic. In the competitive model, workers are paid their marginal product, and any artificial increase in the price of labor above the marginal product prices the worker out of the labor market. Conversely, models that allow for employers' wage-setting power predict wages lower than the marginal product, and, thus, an increase in minimum wages can increase wages without reducing employment rates.[8]

In the Lang and Kahn (1998) model of bilateral search, the effects of minimum wages also differ from the expected effects in the competitive model. In their model, minimum wages affect the quality of the pool of applicants to jobs. Higher minimum wages allow firms to get better applicants for jobs, while reducing the employment prospects of less-productive workers.

On average, youth, women, and the unskilled tend to have lower wages than older, male or skilled workers. Therefore, because minimum wages are more likely to be binding among these workers, the competitive model predicts larger unemployment effects for the first group. In the imperfect competition model, however, the effects are less clear-cut. In principle, the magnitude and sign of the minimum wage effect will depend on how far wages are from their respective marginal products in each subpopulation. If that gap is larger in some groups than in others, an increase in minimum wages may have "competitive" effects on some groups and "non-competitive" effects on others. Given this ambiguity, the sign and magnitude of the effects become an empirical question.

7. See Hamermesh (1993).

8. There are many situations that give rise to imperfect competition in the labor market, such as incomplete information, or imperfectly mobile workers.

2. Labor Market Regulations in Chile

Chile has experienced a very wide range in labor market policies, providing a privileged case scenario for analyzing the impact of regulations on labor market outcomes. We distinguish between job security provisions and statutory minimum wages.[9]

2.1 Job Security Provisions

Among the most interesting aspects of the Chilean experience is that, in the 39 years covered by our sample, Chile has gone from a situation of dismissal at will to a rigid labor market by OECD standards (Heckman and Pagés, 2000). Since their inception in 1966, job security provisions have favored full-time indefinite employment over part-time, fixed-term, or temporary contractual relationships. To this end, in case of a firm-initiated separation, labor codes regulate the following: (1) compulsory advance notice periods; (2) the causes for which a dismissal is considered justified or unjustified; and (3) severance pay related to the tenure of a worker and the cause of dismissal. While the minimum period of advance notice has always been kept constant and equal to one month, the formula for computing severance pay and the causes for just or unjust dismissal have varied widely over the years. This is the variance that we exploit in our empirical work.

Table 1 summarizes the changes in legislation that took place in the 1960-1998 period. From 1960 to mid-1966, firms had to provide a one-month advance notice (or pay the equivalent of one month of salary), but, otherwise, "employment at will" was the norm. In 1966, the congress approved a new law under which firms had to pay compensation equal to one month's wage per year of work to all workers dismissed without just cause. The economic needs of the firm were considered a just cause in the law, and, therefore, a worker dismissed for this reason would not qualify for severance pay. In practice, however, workers would appeal to courts, and judges tended to consider these dismissals unjustified. (Romaguera, Echeverría, and González, 1995). In that event, the employer could choose between paying the mandatory compensation—plus wages foregone during trial—or reinstate the worker in his or her old post. This reform substantially increased the difficulty and the cost of labor force adjustments.

9. See Edwards and Cox-Edwards (1991, 2000) for an excellent summary of labor market reforms in Chile.

Table 1. Employment Protection Provisions in Chile: 1960 – 1998

Period	Prior notice period	Economic reasons just cause for dismissal on the law?/ in the courts?	Compensation for dismissal in case of just cause.	Compensation for dismissal in case of unjust cause.	To whom do changes apply?
1960-1966	1 month	Dismissals at will	Dismissals at will	Dismissals at will	Dismissals at will
1966-1973 Firms could not dismiss workers without a just cause.	1 month	Economic reasons were just cause in the law. In practice labor courts considered most dismissals unjustified.	The law does not mandate any compensation in this case.	One month's pay per year of work at the firm plus forgone wages during trial. Trials could last at most 6 months. There is no maximum in the amount to be awarded.	All workers.
1973-1978	1 month	Labor courts were much more pro-firm. Workers' claims were weaker.	Same as previous period.	Same as previous period.	All workers.
1978-1980 (June 15, 1978): Decree 2,200	1 month	Economic needs were considered just cause.	Zero.	1 month's wage per year of work, without maximum limit.	Only workers hired after June 1978.
1981-1984 (August 14, 1981): Law 18,018	1 month	Economic needs were considered just cause.	Zero.	1 month's wage per year of work with a maximum of 150 days.	Only workers hired after August 1981.
1984-1990 (Dec, 1984): Law 18,372	1 month	Economic needs were no longer considered just cause for dismissal.	Zero.	1 month's wage per year of work with a maximum of 150 days.	All workers.
1990-1998 (Nov. 1990): Firms need to justify dismissals.	1 month	Firms have to justify dismissals, but economic needs are considered just cause for dismissal.	Economic reasons: 1 month's wage per year of work with a maximum of 11 months' pay.	1.2-1.5 month's wage per year of work.	All workers hired after August 1981.

After 1973, a violent change in political regime brought about a *de facto* liberalization. Although job security provisions were not modified in the law, in practice, it was more likely that judges ruled against workers, effectively reducing dismissal costs. In 1978 and 1981, successive modifications reduced the cost of dismissal under the law. In 1981, the maximum amount to be awarded to a worker dismissed without just cause was reduced to the equivalent of five months' pay. This reform substantially reduced the cost of dismissal, particularly for workers with long tenures, although it only applied to newly hired workers.

After 1984, the tide shifted and job security provisions became progressively stricter. In December of that year, the law was modified to exclude economic needs of the firm as a justified cause of dismissal. However, the maximum amount payable to a worker was kept at five months of pay. In 1990, after the return of democracy, a new labor reform further increased the cost of dismissal. This law considered dismissals motivated by the economic needs of the firm justified, but employers were still liable to pay compensation equal to one month's pay per year of work, with a maximum amount of eleven months of pay. It was the responsibility of the firm to prove just cause. If such causality could not be demonstrated, there was a 20 percent surcharge in the amount of compensation.

We summarize this variance in law and court practice by means of a job security measure derived in Pagés and Montenegro (1999).[10] This measure is computed as follows:

$$JS_t = \sum_{i=1}^{T} \beta^i \delta^{i-1}(1-\delta)(b_{t+i} + a_t SP_{t+i}^{jc} + (1-a_t)SP_{t+i}^{uc}) \ ,$$

where δ is the probability of remaining in a job, β is the discount factor, T is the maximum tenure that a worker can attain in a firm, b_{t+i} is the advance notice to a worker that has been i years with a firm, a_t is the probability that the economic difficulties of the firm are considered a justified cause of dismissal, SP_{t+i}^{jc} is the mandated severance pay in that event to a worker that has been i years at the firm, and finally, SP_{t+i}^{uc} denotes the payment to be awarded to a worker with tenure i in case of unjustified dismissal.

10. See the mentioned paper and Heckman and Pagés (2000) for a complete description of the methodology, its application across time and countries, and the relative advantages and costs of using this measure versus other measures of job security.

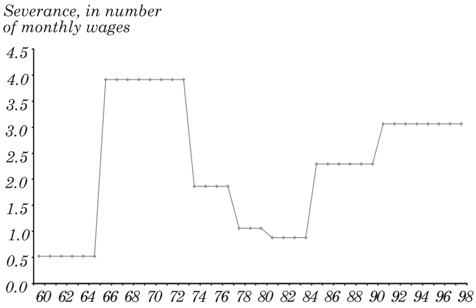
This measure computes the expected cost, at the time a worker is hired, of dismissing this worker in the future. This cost is measured in terms of monthly wages. The advantage of this measure in respect to other measures that compute the cost conditional on having achieved a certain tenure is that our job security measure captures the whole profile of severance pay at each level of tenure. The assumption is that firms evaluate future dismissal costs based on current law. Higher values of this variable indicate periods of relatively high job security, whereas lower values characterize periods in which dismissals were less costly.

Based on the legal information summarized in table 1 and assumptions regarding β, δ, a, and T, we obtain a measure of job security. We take β to be a constant value such that the average real interest is equal to 8.4 percent, which corresponds to the average real interest rate in Chile during the 1960-98 period. The probability of remaining in a job is computed based on the assumption that without job security, turnover rates in Chile would be comparable to those observed in the U.S.[11] Davis and Haltiwanger (1992) report an average annual turnover rate of 12 percent. The probability that a dismissal originated by the economic needs of the firm will be considered just depends on whether it is stipulated in the law and on the disposition of the judges to rule in favor of workers in case of a lawsuit. For the period 1966-84, although economic needs of the firm were considered just cause in the law, we assume a to be larger than zero and determined by the position taken by labor courts. Finally, we assume $T = 25$ (see table 2 for a complete description of the parameters used in the computation of the job security measure).

The evolution of this variable over time is depicted in figure 1. After some years of relatively low employment protection, job security increases eightfold after the introduction of compulsory severance pay in the law. Expected dismissal costs decline markedly in 1973 and then successively in 1978 and 1981. Subsequently, employment protection increases again, but without reaching the levels attained during the late 1960s.

11. Although turnover rates can be measured, this measure is itself affected by labor law. Given this endogeneity, we choose instead to use the U.S. turnover rate, because it is well established that dismissal costs in the U.S. are very small.

Figure 1. Job Security



Source: Pagés and Montenegro (1999).

Table 2. Parameters Used to Compute *Index*

	β	δ	b	a	SP^{jc}	SP^{uc}
1960-1965	0.92	0.88	1	1.0	0	0
1966-1973	0.92	0.88	1	0.2	0	(1)
1974-1977	0.92	0.88	1	0.5	0	(2)
1978-1980	0.92	0.88	1	0.8	0	(2)
1981-1984	0.92	0.88	1	0.8	0	(3)
1985-1990	0.92	0.88	1	0.0	0	(3)
1991-1998	0.92	0.88	1	0.9	(4)	(5)

Notes: To compute β we use the fact that the average real interest from 1960-1998 was 8.4 percent. To compute δ we assume that the average Chilean turnover rate *without* employment protection would be similar to the U.S. rate. According to Davis and Haltiwanger (1995), average turnover rates average 12 percent a year in the United States. (1) Corresponds to one month's pay per year of work augmented by three months to capture the average payments in foregone wages during trial. (2) One month's pay per year of work without upper limit. (3) One month's pay per year of work with an upper limit of five months' pay. (4) One month' s pay per year of work with an upper limit of eleven months' pay. (5) 1.2 months of pay per year of work with eleven months upper limit. We assume the maximum tenure a worker can attain at a firm is twenty-five years.

2.2 Minimum Wages

Columns 2 and 3 in table 3 present the hourly real minimum wage in 1998 pesos; these indices were constructed using Chile's Central Bank Bulletins.[12] It is interesting to note that since 1989 there has been a lower minimum wage for workers eighteen years old or younger. This wage has been fixed at a level between 15 and 20

12. Per hour minimum wages are constructed as monthly minimum wages divided by 4.2*40 hours.

percent below the adult wage. Figure 2 summarizes the evolution of the minimum wage in relation to the average wage for teen and adult workers. The figure shows that, relative to each group's average, minimum wages of teen are much higher than for adult workers. It also shows that the level of teen minimum wages has been quite volatile relative to the average wage.

Figure 2. Minimum to Average Real Wages

Source: Authors' calculations (see data section).

Between 1960 and 1998, adult real minimum wages increased by 186 percent and teen minimum wages by 104 percent. However, because average ages rose more than the increase in the minimum wages, the latter lost ground in relation to the average wage. Despite this long-term secular trend, Chile experienced a wide range of fluctuations in minimum wages, both in its rate of growth (in real terms) and in its level in relation to the average wage. During the 1960s, the real value of minimum wages was held constant, but since real wages increased, the ratio of the minimum to the average real wage declined. In the early 1970s, minimum wages increased substantially, surpassing the growth rate of average wages. In consequence, the ratio of the minimum to the average real wage increased sharply in that period. From 1975 to 1980, minimum wages lost ground relative to the average wage. After the return to democracy in 1990, real minimum wages increased steadily, but they continued declining relative to the average wage. The decline was particularly sharp for the teen group, whose minimum to average real wage rate fell from 1.80 in 1975 to 0.50 in 1998. It is interesting to note that while there are several studies in the Chilean case that suggest that the minimum wage is binding, others such as Bravo and Vial (1997) suggest that it is not.[13]

3. Data

The household surveys used in this study were obtained from the University of Chile's economics department. The economics department's survey monitors the employment-unemployment status in the metropolitan area of Santiago, Chile, four times a year. Unfortunately, only the surveys taken in June of each year contain information about wages and other employment status variables. Therefore, these are the surveys used in this study. The format of the survey and the definition of the variables have been kept constant since 1957, when the survey started, and so the information contained in them is comparable across years.[14] During the period from 1960 to 1998, the surveys interviewed between 10,000 and 16,000 people and around 3,700 and 5,400 active labor force participants each year. During this period, the metropolitan area of Santiago represented about one third of Chile's total population and a higher proportion of gross domestic product.[15] The data set is formed by stacked cross-sectional data sets, which means that individuals are not followed over time. The only restriction applied to our sample is that the people included in the estimates must be at least fifteen years old and no older than sixty-five.

We merge labor policy and macro variables taken at the annual frequency with our individual-level annual data. We include the job security index and the minimum wage data described in section 2. We also include a measure of wage bargaining to control for changes in union activity that can be correlated to our variables and to employment. While perhaps the best measure of the influence of unions on wage determination is union coverage, that is, the share of workers whose wages are affected by collective bargaining, a time series

13. See, for instance, Castañeda (1983), Paredes and Riveros (1989), Chacra (1990), Bravo and Vial (1997), Montenegro (2002), and Cowan, et al. (2003). An excellent review of the impact of minimum wages in the case of the United States can be found in Kosters (1996). A more recent survey on the international evidence of minimum wages can be found in Dowrick and Quiggin (2003).

14. In this study we use data from 1960 on, because the previous years (1957-59) do not have reliable data.

15. According to the 1992 census, the metropolitan area accounted for 39 percent of the total population.

Table 3. Basic Statistics of the Sample

	Job security index	Minimum wage		Bargaining index		Average wage							GDP deviation from trend (%)	Employment rate (%)	Wage employment rate (%)	Self-employment rate (%)
		Under 18 y/o	Over 18 y/o	Original	Smoothed	By sex: Male	By sex: Female	By skill level: Low	By skill level: High	By age group: 15-24	By age group: 25-49	By age group: 50-65				
Year	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
1960	0.52	119	119	3.33	3.33	302	152	157	475	133	283	306	−0.86	52.5	39.8	12.7
1961	0.52	114	114	3.33	3.33	370	179	171	554	164	331	435	−1.41	52.2	41.1	11.1
1962	0.52	126	126	3.33	3.33	373	203	181	615	162	361	418	−1.37	53.2	41.2	11.9
1963	0.52	109	109	3.33	3.33	376	206	n.a.	311	219	342	395	0.20	53.0	41.4	11.5
1964	0.52	107	107	3.33	3.33	268	160	n.a.	230	133	272	296	−2.15	52.9	42.3	10.6
1965	0.52	114	114	3.33	3.33	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	−5.23	54.4	43.3	11.2
1966	3.91	118	118	3.33	3.33	380	211	187	591	179	376	434	1.50	53.0	42.2	10.8
1967	3.91	116	116	3.33	3.35	427	268	222	648	217	420	539	1.50	54.0	43.2	10.8
1968	3.91	111	111	3.33	3.40	466	278	224	699	251	450	502	1.79	53.2	41.9	11.4
1969	3.91	107	107	3.33	3.46	475	279	231	709	218	470	560	2.79	52.4	41.2	11.0
1970	3.91	133	133	3.67	3.54	549	351	256	804	248	536	693	2.97	52.3	41.4	10.9
1971	3.91	183	183	3.67	3.58	689	437	302	957	307	660	779	9.67	53.7	42.1	11.5
1972	3.91	195	195	3.67	3.53	712	457	342	929	359	698	729	7.28	52.7	41.3	11.4
1973	3.91	108	108	3.67	3.41	525	332	279	671	280	512	553	0.37	51.4	39.6	11.8
1974	1.86	204	204	3.00	3.26	435	310	275	561	255	436	496	0.12	49.0	37.1	11.8
1975	1.86	245	245	3.00	3.12	376	277	225	483	214	376	420	−14.58	45.0	34.7	10.4
1976	1.86	259	259	3.00	3.01	486	352	249	635	280	474	542	−12.67	45.8	34.5	11.2
1977	1.86	269	269	3.00	2.88	692	512	320	953	357	696	786	−5.01	48.3	38.1	10.1
1978	1.06	346	346	3.00	2.62	868	517	360	1,090	400	799	1,072	0.87	48.0	37.1	10.9
1979	1.06	345	345	2.67	2.28	913	640	432	1,150	496	904	1,009	6.66	47.8.	36.8	10.9
1980	1.06	354	354	1.33	1.90	890	611	424	1,120	476	881	932	11.83	47.4	36.6	10.7
1981	0.88	334	334	1.33	1.53	1,057	799	510	1,338	590	1,099	1,016	15.64	50.9	39.3	11.6
1982	0.88	365	365	1.33	1.25	1,235	852	508	1,499	618	1,206	1,295	−1.15	41.8	33.0	8.8
1983	0.88	276	276	1.00	1.13	842	622	345	1,056	416	872	721	−6.79	43.5	34.4	9.1
1984	0.88	243	243	1.00	1.06	843	573	355	1,028	371	845	780	−4.19	46.1	35.8	10.3

Table 3. (continued)

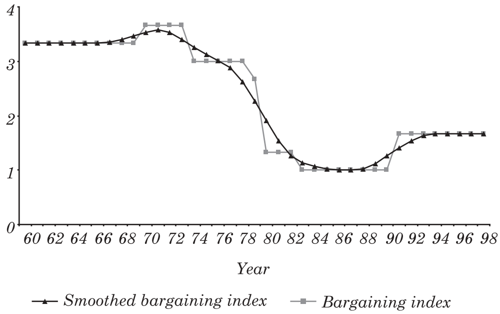
Year	Job security index (1)	Minimum wage: Under 18 y/o (2)	Minimum wage: Over 18 y/o (3)	Bargaining index: Original (4)	Bargaining index: Smoothed (5)	Average wage, By sex: Male (6)	Average wage, By sex: Female (7)	Average wage, By skill level: Low (8)	Average wage, By skill level: High (9)	Average wage, By age group: 15-24 (10)	Average wage, By age group: 25-49 (11)	Average wage, By age group: 50-65 (12)	GDP deviation from trend (%) (13)	Employment rate (%) (14)	Wage employment rate (%) (15)	Self employment rate (%) (16)
1985	2.29	220	220	1.00	1.01	699	480	312	808	323	683	725	−6.19	46.4	36.6	9.8
1986	2.29	215	215	1.00	1.00	653	471	301	742	314	634	731	−5.35	47.0	37.3	9.7
1987	2.29	199	199	1.00	1.00	796	539	288	932	355	764	907	−4.05	50.1	39.5	10.5
1988	2.29	222	222	1.00	1.03	766	542	316	902	376	751	799	−2.93	50.9	38.6	12.2
1989	2.29	293	340	1.00	1.12	869	679	376	981	434	868	973	0.41	53.1	41.6	11.5
1990	2.29	298	346	1.00	1.26	1,003	682	390	1,074	462	960	1,011	−2.83	52.0	40.5	11.4
1991	3.06	278	327	1.67	1.41	971	694	401	1,046	470	951	949	−2.47	53.2	41.2	11.9
1992	3.06	293	340	1.67	1.54	904	726	455	998	503	914	900	1.47	55.7	43.6	12.1
1993	3.06	294	341	1.67	1.64	1,072	832	496	1,158	627	1,054	1,093	0.98	55.9	44.0	11.9
1994	3.06	294	342	1.67	1.67	1,141	840	535	1,194	624	1,101	1,163	−1.22	55.4	42.5	12.9
1995	3.06	302	351	1.67	1.67	1,230	919	566	1,310	657	1,215	1,199	0.81	55.5	42.8	12.7
1996	3.06	279	324	1.67	1.67	1,329	1,047	621	1,412	725	1,283	1,465	1.59	55.8	43.7	12.0
1997	3.06	248	333	1.67	1.67	1,392	1,100	613	1,505	775	1,380	1,335	2.79	56.7	44.1	12.6
1998	3.06	243	341	1.67	1.67	1,356	1,136	759	1,427	792	1,325	1,500	0.70	56.8	43.6	13.2

Source: Authors' calculations (see data section).

of this nature does not exist in Chile. Because union membership is also not available for all years covered in our sample, we measure unions' bargaining power by means of an index that reflects the degree of centralization of collective bargaining constructed by Edwards and Cox-Edwards (2000). This variable takes values from 1 (total decentralization) to 4 (total centralization). The use of this measure is based on the observation that union coverage tends to be larger in countries where collective bargaining is centralized. Finally, we include output gap as a measure of economic activity deviations, with respect to potential GDP. To obtain this variable, we use GDP data from the World Bank and apply a Hodrick-Prescott filter to obtain trend GDP.

Table 3 summarizes some basic statistics of our sample, by year. The first three columns display the value of the job security index and the real minimum wage for people eighteen or younger and for adult workers. The next two columns summarize the index of bargaining (column 4 presents the original index, and column 5 presents the smoothed index). The evolution of these variables over time is depicted in figure 3. Higher values of this measure, like those registered from 1960 to 1970, reflect periods of higher union centralization.[16] The next seven columns summarize the average hourly wage broken down by sex (columns 6 and 7) skill level (columns 8 and 9); and age group

Figure 3. Bargaining Index[a]

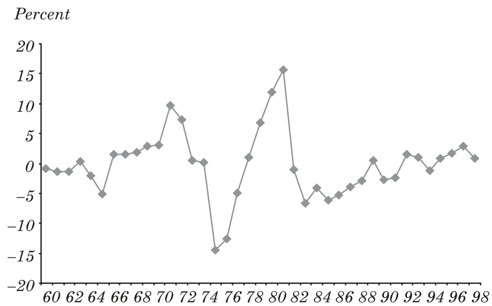


Source: Edwards and Cox Edwards (2000).
a. Bargaining Index measures the degree of centralization of wage bargaining. It takes values from 1 to 4. Higher values indicate higher centralization.

16. Although not shown in the results, we checked the robustness of our results using the strikes index constructed by Edwards and Cox-Edwards (2000) instead of the centralization index. The results were invariant to different specifications.

Figure 4. Employment Rates

Figure 5. GDP Deviation from Trend



Source: Authors' calculations (see data section).

(columns 10, 11 and 12). Column 13 summarizes the deviation of the GDP from its potential or trend value. Finally, columns 14, 15, and 16 present the percentage of total people employed, the percentage of people that work for someone else (wage employment), and the percentage of people self-employed as a proportion of total population between fifteen and sixty-five years old. These three rates are also depicted in figure 4, which, jointly with figure 5 (which shows GDP deviations from its trend), illustrates the violent swings experienced by the Chilean economy during the 1960-1998 period, and in particular between 1970 and 1985.[17] Some additional indicators describing the performance of the Chilean economy are summarized in table 4.

4. Methodology

To estimate the differential impact of labor market regulations across subpopulations we assume that the employment status of an individual is characterized by

$$y^*_{ijt} = X_{it}\beta_1 + X'_{it}Z_t\beta_2 + \gamma_t + \varepsilon_{ijt} \tag{1}$$

where

$y_{ijt} = 1$ if $y^*_{ijt} > 0$

$y_{ijt} = 0$ otherwise ,

and y^*_{ijt} is an unobservable variable that determines whether an individual *i*, in subpopulation *j*, at time *t* will be employed or not, and y_{ijt} is the observable employment status of this individual. The variable takes a value of 1 if the individual is employed and zero if it is not. In some specifications, we focus only on wage employment (and alternatively, self-employment), and, therefore, this variable takes the value of 1 if an individual is wage (self-) employed and zero otherwise. The sample corresponds to the whole population between fifteen and sixty-five years old. In addition, X_{it} is a vector of variables that summarizes the personal characteristics of the individual *i* at time *t*, Z_t is a vector of variables that vary with *t*, γ_t is a year fixed effect, and ε_{ijt} is an error term. Among the personal characteristics, we include age, gender, skill level, number of children, and number of children interacted with gender. In some specifications, we also include age interacted with gender and age interacted with skill to capture differential effects of age across gender and skill groups. Given the number of observations available, we divided the data into three age groups (fifteen to twenty-four, twenty-five to fifty, and fifty-one to sixty-five) and two skill levels (9 years of education or less and more than 9 years). Adding

17. Chilean economic performance has been extensively documented by Edwards and Cox-Edwards (1991, 2000), De la Cuadra and Hachette (1992), Wisecarver D. (1992), Bosworth, Dornbusch, and Labán (1994), Hudson R. (1994), Soto R. (1995), and Cortázar and Vial (1998).

Table 4. General Economic Indicators: Chile 1960-98

Year	*GDP per capita growth (annual %)*	*Inflation, consumer prices (annual%)*	*National unemployment, female (% of female labor force)*	*Santiago unemployment, total (% of total labor force)*	*National unemployment, total (% of total labor force)*	*National unemployment, youth total (% of total labor force ages 15-24)*	*Gini coefficient*
1960	n.a.	n.a.	n.a.	n.a.	n.a.	8.0	42.5
1961	1.5	7.7	n.a.	n.a.	n.a.	7.1	45.2
1962	2.7	14.0	n.a.	n.a.	n.a.	5.7	45.5
1963	3.6	44.1	n.a.	n.a.	n.a.	5.2	n.a.
1964	0.3	46.0	n.a.	n.a.	n.a.	4.9	n.a.
1965	−1.8	28.8	n.a.	n.a.	n.a.	5.0	n.a.
1966	7.6	23.1	n.a.	n.a.	n.a.	6.0	45.2
1967	1.5	18.8	n.a.	n.a.	n.a.	5.9	45.8
1968	1.6	26.3	n.a.	n.a.	n.a.	6.4	48.1
1969	1.5	30.4	n.a.	n.a.	n.a.	7.1	48.0
1970	0.2	32.5	n.a.	n.a.	n.a.	7.0	47.5
1971	7.1	20.0	n.a.	n.a.	n.a.	5.2	47.7
1972	−2.5	74.8	n.a.	n.a.	n.a.	3.7	43.1
1973	−6.5	361.5	n.a.	n.a.	n.a.	3.1	44.1
1974	0.8	504.7	n.a.	n.a.	n.a.	10.3	40.7
1975	−12.8	374.7	n.a.	n.a.	n.a.	16.1	41.1
1976	1.8	211.8	n.a.	n.a.	n.a.	18.0	47.2
1977	7.1	91.9	n.a.	n.a.	n.a.	13.0	48.4
1978	5.9	40.1	n.a.	n.a.	n.a.	12.8	49.8
1979	7.1	33.4	n.a.	n.a.	n.a.	12.5	49.4
1980	6.5	35.1	10.4	10.0	20.8	11.7	49.1
1981	3.2	19.7	11.3	9.9	21.5	9.0	47.3
1982	−11.7	9.9	19.6	18.3	30.5	23.2	51.2
1983	−5.3	27.3	14.6	14.7	24.7	22.7	52.7
1984	6.3	19.9	13.9	n.a.	25.2	18.4	54.2
1985	5.4	29.5	12.1	13.4	22.7	16.2	51.5
1986	3.9	20.6	8.8	9.7	17.3	15.4	48.7
1987	4.9	19.9	7.9	9.3	n.a.	13.5	57.6
1988	5.5	14.7	6.3	7.8	14.3	11.2	53.7
1989	8.7	17.0	5.3	6.1	13.2	9.3	50.8
1990	1.9	26.0	5.7	5.7	13.1	9.7	53.9
1991	6.2	21.8	5.3	5.8	12.7	8.3	52.4
1992	10.4	15.4	4.4	5.6	10.9	6.0	47.4
1993	5.2	12.7	4.5	5.1	11.0	6.4	45.4
1994	4.0	11.4	5.9	6.8	13.2	6.3	45.9
1995	8.9	8.2	4.7	5.3	11.5	6.1	46.3
1996	5.7	7.4	5.4	6.7	12.8	7.2	45.4
1997	6.0	6.1	5.3	6.6	13.0	6.7	n.a.
1998	2.5	5.1	7.2	7.6	16.7	6.9	n.a.

Sources: World Bank World Development Indicators Data Base and Gini coefficient from background data, Montenegro (1998). Note: n.a.= not available.

the skill and the age groups to the gender division, we have twelve different sub-populations, j=1,...12.

In the vector of aggregate variables, Z_t, we include the index of job security, deviations from GDP trend, and the union centralization variable (all in logarithms). We also include the minimum wage index (also in logarithms), but we let it change for individuals eighteen and younger. By construction, the vector of coefficients on the interaction of X_{it} and Z_t, β_2, gives the sign of the *differential* effect. In addition, assuming that the Prob$(y^*_{ijt}>0)$ is distributed as a standard normal distribution, the size of the marginal differential effect is given by $\phi(.)X_{it}\beta_2$, where $\phi(.)$ is the normal density function.

Although specification (1) is a reduced form equation, in some cases it will be useful to add a measure of wages. To construct this variable, w_{ijt}, we assign to all workers i members of j, j=1,.....,12, at period t, the average wage of all employed workers in group j at period t.

Our original intention was to estimate

$$y^*_{ijt} = X_{it}\beta_1 + X'_{it}Z_t\beta_2 + Z_t\beta_3 + \varepsilon_{ijt} \quad . \tag{1'}$$

With such a specification we could recover the *total* marginal effect of a labor policy on subpopulation j as $\phi(.)(X_{it}\beta_2 + \beta_3)$. However, despite finding robust estimates for the differential effects, our estimates for the level effect (β_3) proved to be extremely sensitive to the set of variables included in Z_t, suggesting that our time variables did not properly account for the time variation of the series. In view of these results, we opted for estimating specification (1). This estimation still allows us to compute marginal effects, but the total effects are now absorbed by the constant term. Therefore, we can measure the impact of labor market regulations on the *distribution* but not on the *level* of employment. Nonetheless, estimating equation (1) instead of (1') offers substantial advantages from an econometrics point of view. It allows controlling for macroeconomic trends and cycles as well as policy changes and other unobservable variables that are common to all individuals and that could be correlated to employment and labor market regulations and bias the estimation. In addition to the inclusion of time variables, we minimize the risk of omitted variable biases and spurious correlations in four additional ways.

First, by using individual data from a series of stacked household surveys to estimate specification (1), we can control for changes in the relative size of the population of each group and changes in fertility, which if omitted, could bias our estimates. Second, by controlling for effects of changes in the business cycle (using GDP deviations from its trend) across individuals (that is, including $X'_{it}Z_t$, where Z_t contains the business-cycle variable), we can partially control for changes in policy and institutions that are endogenous to changes in relative employment. This is because such movements are likely to be correlated with changes in the business cycle. Third, by estimating the differential effect of policy while including contemporary labor market policies and institutions, we make sure that our measured effects are not biased by the correlation between these variables and the distribution of employment. Finally, by comparing the estimated effects on the probability of wage employment (which is covered by labor policy) with the results on self-employment (which is not covered) once appropriate pull/push factors from and to self-employment are accounted for, we assess whether we are capturing the effect of policy, or, instead, the effect of some unobservable correlate with group-specific employment.

5. Empirical Results

5.1 The Effect of Job Security on the Distribution of Employment

Our results indicate that job security provisions have a differential impact across demographic subgroups. In table 5, we report the results of estimating our empirical specification (1) assuming normality in the distribution of errors. The reported numbers correspond to the coefficients of the probit model, while the marginal effects for selected sub-populations of workers are reported in table 6. The t-tests, reported next to the coefficients, are robust to the presence of heteroskedasticity of unknown kind using the White (1980) method. Most coefficients on the individual characteristic variables exhibit the expected patterns: female and older workers are less likely to be employed than prime-age (twenty-six to fifty) men. Additionally, the number of children per father increases the probability of being employed, and the number of children per mother decreases the probability of being employed. Instead, the coefficients on the variable young and unskilled change signs across specifications.

In column 1 we report the results of interacting the job security measure with dummies for age (young and older), gender (women)

Table 5. The Effect of Job Security and Minimum Wages, Probit Results

		(1)		(2)		(3)		(4)		(5)		(6)		(7)	
Dependent variable:		*Employed*		*Employed*		*Wage employment*		*Self-employment*		*Employed*		*Employed*		*Employed*	
		β	*t-test*	β	*t-test*	β	*t-test*	β	*t-test*	β	*t-test*	β	*t-test*	β	*t-test*
	Dummy young	−0.895	−104.2	0.492	2.6	0.919	5.0	−0.420	−1.4	−1.170	−6.1	−0.965	−4.9	1.278	9.1
	Dummy old	−0.671	−66.8	−1.651	−7.3	−1.697	−1.4	0.418	1.7	−2.100	−9.1	−2.123	−9.0	−1.410	−8.6
	Dummy women	−0.546	−66.7	−2.026	−12.2	−1.860	−11.6	−0.363	−1.7	−2.411	−14.2	−1.963	−11.3	−2.787	−22.7
	Dummy unskilled	−0.007	0.1	1.864	10.9	1.883	11.2	−0.328	−1.5	1.487	8.6	1.836	10.3	2.287	18.1
	Children per father	0.157	45.0	0.157	44.6	−0.059	25.7	0.027	11.3	0.115	32.0	0.115	31.5	0.156	44.6
	Children per mother	−0.393	−93.9	−0.392	−92.7	−0.315	−86.9	−0.020	−5.4	−0.318	−70.1	−0.316	−68.5	−0.392	−93.1
Interacted with logarithm of job security	Dummy young	−0.094	−10.8	−0.111	−12.7	−0.827	−9.7	−0.016	−1.2	−0.091	−5.6	−0.116	−6.7		
	Dummy old	0.012	1.2	0.020	1.8	0.029	2.7	0.017	1.5	0.025	1.2	0.012	0.6		
	Dummy women	−0.467	−6.1	−0.027	−3.4	−0.002	−0.3	0.027	2.7	−0.055	−4.5	−0.087	−6.8		
	Dummy unskilled	−0.034	−4.2	−0.056	−7.0	−0.073	−9.3	0.034	3.4	−0.038	−3.3	−0.060	−4.8		
	Dummy young* dummy women									0.084	4.7	0.103	5.4		
	Dummy old* dummy women									−0.004	−0.2	0.006	0.3		
	Dummy young* dummy unskilled									−0.038	−2.2	−0.016	−0.9		
	Dummy old* dummy unskilled									0.003	0.2	0.015	0.6		
Interacted with logarithm of minimun wage	Dummy young			−0.140	−8.2	−0.156	−9.3	−0.037	−1.3	−0.011	-0.6	−0.022	−1.2	−0.213	−16.0
	Dummy old			0.091	4.4	0.091	4.4	−0.029	−1.3	0.130	6.2	0.130	6.1	0.072	4.6
	Dummy women			0.146	9.6	0.156	10.7	−0.030	−1.5	0.168	10.8	0.130	8.2	0.210	18.0
	Dummy unskilled			−0.181	−11.6	−0.181	−11.9	0.030	1.5	−0.159	−10.1	−0.181	−11.2	−0.220	−18.3
	Dummy young* dummy women									0.025	11.0	0.022	9.8		
	Dummy old* dummy women									−0.004	−1.3	−0.002	−0.7		
	Dummy young*dummy unskilled									0.040	17.4	0.035	15.2		
	Dummy old* dummy unskilled									0.013	4.9	0.015	5.3		

Table 5. (continued)

		(1)		(2)		(3)		(4)		(5)		(6)		(7)	
		Employed		Employed		Wage employment		Self-employment		Employed		Employed		Employed	
Dependent variable:		β	*t-test*	β	*t-test*	β	*t-test*	β	*t-test*	β	*t-test*	β	*t-test*	β	*t-test*
Interacted with union centralization	Dummy young			0.132	8.2	0.142	9.2	0.080	3.0	−0.301	−13.1	−0.279	−11.9		
	Dummy old			0.027	1.4	0.024	1.2	0.015	0.7	−0.097	−3.2	−0.085	−2.8		
	Dummy women			−0.097	−6.8	−0.122	−8.9	0.080	4.2	−0.245	−13.5	−0.218	−11.6		
	Dummy unskilled			0.076	5.2	0.049	3.4	0.036	1.9	−0.084	−4.6	0.056	−3.3		
	Dummy young* dummy women									0.230	12.3	0.271	10.9		
	Dummy old* dummy women									0.153	5.2	0.136	4.5		
	Dummy young* dummy unskilled									0.349	14.1	0.331	13.0		
	Dummy old* dummy unskilled									0.027	0.9	0.025	0.8		
Interacted with GDP deviation from path	Dummy young			−0.085	−0.9	0.210	2.2	0.021	0.1	−0.293	−1.7	−0.362	−2.1		
	Dummy old			−0.387	−3.1	−0.216	−1.7	−0.004	0.0	−0.790	−3.4	−0.803	-3.4		
	Dummy women			−0.492	−5.5	−0.311	−3.6	0.315	2.7	−0.805	−6.0	−0.896	-6.7		
	Dummy unskilled			0.435	4.8	0.347	3.9	0.078	0.7	0.408	3.2	0.415	3.2		
	Dummy young* dummy women									0.397	2.0	0.502	2.5		
	Dummy old* dummy women									0.386	1.6	0.475	1.9		
	Dummy young* dummy unskilled									−0.246	−1.3	−0.157	−0.8		
	Dummy old* dummy unskilled									0.191	0.8	0.176	0.7		
	Logarithm of hourly wage											0.152	16.9		
	Number of observations	303,945		303,945		303,945		303,945		303,945		295,318		303,945	
	Pseudo R^2	0.20		0.17		0.11		0.08		0.21		0.21		0.20	

Notes: Besides the control variables mentioned in the table, all specifications include yearly dummies (not reported). Standard errors are robust to the presence of heteroskedasticity. The employed dummy variable is defined as 1 if the person is employed and zero otherwise (unemployed or inactive). The wage employment dummy variable is defined as 1 if the person is a dependent employee and zero otherwise (independent, unemployed, or inactive). The self-employed dummy variable is defined as 1 if the person is an employer or if the person works as an independent worker and zero otherwise (dependent, unemployed, or inactive).

and skill level. A negative (positive) sign indicates that periods of more stringent job security provisions are associated with a decline (increase) in the probability of employment of a particular sub-population relative to the omitted category. We find strong age effects. The coefficient on the young- job security interaction is negative and statistically significant, while the coefficient on the older- job security interaction is positive although not statistically significant. Our results suggest that high job security tends to bias the distribution of employment against younger workers. We also find significant effects across the skill divide. The coefficient on the unskilled- job security interaction is negative and statistically significant, suggesting that job security provisions reduce the probability of employment of unskilled workers relative to skilled ones. Lastly, the coefficient on the female- job security interaction suggests a negative effect of job security on the probability of employment of women relative to men.

Column 2 shows the results once we control for the evolution of the minimum wage, union activity, and deviations of GDP with respect to its trend, as well as interaction of these variables with age, gender, and skill dummies. The only difference with respect to column 1 is that the coefficient on the dummy for older workers is now somewhat larger and statistically significant at the 10 percent level, suggesting that job security provisions benefit the employment prospects of older workers relative to prime-age ones. In columns 3 and 4 we report the coefficients resulting from estimating the same specification for wage employment and self-employment separately. Our results are encouraging because they suggest that our findings are driven by policy changes instead of by some unobservable factors correlated with labor policy and employment. The signs and magnitudes of the coefficients for total and wage employment are very similar, except for the coefficients on women. Instead, for self-employment, the coefficients are either not statistically different from zero or going in the opposite direction than for wage employment. This is the case with the coefficients on the gender and unskilled variables, which suggests that more stringent job security regulations increase the probability that women and the unskilled are employed in the self-employment sector relative to men and the skilled.

Column 5 exhibits the results once we allow for further interactions between age, skill, and gender groups. With this finer level of disaggregation we can examine whether the impact of job security is the same across young men and young women, or across young skilled and unskilled workers. These additional variables not only provide a more complete description of the effects of job security on the distribution of employment, but also help to infer the channels through which job security affects that distribution. The coefficients for these additional interaction variables are all statistically significant, and a test for their joint significance strongly rejects the null hypothesis of all the coefficients being zero.

The estimates in column 5 contain some interesting additional information relative to the estimates in columns 1 to 4. We find that an increase in job security tends to reduce the employment probabilities of young men relative to those of young women. However, we also find that this effect is reversed at older ages. Thus, job security provisions seemingly reduce the probabilities of employment of middle-aged and older women relative to those of men in that same age group. Our estimates also suggest that an increase in job security provisions reduces the probability of employment of both skilled and unskilled youth, but the effect is larger for unskilled youth.

Finally, column 6 reports the results of estimating the same specification as in column 5, but in addition controlls by the average wage of each subpopulation group in period *t*. Controlling for the wage level of each group allows us to assess whether some of the observed effects are driven by differences in wage adjustment across subpopulations. Yet the results should be taken with caution because some wage movements may be endogenous to the probability of employment. Overall, we find that holding wages constant does not affect our main results. The only coefficient that changes significance is the interaction between the young unskilled and job security. Holding wages constant reduces the coefficient and the significance of the effect on unskilled youth (relative to more skilled youth). Instead, most of the other coefficients become larger (in absolute value) than the ones reported in column 5. This suggests that more stringent regulations are partly paid by workers in the form of lower wages.

The marginal effects reported in table 6 correspond to the specification reported in column 5 of table 5. They are computed for different combinations of the dummies for gender, age, and skill.

The results indicate that the largest adverse effects are on unskilled youth. However, the effects on skilled youth are also substantial; an increase of 100 percent in job security reduces the probability of employment by 0.066 points (or 6.6 percentage points) for unskilled youth and by 0.0351 for youth skilled workers, relative to prime-age skilled workers. The results in table 6 suggest that skilled prime-age male workers gain relative to all other groups with the exception of

Table 6. Marginal and Total Effects of Labor Market Regulations

	Marginal effects		*Total effects*	
	Job security	*Min. wage*	*Job security*	*Min. wage*
	(1)	*(2)*	*(3)*	*(4)*
Men, 15-25, unskilled	–0.066 (0.000)	–0.052 (0.000)	–0.049	–0.052
Men, 15-25, skilled	–0.0351 (0.000)	–0.004 (0.52)	–0.018	–0.004
Men, 26-50, unskilled	–0.008 (0.001)	–0.036 (0.000)	0.009	–0.036
Men, 51-65, unskilled	–0.004 (0.620)	–0.005 (0.54)	0.014	–0.005
Men, 51-65, skilled	0.008 (0.22)	0.045 (0.000)	0.025	0.045
Unskilled	–0.034 (0.000)	–0.012 (0.09)	–0.017	–0.012
Skilled	–0.015 (0.000)	0.044 (0.000)	0.002	0.044
Women	–0.028 (0.000)	0.046 (0.000)	–0.011	0.046
Men	–0.015 (0.000)	–0.017 (0.000)	0.002	–0.017
Young	–0.039 (0.000)	0.013 (0.08)	–0.022	0.013
Old	–0.008 (0.14)	0.060 (0.000)	0.009	0.060

Note: P-values of the test that the marginal effects are equal to zero are reported in parentheses.

older workers. In addition, the marginal effects suggest that job security policies tend to have more adverse effects on women than on men.

In light of the different theories described in section 1, how do we explain the results presented previously? Although we cannot totally discriminate among different theories, we are at least able to reject some hypotheses. The fact that most of our results remain unchanged when wages are included suggests that the differential effects presented previously cannot be explained by differences in the elasticity of labor supply across demographic groups. The only exception is the larger effect on young unskilled workers, which seems to be driven by a higher labor supply elasticity of this group.[18] Our results also suggest that these differential effects cannot be explained by insider-outsider theories, because in that case the effect would also be through wages. Instead, our results suggest that the differential effects on employment are demand driven: changes in job security provisions bring about changes in hiring and firing rates that selectively affect different types of workers.

A barrier-of-entry effect can explain the negative impact of job security on the employment rates of young workers relative to other demographic groups. However, it cannot account for the estimated differences in impact between young women and young men. One possible way to explain these findings is to consider differences in turnover rates across groups. As discussed in section 1, a higher exogenous turnover rate can bring about two effects. On the one hand, workers with a higher propensity to rotate have lower average tenures and therefore are more likely to be laid off in bad times. On the other higher rotation reduces expected severance payments and, therefore, increases the incentives to hire these workers. Consequently, higher rotation among women can explain why job security provisions affect young women less than young men. It can also explain why middle-aged and older women benefit less from job security than men of the same age.

Differences in turnover rates could also partially explain the results for skilled and unskilled workers. Higher rotation among the unskilled would imply lower tenure rates and higher probabilities of dismissal for middle-aged and older unskilled workers, relative to more skilled ones. This is consistent with the deleterious effect of job security on the employment rates of middle-aged and older unskilled workers, relative to skilled ones. Of course, the higher turnover rates among unskilled workers are less likely to be exogenous to the decisions of employers than female turnover rates. In consequence, a complete discussion of this effect requires a model that explains why turnover rates are different in the first place. This model does not seem to be able to explain why the effect on employment appears more negative on the unskilled than on skilled youth, but as we have seen, this

18. Cowan et al. (2003) find that, in Chile, seemingly high transitions between schooling and the labor market lead to a very elastic labor supply for the young unskilled.

effect seems to be driven by a relatively more inelastic labor supply of the latter.

5.2 Distribution of the Effect of Minimum Wages

Table 5 also reports the results of interacting personal characteristic dummies with the evolution of minimum wages over time. An increase in the statutory wage has qualitative effects on the distribution of employment across age and skill that are similar to the qualitative effects of stricter job security provisions. To account for contemporary employment policies and economic conditions, we include measures of union activity, job security provisions, and GDP deviations, interacted with demographic dummies in all specifications in columns 2 to 6, but not in column 7. As in other studies for developed countries, the results in column 7 suggest that an increase in the minimum wage reduces the employment prospects of young workers relative to older ones. We also find a negative effect on the unskilled. Instead, our results also indicate that minimum wage hikes may increase the probability of employment for women relative to men.

Controlling for the subgroup effects of contemporary changes in policy and the business cycle does not alter the results reported in column 7.[19] The comparison between the results obtained from the wage employment and the self-employment specifications (columns 3 and 4) is also encouraging. As with the coefficients associated with job security provisions, we find that the coefficients on wage employment are very similar to the ones obtained for total employment, while the coefficients on self-employment are not statistically significant. All in all, these results suggest that the effects we are capturing are indeed associated with changes in policy rather than with some unobservable correlate of employment across demographic groups.

In column 5 we present our results once we allow for differential effects across age-skill and age-gender categories and control for contemporaneous changes in policy and economic conditions. As in column 7, we find a negative effect of minimum wages on the employment probabilities of unskilled workers. The effect of minimum wages is negative for young unskilled workers and not statistically significant for young skilled ones. Instead, higher minimum wages tend to shift employment toward older workers. Finally, we find that women, and in particular the young, tend to benefit from minimum wage policies.

19. See column 3 as well.

The former specification assumes that the effect of raising the minimum wage is unrelated to the level of the going wage. However, it is plausible that the effect may be positively related to the distance between the statutory and the going wage. To account for this possibility, we include average wages, computed as described in section 5.[20] The results reported in column 6 indicate that controlling for the time evolution of the average wage of subpopulation $j = 1,\ldots,12$ does not alter the results reported in columns 3 to 5.

Column 2 in table 6 summarizes the marginal effects, which give an estimate of the magnitude of the effects on different demographic groups. A 10 percent rise in the minimum wage reduces the employment probability of young unskilled workers by 0.005 (0.5 percentage points). While the effects on youth skilled workers are insignificant, the results indicate an adverse effect on prime-age unskilled workers. This is an interesting result in the context of a literature that almost exclusively focuses on the effects on youth workers.

While most of our findings are consistent with the competitive model, some are difficult to explain with this paradigm. For instance, this model cannot explain why minimum wages tend to shift employment toward women. One possible interpretation is that while men are able to obtain wages that are close to the competitive ones, women's wages are below their marginal products. This would be consistent with the systematic wage gaps found between observationally identical men and women and with the asymmetric gender effects of minimum wages. If wage gaps are explained by imperfect competition in female labor markets, employers are supply constrained when hiring women. Therefore, an increase in minimum wages reduces the demand for male workers and increases the supply of labor for women.

5.3 Total Effects

In our previous results, all the estimated coefficients measured the effects of labor regulations on each particular subpopulation relative to the omitted category, but they did not provide information on whether the employment probabilities of the different subgroups increased or declined in absolute terms after changes in policy. In

20. Including such variables is tantamount to including a set of noncoverage adjusted, demographic group-specific Kaitz ratios. However, we are not imposing the constraint that the coefficient on the minimum wage is the same as the coefficient on the group-specific average wage.

this section, we attempt to gauge the total effects of labor market policies on the probability of employment by estimating their effect on the aggregate employment rates of prime-age skilled men (the omitted category in the specifications reported in table 5). To do so, we estimate the following error correction specification:

$$\Delta N_t = C - \lambda\left(N_{t-1} - N_t^*\right) + B_1\left(y_t - y_t^*\right) + B_2 \Delta Log\left(W_t\right) + B_3 \Delta N_{t-L} + \varepsilon_t \,, \tag{2}$$

where

$$N_t^* = \gamma_0 + \gamma_1 Log\left(JS_t\right) + \gamma_2 Log\left(MW_t\right) + \gamma_3 Log\left(Union_t\right) \,, \tag{3}$$

and where N_t denotes the employment rate—that is, the employment to population ratio—of prime-age male skilled workers in period t, N_t^* denotes long-run equilibrium employment, $y_t - y_t^*$ denotes GDP deviations from its trend (in logs), W_t denotes average wages for prime-age skilled male workers, JS_t denotes the measure of Job Security, MW_t denotes minimum wages, $Union_t$ denotes the index of wage bargaining and L is the length of the maximum lag. In expression (2), employment changes are a function of: previous period deviations from long-run equilibrium employment; GDP deviations from its trend; changes in wages and short-run dynamics. Expression (3) assumes that, in the long run, employment rates are a function of labor market policies and the structure of wage bargaining.

Using aggregate time series techniques to estimate the effect of policies on the reference group allows us to model short- and long-run employment dynamics. The first step in the estimation of expression (2) and (3) is to test whether the variables are stationary. The first panel in table 7 reports the results of testing for the presence of unit roots using the Augmented Dickey-Fuller test (ADF). The tests are specified with three lags. In those cases in which the plot of the series indicated the presence of a time trend, we included a constant and a time trend in the specification; in the other cases, we included only a constant. While we can reject the unit root hypothesis for GDP deviations from its trend and for changes in hourly wages, we cannot reject nonstationarity for the lagged employment rate, the logarithm of minimum wages, the logarithm of the job security index, and the logarithm of union centralization. However, ADF tests on the first differences of these four series indicate that the hypothesis that these series are integrated of order one, I(1), is not rejected.

Table 7. Unit Root and Coeintegration Tests

Name of the series	*Symbol*	*Specification*	*ADF test statistic*	*5% critical value*
GDP deviation from its trend	$y–y*$	Constant	–4.84	–2.95
Wage growth	$\Delta(logW)$	Constant	–3.85	–2.97
Logarithm minimum wage	$L(MW)$	Trend	–1.47	–3.54
Logarithm job security	$L(JS)$	Constant	–2.43	–2.95
Logarithm union centralization	$L(Union)$	Trend	–2.76	–3.54
Lagged employment rate	N_{t-1}	Constant	–1.67	–2.95
First diff.lagged emp. rate	ΔN_{t-1}	Constant	–3.04	–2.95
Change in log minimum wage	$\Delta L(MW)$	Constant	–2.56	–2.95
Change in log job security	$\Delta L(Index)$	Constant	–2.66	–2.95
Change in log union	$\Delta L(Union)$	Constant	–2.34	–2.95

Source: *Johansen Cointegration Test:*

Series	*Likelihood ratio*	*5% Critical value*	*Hypothesized number of CE*
Nt–1	108.64	53.12	None **
Log (MW)	60.35	34.91	At most 1**
Log (JS)	24.64	19.96	At most 2*
Log (Union)	5.26	9.24	At most 3

* (**) *denotes rejection of the hypothesis at the 5 (1) percent significance level.*

Given the nonstationarity of the employment rate, expression (2) is well defined only if lagged employment deviations, with respect to the long-run equilibrium rate, are stationary. This is equivalent to saying that the series N_t^* has to cointegrate with N_{t-1}. The second panel in table 7 reports the results of the Johansen cointegration test between N^* and N_{t-1}. The likelihood ratio test indicates the presence of three cointegrating equations, indicating that the error correction model is well defined.

Table 8 presents the results of estimating the error correction model (ECM) once expression (3) has been substituted into expression(2). We use the results of the Akaike's Information Criteria (AIC) test to determine the optimal length of the lagged endogenous variable and determine that $L = 1$. We estimate the ECM with and without wages to see whether introducing wages alters our results, and we find the results to be very similar in both cases. Essentially, we find that job security provisions increase the long-run equilibrium rate of prime-age skilled male employment. This is not totally surprising. As mentioned in section 2, job security provisions increase the cost of dismissing workers with long tenure relative to the costs of dismissing less-tenured workers, reducing the layoff rate of the first relative to the layoff rate of the latter. Because prime-age skilled

Table 8. Level Effects on Male Prime-age Employment

Independent variables	*(1)*	*(2)*
N_{t-1}	–0.63	–0.66
	(–3.05)	(–3.24)
Deviations GDPt	0.08	0.10
	(1.21)	(1.48)
Δ log Wt		0.02
		(0.84)
Log (JS)	0.01	0.02
	(1.80)	(2.23)
Log (MW)	–0.01	–0.01
	(–0.93)	(–1.13)
Log (Union)	0.03	0.03
	(1.54)	(1.45)
Constant	0.61	0.65
	(3.55)	(3.92)
ΔN_{t-1}	0.28	0.24
	(1.48)	(1.30)
No. of observations	37	35
Adjusted R^2	0.16	0.23
Long-term effect of JS	0.017	0.023
Long-term effect of MW	0	0

Note: *t-statistics* shown in parentheses.

workers tend to have longer tenures than other, younger, less skilled-workers, job security provisions reduce the layoff rates of prime-age skilled workers relative to the layoff rate of other demographic groups. The positive sign in the ECM suggests that this effect on the layoff rate more than compensates for the negative effect of job security on employment creation. Instead, we do not reject the hypothesis that an increase in the minimum wage does not affect the employment rate of prime-age, skilled male workers, regardless of whether we control for the evolution of wages.

The estimated effect of job security provisions and minimum wages on the employment rate can be used to infer the total effect of these regulations on the employment probabilities of other demographic groups. In order to do so, the coefficients on job security provisions and minimum wages, reported in table 8, should be divided by (minus) the coefficient on the lagged employment variable to obtain the coefficients in expression (3). They reflect the magnitude of the long-run effect of regulations on prime-age skilled male employment. The third and fourth columns of table 6 present our estimates for the total effects. They are obtained by adding the marginal effect reported in the first and second columns of table 6 to the long-run elasticities obtained from specification (1) in table 8.[21]

The total effects reported in columns 3 and 4 suggest that job security provisions not only shift the distribution of employment toward older and skilled workers, but also increase their employment rates. Instead, more stringent job security provisions reduce the employment rates of young workers. Moreover, job security provisions reduce employment opportunities for women while increasing those of men. The magnitudes of these estimated effects are substantial. According to them, the 1990 labor reform, which increased our measure of job security by about one-third, reduced the employment rates of young unskilled male workers by 1.6 percentage points of the population.

We also find non-neutral effects of minimum wage spikes. Our estimates suggest that a 10 percent increase in minimum wages reduces the probability of employment for young unskilled male workers by 0.52 percentage points. Lastly, we find that a 10 percent increase in the minimum wage raises the employment rates of women by 0.46 percentage points.

6. Conclusions

The effect of regulations on employment is far from neutral across demographic subgroups. Paradoxically, job security and minimum wage regulations appear to be detrimental to the employment opportunities of the workers that they are supposed to help. Our results suggest that both minimum wages and job security regulations reduce the employment opportunities of the young and the unskilled—and particularly unskilled youth—while promoting the employment rates of skilled and older workers. We have also found indications that job security regulations may force some workers, particularly women and the unskilled, out of wage employment and into self-employment[22] .

21. The long-run effect of job security on the employment rates of middle-age skilled workers is computed as 0.011 divided by 0.63, which is equal to 0.017.

22. This paper has only examined the effects on employment. A complete analysis of who benefits and who loses from regulations would require examining the effects of regulations on the distribution of wages and benefits as well.

There is an ongoing debate on whether raising minimum wages and job security provisions have any effects on aggregate employment rates. However, even if researchers concluded that job security provisions or minimum wages do not have an effect in the aggregate, it is important to carefully consider these distributional effects when evaluating their desirability. At best, these policies will help some disadvantaged workers although perhaps at the expense of other poor workers. At worse, they distribute jobs from less advantaged to better-off workers.

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UNEMPLOYMENT-POVERTY TRADEOFFS

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Reducing unemployment and alleviating poverty are key policy goals in many developing countries, yet progress remains elusive on both fronts. Although the measurement of poverty and the use of international poverty lines for cross-country comparisons have generated much controversy in recent years (see Deaton, 2001, 2003; Ravallion, 2003), there is some agreement that poverty remains high in many parts of the world and has even increased in some countries. Figure 1 displays the behavior of the headcount ratio, which measures the incidence of poverty (that is, the proportion of individuals or households earning less than a given level of income), in various regions of the developing world, using international poverty lines of $1.08 and $2.16 a day.[1] World Bank data show that between 1990 and 1999, poverty rates fell significantly in East Asia and the Pacific, but they increased in Europe, Central Asia, the Middle East, and North Africa, while countries in Latin America and the Caribbean, South Asia, and Sub-Saharan Africa recorded very little progress. According to the United Nations, poverty rates (measured by the proportion of a country's people living below $1.08 a day) increased in the 1990s in thirty-seven out of sixty-seven countries for which data were available (UNDP, 2003).[2] Projections for

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1. Let y^* be the poverty line; the headcount ratio is defined as $PH = n/N$, where n is the number of households below the poverty line and N is the total number of households.

2. Fifty-four countries also recorded an average growth rate below zero for the last decade, and twenty-one countries experienced a drop in the human development index (a more comprehensive measure of welfare calculated by the United Nations, which includes life expectancy and literacy). Twelve countries registered a decline in primary school enrollment rates, and fourteen countries recorded an increase in child mortality.

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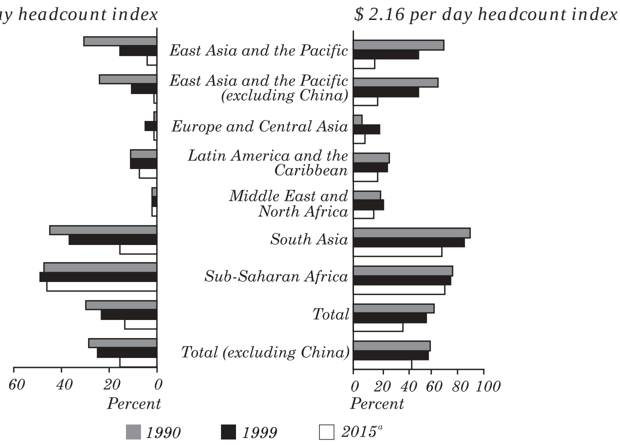
2015 based on current trends indicate that prospects remain bleak for Sub-Saharan Africa.

Unemployment has also become a great source of concern. Women and the young have been particularly hard hit, because their jobs are highly vulnerable to adverse economic shocks. The International Labor Organization (ILO) estimates that the number of unemployed workers worldwide grew by 20 million between the beginning of 2001 and the end of 2002, to reach a record level of 180 million (ILO, 2003). As shown in figure 2, only in transition economies did unemployment rates fall in recent years. They remain well above 10 percent in several countries—and are even close to 20 percent in Bulgaria, Poland, the Slovak Republic, and the former Yugoslavia—despite strong economic growth in recent years.[3] In Latin America, many countries (including those with sustained growth) have experienced major increases in unemployment: the unemployment rate doubled to more than 10 percent in Argentina and Brazil in the 1990s. In the Middle East and North Africa, the population nearly quadrupled during the second half of the past century, and employment growth failed to keep pace with the resulting expansion of the labor force in the 1980s and 1990s.

Figure 1. Poverty Headcount Index, 1990–2015



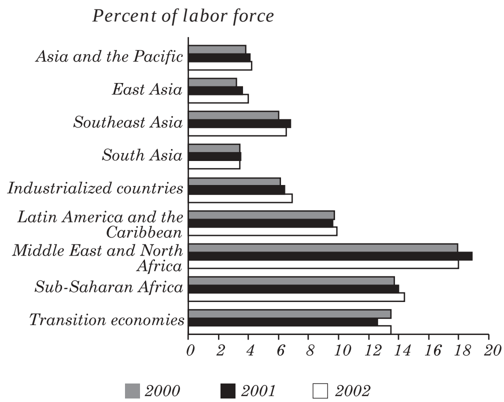
Source: World Bank.
a. Projections.

3. Unemployment was nonexistent at the beginning of the 1990s in Central and Eastern Europe, but it jumped to about 15 percent of the labor force in the early phases of the transition to a market economy.

Consequently, the Middle East and North Africa recorded some of the highest unemployment rates among developing regions in the 1990s. According to the ILO, unemployment rates range from less than 3 percent in the United Arab Emirates to close to 30 percent in Algeria. In 2001, the number of unemployed in the region—mostly the young (or first-time job seekers) and women—was estimated to be over 22 million, or 17.6 percent of the labor force.[4] Based on current trends, prospects remain bleak. The United Nations estimates that the population in the Middle East and North Africa is likely to continue to grow faster than in any other region between 2000 and 2015; the labor force is expected to grow at a rate of about 3 percent, such that unemployment could exceed 25 million by the year 2010 (UNDP, 2002).

Unemployment reduction and poverty alleviation are often viewed as complementary policy goals that involve no tradeoffs. This is not always the case, however. The experience of recent years shows that vulnerable groups (namely, young people, older workers, women, and the unskilled) frequently benefited little from improvements in aggregate macroeconomic conditions, and they often ended up in poorly paid jobs. In Latin America, the share of the so-called working poor (that is, workers who earn less than the $1.08-a-day international poverty line) in total employment rose significantly in many countries.

Figure 2. Unemployment Rates by Region, 2000–02



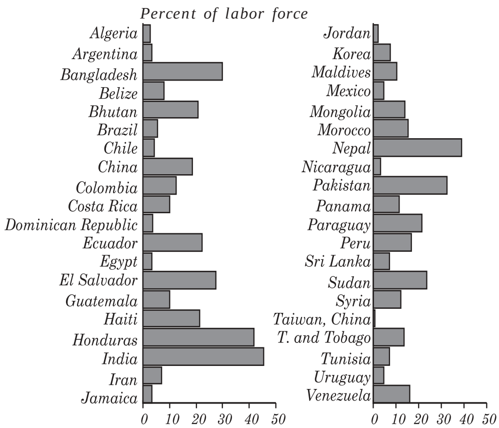
Source: International Labor Organization.

4. In Egypt, for instance, the unemployment rate for women (22.6 percent) is four times that of men, and in Jordan it is almost double. The youth unemployment rate is almost 39 percent in Algeria and exceeds 73 percent in Syria (ILO, 2003). The World Bank (2004) reports a regional unemployment rate of 14.9 percent for 2000–01, with 20 million unemployed.

In Sub-Saharan Africa and South Asia, measured unemployment remains relatively low, yet the share of the working poor in total employment averaged almost 40 percent in both regions, reaching 50 percent in India (ILO, 2003) (see figure 3). In the Middle East and North Africa, the proportion of working poor is also high, as is the case, for instance, in Morocco and Syria. A potential tradeoff between unemployment reduction and poverty alleviation is thus readily apparent: to the extent that the increased growth rates of output and job creation that are needed to absorb the increased labor supply and reduce unemployment require a significant drop in real wages, the deterioration in living standards may lead to a rise in poverty.

Various other sources of potential tradeoffs between reducing poverty and lowering unemployment may arise in both the short and the long term. This paper provides a systematic assessment of the factors that may entail an arbitrage between these two key policy goals. Section 1 presents a broad analytical discussion of the conditions that may trigger unemployment-poverty tradeoffs, focusing in particular on the role of labor market reforms such as a cut in payroll taxes on unskilled labor, a reduction in the minimum wage, and a reduction in firing costs. Section 2 proposes two econometric techniques for empirically assessing the importance of unemployment-poverty tradeoffs. The first is based on a vector autoregression (VAR) model linking the cyclical components of output, real wages, unemployment, and poverty.

Figure 3. Proportion of Working Poor[a]



Source: International Labor Organization.

a. The working poor are workers who do not earn enough to lift themselves and their families above the US$1.08-a-day poverty line.

The second involves cross-country regressions of the determinants of poverty rates, with the unemployment rate among the explanatory variables. Section 3 proposes a third approach, based on a simulation model that integrates a structural macroeconomic component and a household survey to assess the impact of policy shocks on unemployment and poverty. The analysis focuses on labor market reforms as a source of shocks and studies their impact on the composition of both unemployment (skilled and unskilled) and poverty (with a distinction between various categories of urban households). Many economists regard labor market rigidities as being a major obstacle to an expansion of employment in the formal economy and a reduction of urban poverty, which tends to be concentrated in the informal sector.[5] At the same time, the possible existence of tradeoffs between unemployment and poverty reduction has received scant attention in the analytical literature focusing on these reforms. The framework presented in this paper is particularly useful because it allows a description of the transitional dynamics induced by policy shocks. It is therefore possible to assess not only whether such shocks entail the existence of a short-term tradeoff between unemployment and poverty reduction, but also whether this tradeoff tends to persist over time. The last part of the paper offers some concluding remarks and identifies some research perspectives.

1. Sources of Tradeoffs

At the level of an individual country, a tradeoff between poverty and unemployment can surface either at the aggregate (economy-wide) level or at the level of individual household groups (for instance, urban households). Tradeoffs at both levels may entail a temporal dimension, in the sense that they may emerge in the short term but vanish in the long run (or vice versa). This section analyzes the conditions under which aggregate and partial tradeoffs between unemployment and poverty may arise. It also draws the implications of the analysis for predicting and interpreting the correlation between these two variables across countries.

As noted earlier, an obvious reason for an inverse correlation (or the lack thereof) between poverty and unemployment is based on the possibility that reducing unemployment requires a fall in real wages;

5. See, for instance, Saavedra (2003) for a review of the Latin American experience with labor market reform in the 1990s.

this lowers real income and therefore leads to an increase in poverty. The tradeoff may be particularly steep if the expansion in employment (induced by lower real wages and output growth) is skewed toward low-paying jobs. The increase in the number of working poor appears consistent with this interpretation, although the concomitant increase in unemployment observed in some countries would suggest that real wages did not fall sufficiently or that the labor supply expanded simultaneously. Put differently, an increase in the number of working poor induced by lower wages does not necessarily imply an inverse correlation between poverty and unemployment; it depends on the magnitude of the fall in wages and on the strength of the "encouragement effect" associated with higher growth and employment on participation rates.

This discussion suggests, however, that unemployment and poverty are jointly endogenous—and if unemployment and poverty are simultaneously determined, the correlation between them will be driven by factors that are likely to vary over time or across countries, depending on the sources of shocks that prove to be dominant. Although adverse wage shocks may be an important source of negative correlation between unemployment and poverty over time (and across countries or regions), other sources of shocks to labor demand may also matter. In general, if the economy's aggregate production function is not separable in (all) inputs, the demand for labor will depend not only on the cost of labor, but also on all the variables other than labor affecting output, including overall productivity and inputs such as physical capital and imported raw materials. Productivity shocks, in particular, may also affect the unemployment-poverty correlation, either positively or negatively. A positive productivity shock, for instance, may raise labor demand and put upward pressure on wages, thereby lowering both unemployment and poverty (if the increase in the wage rate is sufficient to raise it above the poverty line). But if wages cannot adjust as a result, say, of a binding minimum wage, then the number of working poor may rise. In that case, although unemployment may fall, overall poverty rates may increase if the minimum wage is below the poverty line.

The underlying source of these shocks (whether to wages or productivity) may be policy induced, rather than the result of purely random disturbances. Consequently, changes in real wages and productivity may themselves be endogenous and may need to be analyzed jointly with changes in poverty and unemployment. Policies aimed at improving labor market flexibility, for instance, may indeed entail a tradeoff between unemployment and poverty, through their impact on wages and labor demand. Labor market regulations, particularly job security provisions, have been shown to have a major impact on both the level and distribution of employment in many developing countries.[6] An increase in employment subsidies, for instance, may have a direct, beneficial impact on unskilled employment; at the same time, it may increase poverty if it is financed by an increase in the sales tax on goods sold domestically, because of the impact of the tax hike on the cost of living. Thus, although the subsidy may increase the nominal (and product) wage of the unskilled, their real (consumption) wage may fall. The impact may be particularly large for the poorest households in urban areas, depending on the exact nature of the tax that is used to offset the impact of the increase in spending on the budget (whether it is indeed an increase in the sales tax or a rise in income tax on individuals or firms) and on the composition of household spending. It is possible for poverty to increase in the informal sector (because workers in that sector bear the brunt of the increase in consumer prices, for instance), while at the same time unskilled unemployment falls in the formal economy.

A reduction in the payroll tax on unskilled labor (a policy that is often advocated to reduce unemployment) may have similar results. If the reduction in the payroll tax is financed by a mixture of higher taxes on domestic goods and corporate income, and if the reduction in the net rate of return on physical capital accumulation lowers investment incentives, then the net effect on employment may be mitigated. The demand for labor may not increase over time as much as it would otherwise because of the gross complementarity between capital and labor. Unemployment may thus fall to a limited extent, whereas poverty among the most vulnerable urban groups may increase significantly—again, because higher taxes on domestic goods have a large impact on the cost of living faced by that category of households.

Even labor market reforms that do not have a direct impact on the government budget may entail a tradeoff between unemployment and poverty, as a result of their indirect, general equilibrium effects. A cut in the minimum wage, for instance, may indeed increase the demand for unskilled labor in the urban formal sector; poverty may increase, however, if the cut is large and the elasticity of demand for that category of labor is not high. To the extent that the cut in the

6. See Heckman and Pagés (2000) and Saavedra (2003) for the case of Latin America.

minimum wage reduces the expected wage (because the employment ratio does not rise sufficiently to offset the reduction in labor income), it may also lower the incentive to queue for employment in the formal economy. Consequently, the supply of labor in the informal sector may increase, thereby putting downward pressure on wages there. Urban poverty rates may therefore increase, although in general the effect is ambiguous.[7]

In a growth context, an ambiguous correlation between unemployment and poverty may also emerge from the combination of an inverse correlation between growth and poverty (a sufficient condition for which is a distribution-neutral growth process) and an ambiguous relationship between growth and unemployment, depending on the source of the underlying shock. The source of ambiguity is well illustrated in a simplified version of the model developed by Bean and Pissarides (1993), which considers a two-period economy with overlapping generations and a constant population.[8] Suppose that production in each individual firm in this economy, Y_t, exhibits constant returns to scale in the firm's capital, K_t, and diminishing returns to labor:

$$Y_t = K_t n_t^{\alpha} \, , \tag{1}$$

where $0 < \alpha < 1$ and

$$n_t = \frac{\bar{K}_t N_t}{K_t} \, ,$$

with N_t denoting the firm's employment level and $\bar{K}_t$ the economy-wide stock of capital (which is treated as given by individual firms). Capital depreciates fully in a single period. Technology thus exhibits positive externalities, in Romer-like fashion.

Potential workers and employers have to search for each other, with the number of successful matches increasing in both the number of unemployed and the number of job vacancies. This matching process takes place at the start of the period, and individuals who fail to find a job then have no chance to reenter the labor market later. Given the generational structure, this implies that all matches last

7. The transmission process of a cut in the minimum wage is studied more formally in the context of the structural model described later.

8. The simplifications involve abstracting from intertemporal considerations in household decisions and choosing a specific functional form for the production technology.

exactly one period. The matching technology for aggregate employment, $\bar{N}_t$, may thus be written as

$$\bar{N}_t = m\left(\bar{V}_t, L_t\right), \tag{2}$$

where $\bar{V}_t$, is the aggregate number of job openings at the start of period t and L_t is the number of young households. The matching function is concave, homogeneous of degree one, and increasing in both arguments.[9] These properties can be summarized by the following restrictions:

$$m_i > 0,\ \ m_{ii} < 0,\ \ m\left(0, L_t\right) = m\left(\bar{V}_t, 0\right) = 0\ ,$$

$$\lim_{\bar{V}_t \to \infty} m\left(\bar{V}_t, L_t\right) = L_t\ \text{, and}\ \lim_{L_t \to \infty} m\left(\bar{V}_t, L_t\right) = \bar{V}_t\ .$$

Because the population is constant, one can set $L_t = 1$ and suppress it in what follows, so that

$$m\left(\bar{V}_t, 1\right) = m\left(\bar{V}_t\right).$$

Thus, $\bar{N}_t$(or, respectively, $\bar{N}_t - 1$) can be interpreted as the economy-wide employment (or unemployment) rate.

Hires by an individual firm, N_t, are proportional to the number of vacancies it has relative to the aggregate, that is,

$$N_t = \left(\frac{V_t}{\bar{V}_t}\right) m\left(\bar{V}_t\right). \tag{3}$$

Households are endowed with one unit of labor, which is supplied inelastically in the first period of life. Their propensity to save when young is assumed constant and equal to $0 < \gamma < 1$. In the second period of their lives, households become entrepreneurs and invest directly. A firm's profits, Π_t, are given by

$$\Pi_t = K_t n_t^{\alpha} - w_t N_t - q_t V\ , \tag{4}$$

9. Concavity is assumed in order to capture a congestion externality in the labor market. The higher the number of vacancies opened by firms, the shorter is the search effort of unemployed workers; and the more unemployed workers searching for jobs in the labor market, the faster is the match available for each firm.

where w_t is the wage rate and q_t is the hiring cost per job opening, which is assumed to be proportional to the economy-wide capital stock, $\bar{K}_t$:[10]

$$q_t = \chi \bar{K}_t . \tag{5}$$

The wage is determined after a match has occurred, as the outcome of a Nash bargain between the firm and the individual worker. Workers can only work at one firm, and if both parties fail to reach agreement, neither has the opportunity to look for an alternative match elsewhere.[11] The firm's utility is linear in the marginal profit from employing an additional worker, that is, using equation (1),

$$\alpha \bar{K}_t n_t^{\alpha-1} - w_t \ .$$

Thus, using the wage rate as a measure of the worker's surplus, and assuming that the unemployed receive no benefit and have no alternative source of income, the wage must satisfy

$$w_t = \text{Arg max}\, w_t^{\beta} \left(\alpha \bar{K}_t n_t^{\alpha-1} - w_t \right)^{1-\beta} ,$$

where $0 < \beta < 1$ measures the worker's bargaining strength. This equation yields the first-order condition

$$\beta w_t^{-1} \left(\alpha \bar{K}_t n_t^{\alpha-1} - w_t \right) - (1-\beta) = 0 ,$$

from which the equilibrium wage can be derived as

$$w_t = \alpha \beta n_t^{\alpha-1} \bar{K}_t \ . \tag{6}$$

10. In this setting only firms incur a cost to match workers with their opened vacancies; workers passively wait for a match, comparing their prospective income with the opportunity cost of being unemployed. An alternative approach, following King and Welling (1995), would be to assume that workers bear a direct cost when they decide to actively search for a job. This assumption would be more appropriate for developing economies, where the lack of adequate institutions in the labor market may create informational frictions.

11. This assumption can be relaxed (by assuming instead that it is costly for each agent to change an alternative match) without affecting qualitatively the main results of the model.

Substituting equation (6) into equation (4) and eliminating V_t using equations (2) and (3), together with equations (1) and (5), yields

$$\Pi_t = K_t \left\{ n_t^{\alpha} - \left[\alpha\beta n_t^{\alpha-1} + \chi \frac{m^{-1}\left(\bar{N}_t\right)}{\bar{N}_t} \right] n_t \right\} = K_t \left[\left(1-\alpha\beta\right) n_t^{\alpha} - \chi \frac{m^{-1}\left(\bar{N}_t\right)}{\bar{N}_t} \right.$$

The firm's optimal choice of n_t thus satisfies

$$\frac{d\Pi_t}{dn_t} = \alpha\left(1-\alpha\beta\right) n_t^{\alpha-1} - \chi \frac{m^{-1}\left(\bar{N}_t\right)}{\bar{N}_t} = 0 \ .$$

With a large number of identical firms, and in general equilibrium, $K_t = \bar{K}_t$, $N_t = \bar{N}_t$, and $n_t = N_t$. The above expression thus becomes

$$\alpha N_t^{\alpha-1} = \frac{\chi}{1-\alpha\beta} \frac{m^{-1}\left(N_t\right)}{N_t} \ , \tag{7}$$

which equates the marginal product of labor to an expression that captures both the marginal cost of matching capital and labor and the strategic use of employment by the firm to affect the outcome of the wage bargain (high employment lowers the marginal product and thus also the wage).

Finally, the evolution of the capital stock is determined by the savings of the young, that is, given the assumption of a full depreciation of capital, $K_{t+1} = \gamma w_t N_t$. Using equation (6) with $K_t = \bar{K}_t$ and $N_t = \bar{N}_t$ yields

$$\frac{K_{t+1}}{K_t} = \gamma\alpha\beta N_t^{\alpha} \ . \tag{8}$$

The growth rate of output (or, equivalently here, output per capita) along a balanced growth path with a constant employment rate is $K_{t+1}/K_t - 1$, which is obtained from equation (8). Thus, equations (7) and (8) determine the economy's equilibrium in terms of the employment rate and the growth rate of output.

This framework can be used to analyze the impact of various changes in the parameters along balanced growth paths.[12] A reduction in hiring costs, χ, raises employment, the rate of capital formation, and growth. An increase in the propensity to consume (a reduction in γ) lowers the growth rate but has no effect on employment. The first experiment predicts a negative empirical (cross-sectional) relation between growth and unemployment—and thus a positive relation between the latter variable and poverty—if differences in growth rates are primarily due to differences in hiring costs across countries. By contrast, no systematic relation should be observed if cross-country differences result from differences in saving rates.[13]

An increase in the relative bargaining strength of workers, β, has two opposite effects. On the one hand, from equation (7), it tends to reduce employment and the growth rate, under reasonable conditions.[14] On the other, it tends to increase the growth rate, with no effect on employment. The effect on growth is thus ambiguous. Intuitively, these two effects can be explained as follows. On the one hand, the increase in bargaining strength shifts income from entrepreneurs (who consume all their income here) to workers, which raises savings and fosters growth. On the other, provided that the strategic effect is not too strong, unemployment rises, thereby reducing workers' income and the available pool of savings and dampening growth. The overall impact on growth (and thus poverty) depends on which effect dominates.

The recent growth literature features several other models that may lead to a negative correlation between unemployment and poverty, as a

12. Exercises of this type are complicated, in general, because changes in parameters will generally affect the rate of return and thus the propensity to save. However, these changes are simpler to analyze here because of the assumption of a constant saving rate.

13. In the present framework, an exogenous reduction in the saving rate has the conventional classical effect of lowering investment and reducing the growth rate. Bean and Pissarides (1993) develop a two-sector extension of this model based on imperfect competition in the consumption goods sector, which implies (in the Keynesian tradition) that an increase in the propensity to consume raises both investment and growth.

14. This is most easily shown if the matching technology involves constant elasticity of substitution (CES), that is,

$$\bar{N}_t = \left(\bar{V}_t^{-\rho} + L_t^{-\rho}\right)^{-1/\rho},$$

with $\rho > 0$. The resulting equation (7) may yield multiple solutions, but the implicit function theorem can be used to show that an increase in β does reduce employment.

result of a nonlinear relation between unemployment and growth. These models include Aghion and Howitt (1994), Cahuc and Michel (1996), van Schaik and de Groot (1995), and Aricó (2003). In the Aghion-Howitt framework, for instance, an increase in the growth rate of productivity raises the present discounted value of the profits from opening a new job, which leads firms to open more vacancies and thus reduces unemployment. This is what they call a capitalization effect. At the same time, when productivity growth occurs through the creative destruction of low-productivity jobs and their replacement by new high-productivity ones elsewhere in the economy, then the inflow rate into unemployment will also be increased. This is termed the reallocation effect, which affects workers in the opposite direction of the capitalization effect. Aghion and Howitt show that the reallocation effect dominates at low growth rates, whereas the capitalization effect dominates at high ones, leading to a hump-shaped relation between growth and unemployment. Although the foregoing analysis is based on a causal effect from growth to unemployment (instead of the presumption that growth and unemployment are jointly determined, as emphasized by Bean and Pissarides, 1993), its main point is similar to the one made earlier: tradeoffs between unemployment and poverty reduction may emerge as a result of policy or structural shocks.

In practice, labor is heterogeneous, and households differ in terms of their sources of income. This implies that any analysis of unemployment must take into account its composition; similarly, changes in poverty rates must be examined not only at the aggregate level, but also at the level of various household groups. A policy-induced shock may entail a tradeoff solely between unemployment of one category of workers (say, unskilled workers) and one particular group of households (say, households in the urban informal sector). Consider, for instance, a reduction in the minimum wage, as discussed earlier; to the extent that the wage cut leads formal sector firms to substitute away from skilled labor (which has a higher degree of complementary with physical capital than with unskilled labor), skilled unemployment may increase at the same time that unskilled unemployment falls. In such conditions, the nature of the social welfare function becomes crucial in choosing a given policy path. The simulation framework presented below will help to illustrate partial tradeoffs of this nature.

2. Econometric Techniques

In this section, I use two alternative econometric techniques to assess empirically the importance of potential tradeoffs between unemployment and poverty. The first technique focuses on short-run dynamics; it is based on a vector autoregression (VAR) model involving a small set of stationary variables, which includes unemployment and poverty. The second involves cross-country regressions of poverty rates on a variety of structural and macroeconomic variables, with unemployment being among the explanatory variables.

2.1 A VAR Framework

A first approach to determining whether unemployment and poverty move in opposite directions in response to shocks in the short term is to specify a parsimonious VAR consisting of the detrended components of output, the open unemployment rate, real wages, and the poverty rate. These variables are chosen on the premise that in the short term an output shock, for instance, is transmitted to poverty primarily through two channels: either a change in unemployment or a change in real wages.[15] In general, the impact of a shock on poverty will depend on what group is hit the hardest by the rise in unemployment or the fall in real wages. If movements in these two variables primarily affect prime-age working males with low education, poverty may increase significantly. The VAR may thus need to include a measure of unemployment that closely reflects the labor market conditions faced by young or unskilled workers (as a proxy for vulnerable groups), and a real wage index that is representative of wages earned by the poor—say, an index of unskilled workers' wages or wages in the informal sector.

I applied this procedure to Brazil and Chile, using annual data in both cases. The estimation period is 1981–2002 for Brazil and 1981–2001 for Chile. In both countries, the issue of assessing the impact of macroeconomic variables on poverty has received significant attention.

15. As noted by Agénor (2002a), output shocks may also be accompanied by changes in intrafamily allocation of income or government transfers, which are not captured by movements in wages. Moreover, changes in open unemployment may not be highly correlated with output fluctuations, because adjustment to changes in labor demand takes the form of large movements in the labor force between the formal and informal sectors. Under such conditions, the open unemployment rate should be replaced by a measure of the size of the informal sector.

Paes de Barros, Corseuil, and Leite (2000), for instance, in a study based on microeconomic simulation techniques, find that unemployment has a major impact on the behavior of poverty rates in Brazil. However, none of the existing studies address the issue of potential tradeoffs between unemployment and poverty.

For both countries, I estimate the trend component of each variable by using a modified version of the ideal band pass filter of Baxter and King (1999), as proposed by Christiano and Fitzgerald (2003). The Baxter-King filter is a linear transformation of the data, which leaves intact the components within a specified band of frequencies and eliminates all other components. Its application requires a large amount of data, however. Christiano and Fitzgerald (2003) propose the following approximation. Let y_t be the data series that would result from applying the ideal band pass filter to the raw data, x_t. Then y_t is approximated by $\hat{y}_t$, which is a filter of x_t, with weights chosen to minimize the mean square error:

$$E\left[\left(y_t - \hat{y}_t\right)^2 \middle| x\right].$$

Specifically, $\hat{y}_t$ is computed as

$$\hat{y}_t = B_0 x_t + B_1 x_{t+1} + \ldots + B_{T-1-t} x_{T-1} + \tilde{B}_{T-t} x_T + B_1 x_{t-1} + \ldots + B_{t-2} x_2 + \tilde{B}$$

for t = 1, 2, 4,…, T, where

$$B_j = \frac{\sin(jb) - \sin(ja)}{\pi j},\ j \geq 1,$$

$$B_0 = \frac{b-a}{\pi},\ a = \frac{2\pi}{p_u},\ b = \frac{2\pi}{p_l},$$

and $\tilde{B}_{T-t}$ and $\tilde{B}_{t-1}$ are linear functions of the values of B_j

$$\tilde{B}_{T-t} = -\frac{1}{2} B_0 = \sum_{j=1}^{T-t-1} B_j\ ,$$

and $\tilde{B}_{t-1}$ solves

$$0 = B_0 + B_1 + \ldots + B_{T-t-1} + \tilde{B}_{T-t} + \ldots + B_{t-2} + \tilde{B}_{t-1}\ ,$$

with p_u = 24 and p_t = 2 in the present case.

Consider first the case of Brazil. The variables included in the VAR, which are defined more precisely in appendix A, are the (log of the) output gap and the cyclical components of the (log of the) aggregate unemployment rate, the real minimum wage, and the poverty gap, defined as the average shortfall of the income of the poor with respect to the national poverty line, multiplied by the headcount ratio (as defined earlier).[16] The real minimum wage plays a key role in the distribution of wages in Brazil (as noted, for instance, by Neri and Thomas, 2000), and it is a good proxy for the unskilled real wage, as time-series comparisons indicate that these two series are highly correlated.

Augmented Dickey-Fuller (ADF) stationary tests indicated that all the variables, as defined here, are stationary.[17] A standard VAR approach (that is, one that ignores cointegrating relationships between the variables in level form) can therefore be used.[18] Figure 4 shows the evolution of the cyclical components of all the variables included in the VAR. The data illustrate fairly well the pro-cyclical behavior of the real minimum wage and the counter-cyclical behavior of unemployment and poverty.

Variables in the VAR are ordered as follows: output gap—real minimum wage—unemployment rate—poverty rate. The fact that the output gap and the unemployment rate are placed before the poverty rate in the VAR captures the assumption that shocks to poverty have no contemporaneous impact on these variables. Any contemporaneous correlation between a disturbance to the poverty rate and the output gap, for instance, is thus taken to reflect causation

16. The poverty gap is defined as

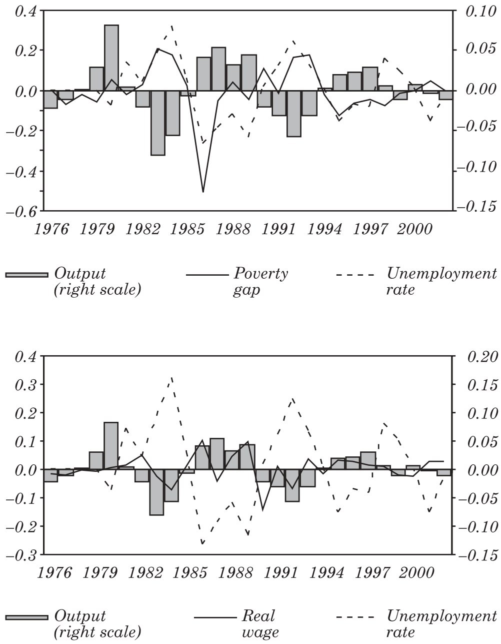
$$P_G = \left(ny^*\right)^{-1} \sum_{i \in L} \left(y^* - y_i\right),$$

where $y^* - y_i$ measures, for individual i in poverty, the gap between income, y_i, and the poverty line, y^*; L is the set of all poor; and n is the total number of poor.

17. The ADF test statistics were –3.418 for the cyclical component of the poverty rate (significant at the 5 percent level, using MacKinnon critical values for rejection of the null hypothesis), –2.978 for the detrended component of unemployment (significant at the 10 percent level), –3.889 for the cyclical component of the real minimum wage (significant at the 1 percent level), and –4.975 for the detrended component of output (significant at the 1 percent level).

18. Alternatively, all variables in the VAR could be measured in levels, despite being nonstationary. As shown by Sims, Stock, and Watson (1990), least-squares estimates are consistent for the levels specification (whether or not cointegration exists), whereas a differenced specification is inconsistent if some variables are cointegrated. In the absence of cointegration, the estimated standard errors of the levels specification are not consistent, so conventional inference could potentially be misleading.

Figure 4. Brazil: Cyclical Components of Real GDP, Unemployment Rate, Real Wage, and Poverty Gap, 1976-2002[a]



a. The cyclical component of each variable is defined as the log difference of the variable from its trend value calculated by using the Baxter-King filtering method.

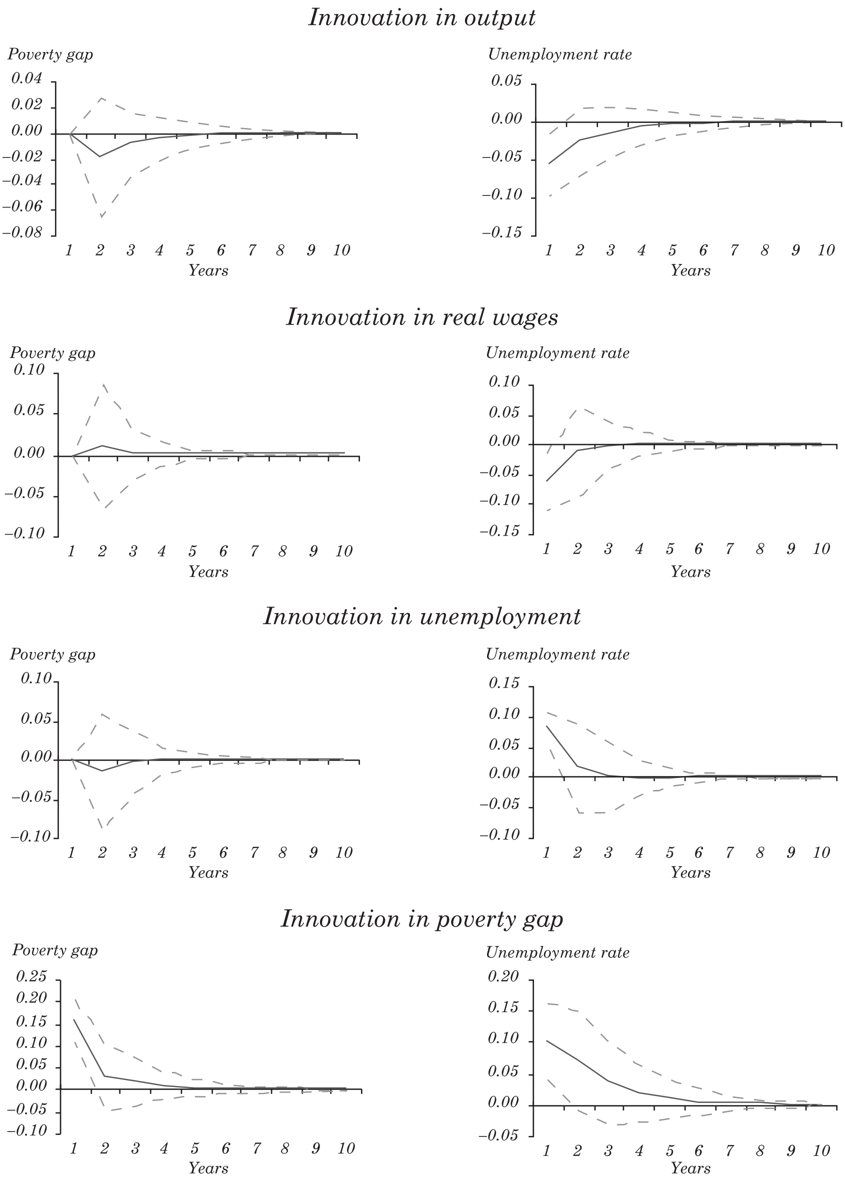
from output to poverty, and not the other way around.[19] I used the Akaike criterion to choose the optimal lag length. Given the relatively small size of the sample, I only compared models with one and two lags. The test led to the selection of one lag as the optimal choice.

The impulse response functions of the poverty gap and unemployment associated with a one-standard-deviation shock to the innovation in all the variables included in the VAR are shown in figure 5. The solid lines in the figures represent the impulse responses themselves, whereas the dotted lines are the associated 95 percent upper and lower confidence bands.[20] An innovation in output lowers unemployment (as

19. Alternative orderings were also considered, with either the poverty rate or the unemployment rate always appearing last in the sequence. The results were virtually unchanged.

20. The confidence intervals were generated with EViews, using a procedure based on analytical derivatives.

Figure 5. Brazil: Impulse Response Functions[a]



a. Cyclical components of each variable are used. The VAR model is estimated using a one-period lag. Each innovation corresponds to one standard deviation of the respective variable.

expected) but has no statistically significant effect on poverty. An innovation in real wages has, again, no effect on poverty and a perverse effect on unemployment in the first period. An innovation in the unemployment rate raises unemployment, of course, with no effect on poverty, whereas an innovation in the poverty gap has a positive and significant effect on both variables.

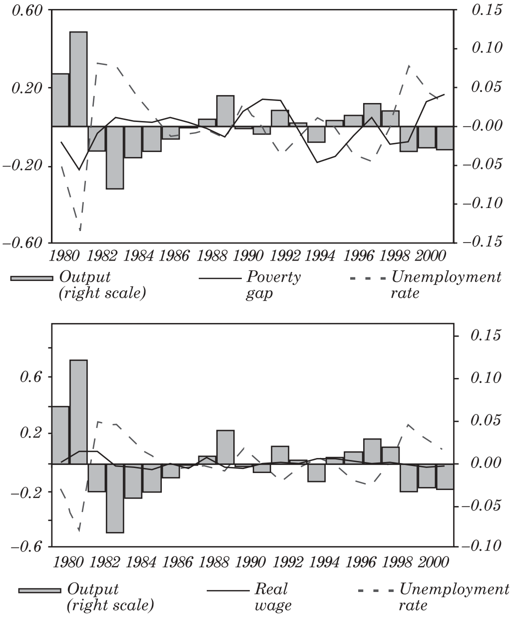
Consider now the case of Chile. The variables included in the VAR are the (log of the) output gap and the cyclical components of the (log of the) urban unemployment rate, the real wage for unskilled labor, and the headcount poverty index for the Santiago Metropolitan area.[21] ADF tests also indicated that all these variables are stationary.[22] Figure 6 displays the cyclical components of all the variables. Although real unskilled wages seem to fluctuate relatively little over time, they do show some degree of procyclicality. Both unemployment and poverty are countercyclical. Unemployment seems to fluctuate a lot more than poverty, however, and in the 1990s the two variables appear to be negatively correlated—an observation that would be consistent with a tradeoff between them, despite the fact that the sample period is small. Using the same ordering as before and uniformly selecting one lag (based on the Akaike criterion), I calculated the impulse response functions of the poverty gap and unemployment. The results, illustrated in figure 7, indicate that a positive innovation in output lowers unemployment and raises unskilled wages (again, as expected), but it has no direct, discernible effect on poverty. An innovation in real wages has no statistically significant effect on either one of the variables of the system. Unemployment shocks have no significant impact on poverty, and conversely poverty shocks do not affect unemployment.

Overall, therefore, the results for Brazil and Chile do not indicate the existence of a short-term tradeoff between poverty reduction and unemployment. This result may be due to a variety of factors, including limitations in the data. For instance, the aggregate unemployment rate was used in both cases, instead of the unskilled unemployment rate; the latter would be more appropriate given the correlation between education and poverty levels. More advanced approaches might

21. More precise definitions of these variables are provided in appendix A. The VAR model was also estimated with a measure of extreme poverty and with an index of average wages in the urban sector. In both cases, the impulse response functions obtained were very similar to those reported here.

22. The ADF test statistics were –4.479, –3.461, –3.022, and –3.064 for the detrended components of, respectively, the poverty rate, the unemployment rate, the real unskilled wage, and real GDP. All these statistics are significant at a 5 percent threshold or higher.

Figure 6. Chile: Cyclical Components of Real GDP, Unemployment Rate, Real Wage, and Poverty Gap, 1980-2001[a]



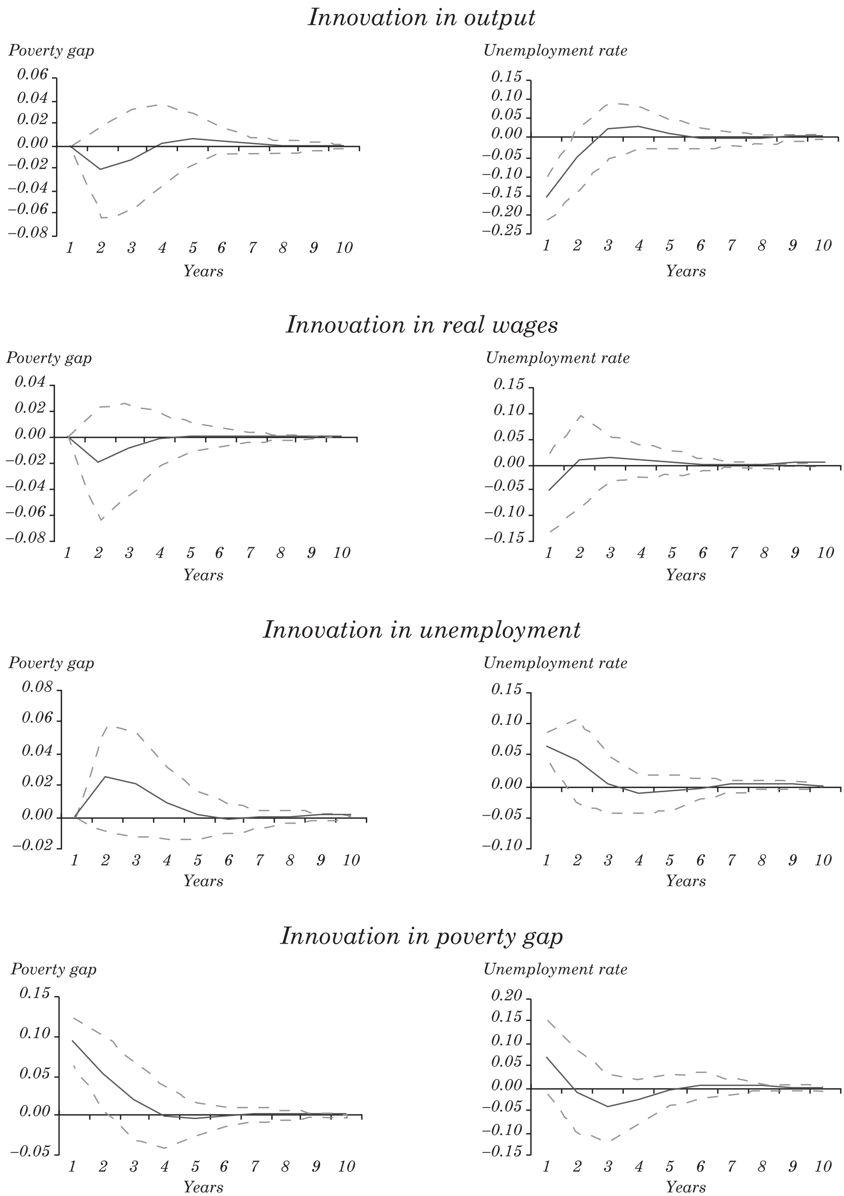
a. The cyclical component of each variable is defined as the log difference of the variable from its trend value calculated by using the Baxter-King filtering method.

also provide different results. One line of investigation would be to develop a structural VAR model, which would allow one to disentangle the importance of, say, real wage shocks, as opposed to, say, productivity shocks, in the behavior of poverty and unemployment. Alternatively, an error-correction framework would allow a possible distinction between short- and longer-term tradeoffs. This could be important because the fact that output shocks appear to have no effect on poverty in a VAR in which all variables are entered in detrended form does not preclude the existence of a cointegrating relationship between the raw output and poverty series themselves.

2.2 Cross-country Regressions

As noted earlier, if both unemployment and poverty are viewed as jointly endogenous, a key issue then becomes to identify the ultimate source of the differences in unemployment, growth, and poverty, either over time (at the level of an individual country) or across countries.

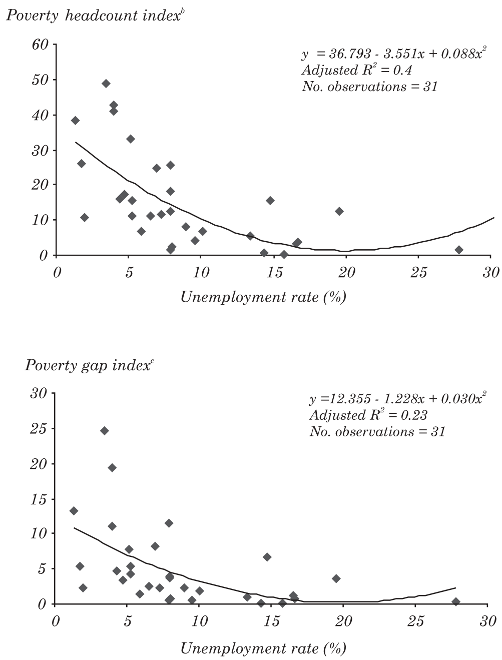
Figure 7. Chile: Impulse Response Functions[a]



a. Cyclical components of each variable are used. The VAR model is estimated using a one-period lag. Each innovation corresponds to one standard deviation of the respective variable.

Figure 8 displays data for a group of thirty-one developing countries on two standard measures of poverty (the headcount index and the poverty gap, both defined earlier) and the open unemployment rate. The number of countries corresponds to all those for which matching data were obtained between the World Bank and the ILO databases on these variables. Each data point is an average of all available observations for each country. The figure suggests a negative correlation (and thus a potential tradeoff) between poverty and unemployment across countries. Moreover, a simple cross-section regression of poverty on unemployment (also shown in the figure) suggests that the relation between these variables is convex: the correlation appears to turn positive beyond a rate of unemployment of about 20 percent—specifically, 3.551/(2*0.088) = 20.2 for the headcount index, and 1.228/(2*0.03) = 20.5 for the poverty gap.

Figure 8. Developing Countries: Unemployment and Poverty[a]



Source: World Bank Global Poverty Monitoring and ILO.

a. Samples consist of thirty-one countries for which data are available in the World Bank Global Poverty Monitoring (www.worldbank.org/research/povmonitor/) and ILO.

b. Proportion of the population earning 1.08 US dollar or less a day, various survey years.

c. Poverty gap at 1.08 US dollar or less a day, various survey years.

However, given the small number of data points in that range, it is difficult to draw much from this increasing portion of the curve, despite the statistical significance of the quadratic term in unemployment in the regression.

A simple explanation for the negative correlation between unemployment and poverty shown in the figure is that it reflects the fact that poor countries often have a large informal sector, such that open (or officially measured) unemployment tends to be small. At the same time, the urban poor tend to be highly concentrated in the informal sector. Thus, a large informal sector corresponds to a lower open unemployment rate (reflecting higher disguised unemployment) and a higher poverty rate. There are two problems, however, with this interpretation. First, it does not appear to hold in some regions. In the Middle East and North Africa, most notably, a good part of unemployment is voluntary in nature and affects the educated. The link between unemployment and poverty thus tends to be weak: the World Bank (2004) finds, using microeconomic data, that poverty and labor market status are not closely correlated in that region. Second, this interpretation does not appear to be sufficient; in the cross-country econometric results discussed later, I control indirectly for the size of the informal sector by using income per capita as a regressor (the lower the standards of living, the larger is the size of the informal economy), and the negative correlation between unemployment and poverty persists.

To assess the relation between these two variables over time and across countries, I specify and estimate a cross-country regression model, using unbalanced panel data for a group of developing economies. The dependent variable is either the headcount index or the poverty gap, based on the international poverty line of $1.08 a day. Based on my previous results (see Agénor, 2002a, 2002b, 2004a), I included the following explanatory variables in the regressions, in addition to the unemployment rate: INFL is the inflation rate in terms of consumer prices; LGDPPC is the log of gross domestic product (GDP) per capita at purchasing power parity (PPP) exchange rates, which captures the level of economic development and the effect of economic growth on standards of living; REALEX is the rate of change of the real effective exchange rate (defined such that an increase is a depreciation); VREALXL is a measure of macroeconomic volatility, which consists of rolling standard deviations of the real exchange rate; and TARIFF is the average tariff rate (total tariff revenue divided by the value of imports).[23]

23. See appendix A for more precise definitions and sources.

I have discussed at length elsewhere the rationale for considering these variables (see Agénor, 2002a, 2002b, 2004a), so only a brief justification is offered here. Inflation (which is a tax on nonindexed financial assets, such as currency holdings) lowers the overall purchasing power of households and tends to raise poverty. An increase in real GDP per capita is expected to be negatively correlated with the poverty rate. The effect of a real exchange rate depreciation is ambiguous. It may lead to a reduction in poverty if it benefits small farmers in the tradable sector (as is the case in many low-income developing countries); but overall poverty may increase if the depreciation is accompanied by a significant increase in the cost-of-living index in urban areas (as a result of a rise in the domestic price of imported goods). The average tariff rate is a proxy for the degree of trade openness, or "real" globalization, and is expected to have a nonlinear effect on poverty (see Agénor, 2004a): to the extent that trade liberalization entails short-run adjustment costs (as a result of a reduction in employment in import-substitution industries, for instance), poverty may rise initially; over time, as liberalization progresses and tariffs continue to fall, the expansion of employment in export industries may lead to lower poverty. This is tested by using both the average tariff rate and its squared value as regressors. The tariff rate itself is expected to have a negative effect on poverty, whereas its squared value is expected to have a positive effect.

The data on poverty rates are taken from the World Bank and cover countries for which data on the unemployment rate are simultaneously available from the ILO, with at least two observations available for each country. These requirements give a relatively small sample, consisting of eleven countries and forty observations (see appendix A). The first estimation method that I use is ordinary least squares (OLS) with fixed effects. The results are reported in table 1, columns 1 and 2 for the headcount index and columns 4 and 5 for the poverty gap. The difference between columns 1 and 2 and columns 4 and 5 is that the change in the real exchange rate, and the volatility measure based on it, are entered separately, because of colinearity between the variables. The results are very similar, however. Inflation raises poverty, whereas higher income per capita tends to reduce it. A real exchange rate depreciation and a high degree of real exchange rate volatility both tend to increase poverty. The tariff rate and its squared values have the expected sign—increased trade openness (a reduction in tariffs) tends to increase poverty at first and then reduces it beyond a certain threshold, a result consistent with the globalization-poverty curve discussed by Agénor (2004a) in a more general setting. The open unemployment rate also appears to have a nonmonotonic effect on poverty: lower unemployment is associated with higher poverty, but an opposite effect kicks in at levels of unemployment of 0.033/(2*0.002) = 8.3 percent for the headcount index and 0.01/(2*0.001) = 5 percent for the poverty gap (regressions 1 and 4). These results corroborate at much smaller levels those shown in figure 8, which are based on a simple cross-section regression.[24] But again, caution is needed in interpreting the positive segment of the curve, given the small number of data points in that range.

To account for possible simultaneity problems with the control variables, I also used an instrumental variables procedure (together with fixed effects). In the first step, inflation, unemployment, income per capita, and the rate of depreciation of the real exchange rate (or the index of volatility based on it) were all regressed on the lagged values of each variable at $t-1$, $t-2$, and $t-3$, as well as on the tariff rate and its squared value. In the second step, the predicted values from these regressions were introduced in the poverty regression, together with linear and quadratic terms in the tariff rate. The estimation results are shown in columns 3 and 6 for the two measures of poverty and the percentage change in the real exchange rate. By and large, the estimates obtained with OLS are unaffected, except that the real exchange rate variable loses some of its significance. Most importantly for the issue at hand, the degree of significance of the coefficients on the unemployment rate and its squared value, as well as their size, increases. This implies slightly higher threshold levels for the unemployment rate to be positively correlated with poverty: 0.091/(2*0.05) = 9.1 percent for the headcount index and 0.028/(2*0.002) = 7.0 percent for the output gap.

Finally, I reran all the regressions using the employment ratio (as measured by the share of employment in total population) instead of the open unemployment rate, on the grounds that employment and total population are measured with a greater degree of precision than the labor force—perhaps because of the difficulty of accurately measuring changes in participation rates. The results, shown in table 2, are very similar to those reported in table 1, except for the fact that the coefficients on the linear and quadratic terms in the employment ratio have the opposite sign (as expected), and the quadratic term in the employment ratio, when the poverty gap is used and the instrumental variables methodology is applied, is only borderline significant.

24. The difference is that the cross-section regression attempts to explain the cross-country variation in poverty rates on the basis of the independent variables only, whereas the panel regressions explain some of the variation through separate intercepts (or fixed effects). The coefficient of the quadratic term in the panel regressions is determined with greater precision than in the cross-section regression, owing to the larger number of observations.

Table 1. Unemployment Rate and Poverty in Developing Countries, 1981–98[a]

	Headcount poverty index			*Poverty gap*		
Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
UNEMP	–0.033	–0.029		–0.010	–0.058	
	(–1.895)	(–2.048)		(–2.153)	(–2.034)	
UNEMP_SQ	0.002	0.001		0.001	0.001	
	(2.002)	(1.919)		(2.199)	(1.852)	
IVUNEMP			–0.091			–0.028
			(–3.497)			(–2.491)
IVUNEMP_SQ			0.005			0.002
			(2.409)			(2.200)
INFL	0.007	0.006		0.003	0.003	
	(4.132)	(4.115)		(7.410)	(6.842)	
IVINFL			0.047			0.021
			(15.192)			(9.318)
REALEX	0.087			0.021		
	(2.139)			(1.759)		
IVREALEX			0.373			0.129
			(1.472)			(1.341)
VREALXL		0.198			0.034	
		(2.103)			(1.249)	
LGDPPC	–0.176	–0.200		–0.053	–0.056	
	(–9.802)	(–3.299)		(–8.241)	(–3.310)	
IVLGDPPC			–0.188			–0.047
			(–5.374)			(–3.309)
TARIFF	–1.282	–1.420	–2.049	–0.606	–0.652	–0.902
	(–3.074)	(–3.679)	(–5.712)	(–4.027)	(–4.437)	(–6.106)
TARIFF_SQ	3.202	3.361	5.503	1.860	1.973	2.929
	(2.764)	(3.156)	(5.629)	(4.304)	(4.921)	(7.044)
Summary statistic						
Adjusted R^2	0.816	0.807	0.817	0.762	0.743	0.772
Total panel observations	40	38	38	40	38	38
Standard error of regression	0.041	0.042	0.041	0.014	0.014	0.014

a. The dependent variable is the headcount poverty index in regressions 1 through 3 (that is, the ratio of population earning less than USD 1.08 per day) and the poverty gap in regressions 4 through 6 (that is, the mean shortfall from the poverty line of USD 1.08 per day, expressed as a percentage of the poverty line). The estimation technique is ordinary least squares with fixed effects in regressions 1, 2, 4, and 5 and two-stage least squares with fixed effects in regressions 3 and 6. UNEMP is the rate of unemployment; UNEMP_SQ is its squared value. IVUNEMP is the instrumental variable of UNEMP (fitted values obtained by regressing UNEMP on the growth rate of GDP per capita (purchasing power parity) at $t-1$, $t-2$, and $t-3$, TARIFF, and TARIFF_SQ). IVUNEMP_SQ is the squared value of IVUNEMP. INFL is the annual change in the consumer price index. IVINFL is the instrumental variable of INFL (fitted values obtained by regressing INFL on INFL at $t-1$, $t-2$, and $t-3$, TARIFF, and TARIFF_SQ). REALEX is the annual change in the real effective exchange rate index (a rise is a depreciation). IVREALEX is the instrumental variable of REALEX (fitted values obtained by regressing REALEX on REALEX at $t-1$, $t-2$, and $t-3$, TARIFF, and TARIFF_SQ). VREALXL is the volatility measure of the real effective exchange rate, calculated as the ratio of the standard deviation of the variable for t, $t-1$, $t-2$, and $t-3$ to the average value for the same period. LGDPPC is the log of the GDP per capita (purchasing power parity). IVLGDPPC is the instrumental variable of LGDPPC (fitted values obtained by regressing LGDPPC on LGDPPC at t-1, t-2 and t-3, TARIFF, and TARIFF_SQ). TARIFF is the average tariff rate and TARIFF_SQ is its squared value; t statistics are in parentheses.

Table 2. Employment Ratio and Poverty in Developing Countries, 1981–98[a]

	Headcount poverty index			*Poverty gap*		
Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
EMP	8.513	9.036		2.661	2.310	
	(3.325)	(3.539)		(2.850)	(2.504)	
EMP_SQ	–8.438	–9.567		–2.388	–2.124	
	(–3.169)	(–3.411)		(–2.451)	(–2.156)	
IVEMP			6.393			1.414
			(6.546)			(2.206)
IVEMP_SQ			–6.295			–1.036
			(–5.628)			(–1.572)
INFL	0.005	0.005		0.003	0.002	
	(2.431)	(2.596)		(4.812)	(5.032)	
IVINFL			0.033			0.013
			(2.008)			(1.523)
REALEX	0.088			0.023		
	(3.014)			(2.642)		
IVREALEX			0.241			0.201
			(0.982)			(1.856)
VREALXL		0.218			0.038	
		(2.914)			(1.703)	
LGDPPC	–0.165	–0.172		–0.060	–0.063	
	(–3.208)	(–3.519)		(–3.370)	(–3.092)	
IVLGDPPC			–0.185			–0.073
			(–3.697)			(–3.184)
TARIFF	–0.763	–0.948	–1.291	–0.411	–0.497	–0.392
	(–1.628)	(–2.255)	(–2.222)	(–2.508)	(–3.385)	(–1.687)
TARIFF_SQ	3.011	3.236	4.720	1.703	1.856	1.635
	(2.756)	(3.274)	(2.857)	(4.254)	(5.033)	(2.430)
Summary statistic						
Adjusted R^2	0.850	0.839	0.889	0.788	0.755	0.788
Total panel observations	40	38	31	40	38	31
Standard error of regression	0.037	0.038	0.033	0.013	0.014	0.013

a. The dependent variable is the headcount poverty index in regressions 1 through 3 (that is, the ratio of population earning less than USD 1.08 per day) and the poverty gap in regressions 4 through 6 (that is, the mean shortfall from the poverty line of USD 1.08 per day, expressed as a percentage of the poverty line). The estimation technique is ordinary least squares with fixed effects in regressions 1, 2, 4, and 5 and two-stage least squares with fixed effects in regressions 3 and 6. EMP is the ratio of employment to total population; EMP_SQ is its squared value. IVEMP is the instrumental variable of EMP (fitted values obtained by regressing EMP on the lagged values of EMP at $t-1$, $t-2$, and $t-3$, TARIFF, and TARIFF_SQ). IVEMP_SQ is the squared value of IVEMP. INFL is the annual change in the consumer price index. IVINFL is the instrumental variable of INFL (fitted values obtained by regressing INFL on INFL at $t-1$, $t-2$, and $t-3$, TARIFF, and TARIFF_SQ). REALEX the annual change in the real effective exchange rate index (a rise is a depreciation). IVREALEX is the instrumental variable of REALEX (fitted values obtained by regressing REALEX on REALEX at $t-1$, $t-2$, and $t-3$, TARIFF, and TARIFF_SQ). VREALXL is the volatility measure of the real effective exchange rate, calculated as the ratio of the standard deviation of the variable for t, $t-1$, $t-2$, and $t-3$ to the average value for the same period. LGDPPC is the log of the GDP per capita (purchasing power parity). IVLGDPPC is the instrumental variable of LGDPPC (fitted values obtained by regressing LGDPPC on LGDPPC at $t-1$, $t-2$, and $t-3$, TARIFF, and TARIFF_SQ). TARIFF is the average tariff rate and TARIFF_SQ is its squared value; t statistics are in parentheses.

Overall, therefore, the results suggest that when unemployment is below a threshold of about 10 percent, a tradeoff seems to exist between poverty and unemployment across countries. The next step would be to determine the exact source of this tradeoff—for instance, changes in labor market regulations during the sample period, as suggested by the model of Bean and Pissarides (1993) discussed earlier. This could be done by estimating a simultaneous equations system in unemployment and poverty rates, with the explicit introduction of an index of labor market regulations and other variables likely to affect unemployment (such as the presence of a binding minimum wage or a compensation scheme for the unemployed).

3. A Structural Approach

Another approach that can be used to gauge the extent to which poverty-unemployment tradeoffs are important, depending on the origin of shocks, is to use a numerical model and perform relevant simulations. I do so here with the Mini Integrated Macroeconomic Model For Poverty Analysis (mini-IMMPA) model, which was developed at the World Bank to quantify poverty reduction strategies in developing countries.[25] An appealing feature of the model is its detailed treatment of the labor market and the sources of unemployment in a typical developing-country context. I first describe the macroeconomic component of the model, emphasizing the production side and the structure of the labor market, and explain briefly how it is linked to a household survey for poverty analysis. Other features of the model (such as the composition of aggregate demand, the determination of prices, and the distribution of income flows) are briefly summarized in appendix B. I then report simulation results associated with two types of labor market reforms: a cut in the minimum wage and a reduction in payroll taxes on unskilled labor in the formal sector.

3.1 Production and the Labor Market

The structure of production and the labor market in Mini-IMMPA are summarized in figure 9. Production activities take place in both

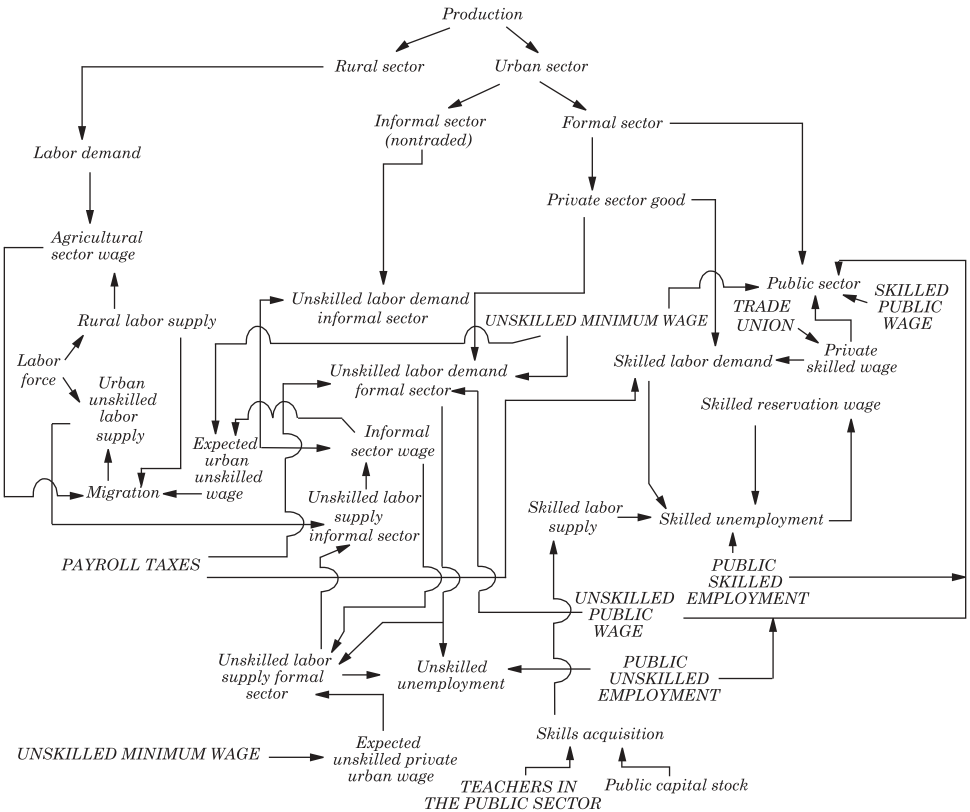
25. See Agénor (2003), Agénor and El Aynaoui (2003), Agénor, Izquierdo, and Fofack (2003), Agénor, Fernandes, and Haddad (2003), and Agénor and others (2004).

rural and urban areas. The rural sector produces only one good, which is sold either on domestic markets or abroad. Urban production includes both formal and informal components; in addition, the urban formal economy is separated between production of private and public goods. Gross output of each type of good is given by the sum of value-added and intermediate consumption. Value-added in the rural sector is assumed to be produced with land (which is in fixed supply) and a composite factor, which consists of unskilled labor and public capital. Value added in the urban informal sector depends only on labor and is subject to decreasing returns to scale. Value-added in the public sector is measured by the government wage bill, and employment is exogenous. Private formal production uses as inputs both skilled and unskilled labor, as well as public and private capital. Skilled labor and private physical capital have a higher degree of complementarity (lower degree of substitution) than the physical capital-skilled labor bundle with unskilled labor. Firms in the urban formal sector are subject to a payroll tax on unskilled labor.

Unskilled workers are employed in both the rural and urban sectors, whereas skilled workers are employed only in the urban formal economy. Wages in the rural and urban informal sectors adjust to equilibrate supply and demand. Unskilled workers in the urban economy may be employed in either the formal sector, in which case they are paid the minimum wage, or the informal economy, where they receive the going wage. The nominal wage for skilled labor in the private sector is determined on the basis of a monopoly union approach, as in Agénor (2005). The consumption real wage is set by a representative labor union, whose objective is to maximize a utility function that depends on deviations of both employment and the consumption wage from their target levels, subject to the firm's labor demand schedule. The union's target wage is related negatively to the skilled unemployment rate. Education is a pure public good; the flow of unskilled workers who become skilled is a function of the effective number of teachers in the public sector and the stock of public capital in education.

Incentives to rural-urban migration depend on the differential between expected rural and urban wages, as in Harris and Todaro (1970). The expected (unskilled) urban wage is a weighted average of the minimum wage in the formal sector and the going wage in the informal sector. The degree of mobility of the unskilled labor force between the formal and informal sectors is also imperfect and is a function of expected income opportunities. The supply of labor in the

Figure 9. Production Structure and the Labor Market



Source: Agénor (2003).

informal economy is obtained by subtracting the number of unskilled job seekers in the urban formal sector from the urban unskilled labor force, which increases as a result of natural urban population growth and migration from the rural economy and which falls because some unskilled workers acquire skills and leave the unskilled labor force to increase the supply of skilled labor.

3.2 Link with a Household Survey

The procedure followed here to assess the poverty effects of policy shocks involves linking the structural macroeconomic component described earlier to a household income and expenditure survey, in order to calculate both the headcount index and the poverty gap. This procedure, which is discussed at length in Agénor, Chen, and Grimm (2004) and Agénor, Izquierdo, and Fofack (2003), involves six steps. The first step is to classify the data in the household survey into the five categories of households contained in the macroeconomic framework—workers in the rural sector, unskilled workers in the urban informal sector, unskilled workers in the urban formal sector, skilled workers in the urban formal sector, and profit earners (see appendix B).

The second step is to initiate a shock and then generate the subsequent real growth rates in real per capita consumption and disposable income for all categories of households, up to the end of the simulation horizon. The third step is to apply these growth rates separately to the per capita (disposable) income and consumption expenditure for each household in the survey. This gives a new vector of absolute income and consumption levels for each individual in each group.

The fourth step is to update the initial rural and urban poverty lines to reflect increases in rural and urban price indexes and then to calculate poverty indicators using the new vector of absolute levels of income and consumption. The fifth step uses employment and unemployment growth rates to adjust the composition of the sample of each household group, as given in the survey. Finally, the sixth step is to compare the post-shock poverty indicators with the baseline values to assess the impact of the shock on the poor.

3.3 Policy Shocks

This subsection examines the poverty and employment effects of two types of labor market reforms: a cut in the minimum wage and a reduction in the payroll tax rate on unskilled labor paid by firms in the private formal sector. Discussions of the employment effects of changes in minimum wages and the taxation of labor figure prominently in the recent debate on labor market reforms in developing countries (see, for instance, Agénor, 2004b; World Bank, 2004). Assessing whether these policies may entail tradeoffs between unemployment reduction and poverty alleviation is thus timely. In both simulations, I consider only permanent shocks and focus on the first ten periods after the shock. For the payroll tax experiment, I consider three alternative budget financing rules: domestic borrowing with no offsetting tax change; and offsetting, revenue-neutral increases in either sales taxes on private formal sector goods or income taxes on profit earners.[26] In all of these experiments, the government borrows domestically to finance its deficit, and private capital formation is determined residually to maintain a continuous equilibrium between aggregate savings and investment.[27]

Reduction in the Minimum Wage

Simulation results associated with a 10 percent reduction in the minimum wage are shown in table 3, which displays absolute percentage changes from the baseline solution of unemployment (both skilled and unskilled) and poverty rates (for informal sector households, formal unskilled households, and formal skilled households), as measured by the poverty gap.

The impact (or first year) effect of the reduction in the minimum wage is an increase in the demand for unskilled labor in the private sector on the order of 4.3 percent. The increase in demand is met by the existing pool of unskilled workers seeking employment in the urban sector. As a result, the unskilled unemployment rate drops significantly, by 2.9 percentage points in the first year. The cut in the minimum wage reduces the relative cost of unskilled labor, which leads to substitution among production factors not only on impact but also over time. Because unskilled labor has a relatively high elasticity of substitution with respect to the composite factor consisting of skilled labor and physical capital, the reduced cost of that category of labor gives private firms in the formal sector a fairly strong incentive

26. The calibration procedure and parameter values used in these simulations are described in appendix C. Detailed tables summarizing the simulation results are available on request.

27. How this transfer of private savings to the government takes place is not explicitly specified; one can think of a pure financial intermediary operating in the background.

to substitute away from skilled labor and physical capital. The fall in the demand for that category of labor puts downward pressure on skilled wages, which drop by 1.6 percent in the first period. On impact, labor supply is fixed in the rural and informal sectors, so the level of employment does not change in either sector—and neither does the level of activity (that is, real value-added in both sectors is constant). The rise in real disposable income and real consumption of rural and informal sector households leads to higher value-added prices and higher wages in both sectors. Value-added prices go up by slightly more than wages in the second and subsequent periods, implying a fall in the product wage in both sectors and a rise in employment.

Over time, changes in wage differentials affect both rural-urban and formal-informal migration flows and, therefore, the supply of labor in the various production sectors. The expected unskilled wage in the formal economy is constant on impact. Despite the increase in unskilled employment in the private sector in the first period (implying a higher perceived probability of finding a job in that sector), the fall in the minimum wage is large enough to entail a reduction in the urban expected wage. At the same time, rural sector wages rise, thereby magnifying the fall in the expected urban-rural wage differential. In the second period, the drop in this differential (measured as a proportion of the rural wage) is 8.7 percentage points; it persists over time, although it narrows somewhat. As a result, the inflow of unskilled workers in the informal sector (measured as a proportion of the total supply of unskilled labor in the urban sector) falls by about 1.2 percentage points in periods 2 and 3. The reduction in the labor supply leads, in turn, to an increase in informal sector wages throughout the adjustment period. This increase in the informal sector wage, coupled with the reduction in the minimum wage (as well as the expected wage in the private urban formal sector, despite the higher employment probability), leads to a sharp drop in period 2 in the expected formal-informal wage differential. As a result, the number of unskilled workers willing to queue for employment in the formal private sector falls. The reduction in the number of job seekers, together with the sustained effect of the cut in the minimum wage on labor demand, explains the large impact on unemployment, which averages about 11 percent in the long run.[28]

28. Unskilled employment in the formal (private) sector increases by about 10 percent in the long run, whereas the number of unskilled job seekers in the formal economy drops by 4.5 percent.

Table 3. Simulation Results

Absolute deviations from baseline, unless otherwise indicated

Simulated policy, affected variable, and sector[a]	*Period* (1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Unskilled labor minimum wage										
Unemployment rate (urban formal sector)										
Unskilled	–2.90	–11.63	–9.63	–10.52	–10.63	–10.89	–11.08	–11.25	–11.39	–11.51
Skilled	0.29	0.14	0.19	0.19	0.19	0.20	0.20	0.20	0.21	0.21
Poverty gap (urban)										
Informal	–1.52	–0.58	–0.96	–0.97	–1.06	–1.12	–1.17	–1.22	–1.26	–1.30
Formal unskilled	1.01	–0.90	–0.51	–0.69	–0.72	–0.78	–0.82	–0.86	–0.89	–0.90
Formal skilled	0.31	0.17	0.22	0.21	0.22	0.22	0.22	0.21	0.22	0.20
Unskilled labor payroll tax rate—nonneutral										
Unemployment rate (urban formal sector)										
Unskilled	–0.90	–5.21	–3.79	–3.86	–3.55	–3.34	–3.10	–2.88	–2.66	–2.46
Skilled	0.31	0.23	0.25	0.24	0.25	0.25	0.25	0.25	0.25	0.25
Poverty gap (urban)										
Informal	–1.55	–0.92	–1.03	–0.94	–0.90	–0.85	–0.80	–0.75	–0.71	–0.66
Formal unskilled	0.06	–0.91	–0.60	–0.61	–0.53	–0.48	–0.42	–0.36	–0.30	–0.25
Formal skilled	0.33	0.25	0.28	0.27	0.27	0.26	0.25	0.25	0.25	0.26
Unskilled labor payroll tax rate—sales tax neutral										
Unemployment rate (urban formal sector)										
Unskilled	–2.07	–2.05	–1.88	–1.71	–1.54	–1.37	–1.21	–1.05	–0.90	–0.76
Skilled	0.25	0.25	0.25	0.25	0.26	0.26	0.26	0.26	0.26	0.26

Table 3. (continued)

Simulated policy, affected variable, and sector[a]	*Period*									
	(1)	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*	*(7)*	*(8)*	*(9)*	*(10)*
Unskilled labor payroll tax rate—sales tax neutral										
Poverty gap (urban)	−0.49	−0.44	−0.40	−0.37	−0.33	−0.29	−0.25	−0.22	−0.19	−0.15
Informal	−0.31	−0.30	−0.25	−0.20	−0.15	−0.11	−0.06	−0.02	0.02	0.06
Formal unskilled	0.24	0.24	0.25	0.25	0.25	0.24	0.24	0.24	0.25	0.26
Formal skilled										
Unskilled labor payroll tax rate—income tax neutral										
Unemployment rate (urban formal sector)										
Unskilled	−3.56	−3.49	−3.40	−3.30	−3.19	−3.09	−2.98	−2.87	−2.76	−2.66
Skilled	0.18	0.19	0.19	0.19	0.19	0.20	0.20	0.20	0.20	0.20
Poverty gap (urban)										
Informal	−1.05	−1.03	−1.00	−0.98	−0.95	−0.93	−0.90	−0.87	−0.85	−0.82
Formal unskilled	−0.35	−0.34	−0.32	−0.30	−0.27	−0.24	−0.22	−0.19	−0.17	−0.14
Formal skilled	0.44	0.46	0.48	0.49	0.49	0.48	0.48	0.49	0.52	0.54

a. All simulations involve a 10 percentage point cut in the indicated policy.

Although the behavior of nominal wages in the rural sector essentially reflects changes in value-added prices on impact, over time it is also affected by changes in output (induced by changes in households' disposable income and expenditure) and migration flows. After an initial increase in nominal wages, lower migration flows to urban areas begin to put downward pressure on rural wages, which end up falling (in nominal terms) by slightly less than 2 percent after ten years. As indicated earlier, the reduction in the cost of unskilled labor induces a substitution away from skilled labor, which brings a sustained fall in skilled wages in nominal terms. The overall effect on labor demand is not large; skilled employment in the private formal sector falls in the long run only slightly. Because the supply of skilled labor remains roughly constant throughout (public investment in education and the number of school teachers are held constant at their baseline values), the increase in the skilled unemployment rate (of about 0.3 percentage points in the long run) mirrors the drop in employment. The reason for the small effect on skilled employment is that the direct substitution effect associated with the reduction in the minimum wage is mitigated by a fall in the skilled wage, resulting from general equilibrium effects. The drop in the nominal skilled wage is larger than the fall in the value-added price of the private urban formal sector, implying a drop in the product wage. This, in turn, stimulates the demand for that category of labor.

Changes in real consumption and disposable income lead to significant differences in poverty patterns among urban households. As shown in table 3, poverty drops by 1.5 percentage points for informal sector households on impact, but it increases for both categories of workers in the formal sector (by 1.0 and 0.3 percentage points, respectively, for skilled and unskilled households). In the long run, poverty falls for unskilled workers in both the informal and formal sectors, whereas the slight increase in poverty recorded on impact for skilled workers persists. For that group of workers, the behavior of poverty tends to mirror the behavior of unemployment. Thus, the simulation results suggest the existence of a potential short-run tradeoff between unemployment and poverty: although the reduction in the minimum wage raises unskilled employment in the formal sector, it also increases poverty for households employed in that sector. Moreover, a longer-run tradeoff could potentially result from the fact that poverty among skilled workers increases (albeit slightly) in both the short and long term.

Cut in Payroll Tax on Unskilled Labor

Simulation results associated with a 10 percentage point reduction in the payroll tax rate on unskilled labor are also shown in table 3. The results correspond, as noted earlier, to three alternative budget financing rules: a nonneutral change involving domestic borrowing with no offsetting tax change; a revenue-neutral change on impact involving an increase in sales taxes on private formal sector goods; and a revenue-neutral change implying an offsetting increase in income taxes on profit earners.

Consider first the nonneutral experiment. The impact effect of a reduction in the payroll tax rate is qualitatively similar to a cut in the minimum wage, as discussed earlier: by reducing the effective cost of unskilled labor, it tends to increase immediately the demand for that category of labor. The unskilled unemployment rate drops by 0.9 percentage point in the first year and by an average of 2.5 percentage points in the long run. The reduction in the effective cost of unskilled labor also leads firms in the private urban formal sector to substitute away from skilled labor and physical capital, causing skilled employment to fall by about the same amount as in the previous experiment. The behavior of the (expected) urban-rural wage differential follows a pattern qualitatively similar to the one described in the previous experiment, although the magnitude of the initial effects are not as large. Most importantly, the expected formal-informal wage differential now increases in the second period. The reason is that the minimum wage does not change this time around, and the expansion in unskilled employment in the private formal sector raises the probability of finding a job there, thereby increasing the expected formal sector wage. The number of unskilled job seekers in the formal economy therefore increases, which explains why the reduction in the unemployment rate is significantly lower than in the previous case.[29] Changes in poverty among urban household groups follows a similar pattern as before. The long-run reduction in poverty in the informal sector is less marked, however, largely because wages do not increase by the same amount—the fall in open unskilled unemployment is less dramatic, so fewer workers seek employment in the formal sector. The impact effect on poverty among formal unskilled

29. This time, unskilled employment in the formal (private) sector increases by about 7.5 percent in the long run, but the number of unskilled job seekers in the formal economy increases as well, by 2.8 percent.

households is about the same, so the same type of tradeoffs identified earlier emerge.

Consider now the case in which the effect of the payroll tax cut on overall tax revenue is offset by either an increase in sales taxes on private formal sector goods or an increase in taxation of profit earners. In both cases, the impact and longer-run effects of the shock are qualitatively similar to those described earlier, although their magnitude differs. In particular, movements in the informal sector wage are less pronounced, in part because changes in rural-urban migration flows are not as large. The most important difference is that when the cut in payroll taxes is financed by an initial increase in income taxes, the fall in unskilled unemployment is larger in both the short and the long term, because the reduction in the after-tax rate of return on investment lowers the demand for physical capital, which has a high degree of substitution with unskilled labor. The reduction in poverty among informal and formal unskilled households and the increase in poverty among skilled households are also both larger on impact. Moreover, the long-run impact on poverty among formal unskilled households is negligible with an increase in sales taxes, whereas the long-run effects remain quite significant (and are even stronger for skilled households) with an increase in income taxes. As in the nonneutral experiment, poverty increases among skilled households but falls among unskilled households (both formal and informal). Unemployment among skilled workers rises at the same time that it falls among the unskilled.

Overall, the results indicate that there may be short- and longer-term tradeoffs between unemployment reduction and poverty alleviation among household groups. The magnitude of these tradeoffs depends on the nature of the financing rule that accompanies the shocks. While the results are specific to the policy shocks examined here (as well as to the nature of the model and the parameter values chosen for its calibration), one may surmise that these tradeoffs are more than mere curiosities and may well occur with other types of policy changes.

4. Concluding Remarks

The purpose of this paper has been to discuss analytically and assess empirically the potential short- and long-term tradeoffs that may arise between reducing poverty and lowering unemployment in developing countries. The first part provided a general discussion of the channels through which such tradeoffs may arise. The discussion noted that the expansion in employment (resulting from either favorable productivity shocks or lower wages) may be skewed toward low-paying jobs, and as long as the labor supply does not increase significantly, the increase in the numbers of working poor may translate into both lower unemployment and higher poverty. Furthermore, poverty and unemployment are both endogenous variables, and the correlation between them may depend on the type of shocks affecting the economy, either over time or across countries. This general proposition was illustrated in a growth context using a simple overlapping-generations model based on Bean and Pissarides (1993). In the model, unemployment is created by matching frictions in the labor market. The analysis showed that an increase in workers' bargaining power leads to higher wages, which discourages firms from opening new vacancies. This tends to raise unemployment. At the same time, a higher income for workers increases savings, which can stimulate growth and reduce poverty (assuming that growth is distribution neutral). The net effect on the pool of savings cannot be determined a priori—and thus neither can the effect on growth and unemployment. Nevertheless, the model can generate an inverse correlation between unemployment and poverty as a result of this type of shock.

The second part of the paper used two econometric techniques to assess empirically the relevance of these tradeoffs: a VAR framework and cross-country regressions. Impulse response functions derived from VAR models estimated for Brazil and Chile showed no short-run tradeoff between these variables, for neither output nor wage shocks. However, improvements in the quality of the data used, and the application of more sophisticated forms of VAR models, could deliver different results. The regression results, by contrast, show a negative relation between unemployment and poverty (as long as unemployment is below a certain threshold), even after controlling for various other determinants of poverty (such as inflation, real income per capita, changes in the real exchange rate, macroeconomic volatility, and the degree of trade openness), and using different econometric estimation techniques (specifically, OLS and instrumental variables with fixed effects).

The third part used a structural macroeconomic model built specifically for labor market and poverty analysis—namely, the Mini-IMMPA framework developed by Agénor (2003). Simulation results showed that labor market reforms can induce both short- and long-run tradeoffs between the composition of unemployment and poverty. In particular, following a cut in the minimum wage, unskilled unemployment and poverty rates in the formal sector may well move in opposite directions for particular household groups. Similarly, unskilled unemployment and poverty among urban unskilled households may both fall in the long run, while skilled unemployment and poverty among urban skilled households may well increase. A tradeoff may therefore exist across labor categories. To the extent that such tradeoffs exist, the nature of the social welfare function (that is, the relative importance of the various labor or household groups in shaping government preferences) becomes crucial in choosing a given policy path.

Appendix A
Data Sources and Definition of Variables

This appendix first describes the sources of the data for Brazil and Chile used in this paper. It then lists the countries included in the regression results presented in tables 1 and 2 and provides a more precise definition of the variables used in the regressions, with their respective data sources.

VAR estimates are based on the period 1981–2002 for Brazil and 1981–2001 for Chile. All series are detrended using the modified band-pass filter proposed by Christiano and Fitzgerald (2003), as discussed in the text, and are defined as follows:

Y_CYC is the cyclical component of real GDP calculated as the log difference of real GDP and its trend component. Data sources for real GDP are the World Bank's *World Development Indicators* (WDI) for Brazil, and the Central Bank of Chile (CBC) for Chile.

POVER_CYC represents cyclical components of the poverty gap (for Brazil) and the urban headcount index (for Chile). For Brazil, the source is IPEA (www.ipea.gov.br), and for Chile unpublished estimates by the CBC, which are based on an urban poverty line defined as twice the cost of a representative basket of food.[30]

WAGE_CYC is the cyclical component of the real minimum wage (for Brazil) and the unskilled real wage (for Chile). The source is IPEA for Brazil and for CBC (based on INE surveys) for Chile.

UNEMP_CYC is the cyclical component of the aggregate unemployment rate (for Brazil), and the unemployment rate in the Santiago metropolitan area (for Chile). The source is IPEA (from the monthly employment survey of IBGE) for Brazil and the CBC (based on the monthly survey of the Universidad de Chile) for Chile.

Regressions are based on the following list of countries (years of observation on poverty and unemployment rates in parentheses): Brazil (1985, 1988, 1989, 1993, 1997), Colombia (1988, 1991, 1995, 1996), Costa Rica (1986, 1990, 1993, 1996), Indonesia (1996, 1998), Mexico (1992, 1995), Pakistan (1990, 1993, 1996), Peru (1994, 1996), Philippines (1985, 1988, 1991, 1994, 1997), Sri Lanka (1990, 1995), Thailand (1981, 1988, 1992, 1996, 1998), and Venezuela (1981, 1987, 1989, 1993, 1995, 1996). These countries are all those for which at least two data points on poverty (as measured by the poverty gap or the headcount

30. An unpublished note (available on request) prepared by Elías Albagli of the Central Bank of Chile describes these estimates of the poverty rate in more detail.

index) and the unemployment rate were simultaneously available in the ILO and World Bank databases.

The variables used in the regressions are defined as follows:

POV is the poverty gap and headcount index, calculated with a poverty line of $1.08 a day. Source: World Bank Global Poverty Monitoring Database.

UNEMP is the unemployment rate, defined as the ratio of the labor force that is without work but is available for and seeking employment, to the total labor force. Source: *Key Indicators of the Labor Market* database (ILO).

INFL is the inflation rate in terms of consumer prices. Source: WDI.

REALEX is the percentage change in the real effective exchange rate. A rise is a depreciation. Source: *International Financial Statistics*, IMF.

LGDPPC is the log of GDP per capita measured at purchasing power parity exchange rates. Source: WDI.

TARIFF is the average tariff rate, defined as the ratio of import duties over imports. Source: WDI.

Appendix B
Other Features of Mini-IMMPA

This appendix briefly summarizes some of the other features of Mini-IMMPA, in addition to the production and the labor market structure described in the text.

Both the informal and public sector goods are nontraded. Total supply in each sector is thus equal to gross domestic production. Rural and private urban formal goods, by contrast, compete with imported goods. The supply of the composite good for each of these sectors consists of a combination of imports and domestically produced goods. The demand for imported versus domestic rural and private urban goods is a function of relative domestic and import prices and of the elasticity of substitution between these goods. Allocation of output of rural and private urban formal sector goods to exports or the domestic market occurs along each sector's production possibility frontier. Efficiency conditions require that firms equate the domestic-export relative price to the opportunity cost in production.

For the rural and informal sectors, aggregate demand consists only of intermediate consumption and demand for final consumption (by both the government and the private sector), whereas aggregate demand for the public and private goods consists, in addition, of investment demand. Total demand for intermediate consumption of any good is the sum of the share of this good in the consumption of other sectors. Final consumption for each production sector is the summation across all categories of households of nominal consumption of this sector's good. Total investment by urban firms consists of purchases of private urban formal goods only.

The net or value-added price of output is given by the gross price net of indirect taxes, less the cost of intermediate inputs. World prices of imported and exported goods are exogenously given. The domestic currency price of these goods is obtained by multiplying the world price by the exchange rate, with import prices also adjusted by the tariff rate. Because the transformation function between exports and domestic sales of the rural and private urban goods is linearly homogeneous, the domestic sales prices are derived from the sum of export and domestic expenditure on rural and private goods in nominal terms divided by the quantity produced of these goods. For the informal and public sectors, the composite price is equal to the domestic market price, which is in turn equal to the output price. For the rural sector and private urban production, the substitution function between imports and domestic goods is also linearly homogeneous, and the composite market price is determined accordingly by the expenditure identity. The nested production function of private urban formal goods is once again linearly homogeneous; prices of the composite inputs are derived in a similar fashion. The price of capital is equal to the price of private urban formal goods, because investment expenditure involves only purchases of that category of goods (as noted earlier). Consumption price indices for the rural sector and for urban unskilled and skilled workers are defined as weighted averages of prices of composite goods, with weights reflecting the share of these goods in each group's consumption basket.

Firms' profits in all sectors are defined as revenue minus total labor costs. Firms' income in the rural and informal sectors is equal to their profits, whereas firms' income in the urban formal economy is equal to their profits minus corporate taxes and interest payments on foreign loans. Household income consists of salaries, distributed profits, and government transfers. Households are defined according to both the type of labor and their sector of location, which yields five categories: workers in the rural sector, workers in the urban informal sector, skilled workers in the urban formal sector, unskilled workers in the urban formal sector, and profit earners. The rural household comprises all workers employed in the rural sector. The urban informal household consists of workers in the informal sector. The unskilled (skilled) urban formal household consists of all unskilled (skilled) workers employed in the formal sector, both public and private. Households in the rural sector and in the urban informal economy own the firms in which they are employed. Income of rural sector households is thus equal to the sum of production revenue and transfers from the government. Income of the urban formal skilled and unskilled households depends on government transfers and salaries. Firms provide no direct income, because these groups do not own the production units in which they are employed. Firms in the private urban sector retain a portion of their after-tax earnings to finance investment, and they transfer the remainder to profit earners (who also receive transfer payments).

Each category of households saves a constant fraction of its disposable income, which is equal to total income minus income tax payments. The portion of disposable income that is not saved is allocated to consumption. The accumulation of capital over time depends on the flow level of investment and the depreciation rate. The aggregate identity between savings and investment implies that total investment must be equal to total savings, equal to firms' after-tax retained earnings, total after-tax household savings, government savings, and foreign borrowing by firms. In the simulations reported in the text, this equation is solved residually for the level of private investment.

All value added in the production of public goods is distributed as wages. Government expenditures consist of government consumption and public investment, which consists of investment in infrastructure, education, and health. Infrastructure and health capital affect the production process in the private sector, as they both combine to produce the stock of public capital. Tax revenues consist of revenue generated by import tariffs, sales taxes, income taxes (on both households and firms in the private urban sector), and payroll taxes. Thus, the fiscal deficit is equal to tax revenue minus transfer payments, current expenditure on goods and services, total wage payments, and total investment expenditure. Finally, the external constraint implies that any current account surplus (or deficit) must be compensated by a net flow of foreign capital, given by the change in private and public foreign borrowing. This is obtained by an adjustment of the real exchange rate.

Appendix C
Calibration and Parameter Values

This appendix presents the characteristics of the data underlying the calibration procedure for the Mini-IMMPA prototype described in the text (see Agénor, 2003). The basic data set consists of a social accounting matrix (SAM) and a set of initial levels and lagged variables. The SAM encompasses twenty-seven accounts, including production and retail sectors (four accounts), labor production factors and profits (three accounts), enterprises (one account), households (five accounts), government current expenditures and taxes (nine accounts), government investment expenditures (three accounts), private investment spending (one account), and the rest of the world (one account). The data satisfy the double-entry accounting principle and can therefore be used to initialize model variables and calibrate level parameters, such as effective tax rates.

The characteristics of the SAM data and other data (including initial labor market quantities and debt and capital stocks) are summarized as follows. On the output side, agriculture and the informal sector account for 12 and 35 percent of total output, respectively. On the demand side, private current and capital expenditures account for 78 percent of GDP, whereas overall government expenditures account for 18 percent of GDP. The economy has a balanced current account but runs a trade surplus, amounting to 4 percent of GDP, to finance foreign interest payments.

Total investment expenditures amount to 22 percent of GDP, and the private sector account for two-thirds of these outlays. This implies that investment spending accounts for 19 percent of private expenditures and 40 percent of public expenditures. The public sector investment budget allocates 30 percent of expenditures to investment in the health sector, 30 percent to investment in the education sector, and 40 percent to investment in infrastructure. Furthermore, the public sector wage bill makes up 30 percent of overall public sector expenditures. The government is assumed to run a balanced budget in the base period and thus does not resort to domestic or foreign borrowing. Sales taxes and import tariffs make up more than 70 percent of total government revenues, whereas private income and corporate taxes account for less than 20 percent of revenues.

The trade balance is dominated by nonagricultural imports and exports. Agricultural exports account for only 8 percent of total export earnings, whereas nonagricultural imports account for 92 percent of total imports. The level of trade openness (measured by the ratio of the sum of imports and exports to GDP) amounts to a moderate 40 percent. Because the economy runs a balanced current account in the base period, there is no private or public foreign borrowing. Nevertheless, the stock of external debt in the base period amounts to 51 percent of GDP (or 233 percent of export earnings), whereas foreign interest payments amount to 4 percent of GDP (or 18 percent of exports earnings).

Rural areas account for 29 percent of the total labor force, and the rest is concentrated in urban areas. Altogether, 47 percent of the workers are employed in some kind of urban informal occupation, whereas only 22 percent of the labor force is employed in the urban formal sector. Open unemployment among urban formal workers amounts to 2 percent of the total labor force. The formal labor force consists of 58 percent unskilled workers and 42 percent skilled workers, and unemployment rates are 10 percent among formal unskilled workers and 8 percent among skilled workers. Migration from rural to urban areas amounts to 1.3 percent of the rural population, and the urban-rural wage differential amounts to 54 percent of the rural wage. In comparison, unskilled labor migration from the informal to the formal sector amounts to 0.8 percent of the informal sector labor force, and the formal-informal wage differential amounts to 106 percent of the informal wage.

Seventeen elasticity parameters, which cannot be derived from the calibration procedure, have to be estimated. These parameters include constant elasticity of substitution (CES) in rural agricultural and private formal production (four parameters); CES Armington elasticities and constant elasticity of transformation (CET) for aggregating domestic composite goods and transforming domestic production (four parameters); elasticities related to rural-urban, and formal-informal sector migration (two parameters); elasticities related to the computation of ordinary and congested government capital (two parameters); the elasticity of effort by teachers and the elasticity of substitution between labor and capital in skill upgrading (two parameters); the elasticities related to determination of skilled labor wages (two parameters); and the elasticity of investment with respect to the desired private capital stock (one parameter). In addition, a set of minimum consumption levels (fifteen parameters) has to be determined, because they similarly cannot be derived from the calibration procedure.

The substitution elasticity between labor and government capital in rural production is set at 0.7, whereas elasticities in the nested private, formal sector production structure range from 0.7 between skilled labor and capital to 1.2 between the skilled labor-capital bundle and unskilled labor. Import and export elasticities are uniformly set at 0.7 for agriculture and 1.5 for the private urban formal sector. This is again meant to reflect a lower-middle-income economy with low agricultural potential. The elasticity of rural-urban migration with respect to the relative rural-urban wage-differential is set at 0.4, while the elasticity of formal-informal migration with respect to the formal-informal wage ratio is set at 0.8.

The substitution parameter between infrastructure and health capital stocks are set at 0.5, and congestion is assumed to be absent (zero elasticity). The substitution elasticity between teachers and public capital in education in the production of skilled labor is set at 0.3, while the effort elasticity with respect to the relative wage ratio—a specification that follows Agénor and Aizenman (1999)—is set to 0.8. Furthermore, skilled wages in the private urban formal sector are assumed to be affected only by the skilled unemployment rate, with an elasticity of –2.0. Finally, the elasticity of private investment with respect to the desired growth rate of the private capital stock is set at 0.3.

Among the remaining parameters, the foreign interest rate on private borrowing is calibrated to 3.8 percent, while the public foreign interest rate is calibrated to 4.9 percent. In addition, the initial depreciation rates are calibrated to 6.4 percent for private capital and 3.9-5.8 percent for public capital (depending on whether investment is in education, health, or infrastructure).

Turning to the government budget, output and value-added tax rates range from 3.0 to 3.7 percent, whereas the tax rate on sales of the private urban formal sector and the payroll tax rate paid by firms in that sector are calibrated to 12.1 and 20.1 percent, respectively. Import tariffs range from 34 percent on private formal sector goods to 167 percent on agricultural goods, reflecting a country with significant protection on agriculture. Finally, the corporate income tax rate is set at 7.6 percent, while income tax rates on households range from 2.2–3.9 percent for rural agricultural and urban unskilled groups to 9.6–12.5 percent for the urban skilled group and capitalists-rentiers. As noted in the text, workers in the urban informal sector do not pay income taxes.

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Are Labor Market Regulations an Obstacle for Long-term Growth?

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Labor markets across the world are usually characterized by a set of institutions that limit the ability of private agents to determine wages and the amount of labor required and by tax systems that transfer resources from the working to the nonworking population via unemployment benefits, employment protection laws, and active employment policies by the government (Saint-Paul, 1999).

One strand of the economic analysis claims that labor institutions reduce the rate of job creation and increases unemployment (Salvanes, 1997; Nickell, 1997; Blanchard and Wolfers, 2000). This process has an adverse impact on economic growth (Besley and Burgess, 2004; Forteza and Rama, 2002). Supporters of this approach usually suggest the reduction or elimination of labor market regulations in order to foster labor reallocation and increased competition, which in turn enhances growth (Burki and Perry, 1997). Labor market reforms, however, have proved to be politically unfeasible and have faced significant opposition from powerful sectors of the economy (Alesina and Drazen, 1991).

A second strand of the analysis holds that the behavior of labor markets is far from competitive (Freeman, 1993a; Blanchard, in this volume). Proponents suggest that in the presence of market failures, governments should set up regulations for the proper functioning of the labor markets. Labor market regulations are introduced to enhance the welfare of workers and insure them from unexpected shocks. For

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example, legislation on social security and mandated benefits was designed to secure the workers' income in case of old age, sickness, disability, and work-related accidents. Job security provisions are similarly undertaken to insure an income for workers who lose their jobs during economic downturns (Heckman and Pagés, 2000).

This paper tests whether labor market regulations have been an obstacle to long-run growth. Using two recently developed databases on labor regulation by Rama and Artecona (2002) and Djankov and others (2003), we perform our regression analysis for a sample of seventy-six countries over the period 1970–2000 in the tradition of empirical growth literature (Barro, 1991, 1997). Our analysis is performed on both cross-sectional and panel data; in both cases we control for the likely endogenous regressors. Our set of instruments consists of both external instruments chosen from theories of selection of labor regulations (Djankov and others, 2003a) and internal instruments, which are lagged levels or differences in panel data models (Arellano and Bond, 1991; Arellano and Bover, 1995).

Our cross-sectional results indicate that growth in industrial countries is hampered by thicker labor codes. The effect of de facto labor regulations on growth is mixed, depending on the data, indicator, or sample used. Finally, regulations stipulated in labor laws regarding employment, industrial relations, and social security seem to have an impact on growth only for industrial economies.

Our panel data regression analysis yielded five main results. First, growth among developing countries could be fostered by reducing the regulations stipulated in the national labor codes. However, deregulation processes in labor markets commonly succeed at reducing the number of regulations in labor laws, but they cannot improve the strength of legal enforcement mechanisms.

Second, growth in industrial countries could be enhanced by lower de facto regulation. According to our estimates, a one-standard-deviation decline in the labor regulations index could generate a 2 percent increase in growth in advanced economies. Achieving these growth effects would require an enormous deregulation effort from this group countries—especially the European economies. Unfortunately, most European countries have only marginally changed their labor institutions (Siebert, 1997).

Third, a decline in the degree of de facto labor regulation may also improve the growth rate of developing countries. Our regression results suggest that growth in developing countries would increase by 0.6 percentage points in response to a one-standard-deviation decrease in the de facto labor regulations. However, obtaining significant growth benefits in these nations would require substantial labor market reform (that is, a sharper decline in labor regulations than that seen in the data over the period of analysis).

Fourth, the transmission channels of the adverse growth effects of higher labor regulations within developing countries are minimum wages and trade unions. Again, we find that the growth effects obtained from one-standard-deviation reductions in both variables are not plausible unless the countries embark on serious efforts to deregulate labor markets.

Finally, we construct a scorecard that summarizes the different panel estimation results presented in this paper. Our most robust results are that thicker codes are negatively associated with growth for the full sample and for developing countries; de facto regulations seem to hinder growth among developing countries; and minimum wages have an adverse impact on growth.

The rest of the paper is divided into four sections. Section 1 presents a brief review of labor regulations and economic performance. Section 2 discusses the data used and the methodology applied. Section 3 empirically evaluates whether labor regulations have hindered long-term growth. Finally, section 4 concludes.

1. Literature Review

There is a vast empirical literature evaluating the effects of economic policy on growth for a cross-section of countries. Recent evidence indicates that government interventions may have an important effect on growth. Hall and Jones (1999) show that poor social infrastructure (which they approximate by poor contract enforcement, low bureaucratic quality, and government repudiation of contracts, among others) is negatively related to long-term growth. The degree and type of government intervention vary across countries. Djankov and others (2002) analyze the regulations on starting a business across countries; they find that countries with extensive regulations on entry may enlarge their informal sectors and, hence, have a poor economic performance.[1]

1. Djankov and associates argue that this empirical result is consistent with economic regulations legislated and imposed by government officials or insiders that extract rents (see Shleifer and Vishny, 1998).

In policy circles, the differing viewpoints on the role of labor market regulations in the economic process fall into two broad groups (Freeman, 1993a). On the one hand, the distortionist view argues that government regulations in labor markets (in the form of minimum wages, social security contributions, job security, and collective bargaining) create distortions in an optimal world (World Bank, 1990). According to this view, labor market regulations are major obstacles to growth and employment for three main reasons. First, labor market regulations prevent wages from equaling their marginal product in equilibrium and thus lead to the misallocation of resources. Second, regulations may hinder the adjustment of labor markets to economic shocks. Finally, labor regulations that redistribute economic rents from capital to labor may reduce the profitability of investment; examples include collective bargaining schemes and expansionary fiscal programs to fund public employment. This reduced profitability may discourage investment and lower growth.[2]

On the other hand, the institutionalist view claims that market failures generate divergences from the ideal world and emphasizes the benefits of government interventions in the labor markets (ILO, 1991). Labor regulations may fulfill redistributive roles to low-wage workers or constitute an insurance against adverse market outcomes (Standing and Tokman, 1991). Labor standards force employers to focus on enhancing their labor force through either training or technical innovations (Freeman, 1993a, 1993b). Finally, standards on mandated benefits may help solve moral hazard or selectivity issues that prevent firms from offering socially desirable benefits or contracts (Summers, 1988).

Forteza and Rama (2002) evaluate the role of labor market regulations in the success of economic reforms. They find that wage adjustment and labor reallocation in outward-oriented economies will be faster if labor markets are flexible. International competition in the goods markets will drive down wages in the import-competing sectors and labor costs in the economy, thus making the export sector more competitive. If labor markets are not flexible, the adjustment in the economy will be slower and the unemployment rate will

2. Freeman (1993a, 1993b) argues that the distortionist view is not consistent with macro- and microeconomic propositions derived from economic theory. For example, Ricardian equivalence is rejected by those who argue that social security contributions have a negative impact on investment and savings. Also, the Coase theorem is not taken into account when distortionists claim that employment laws have efficiency costs.

be higher (Rama, 1997). Furthermore, current labor laws have been an obstacle to absorbing workers displaced by economic reform (IDB, 1997). The usual recommendation, therefore, is to eliminate government interventions that make labor costly and risky (Burki and Perry, 1997).

Potential losers from economic reforms, such as workers in the public sector or unionized labor, usually try to hinder or delay the economic adjustment process (Alesina and Drazen, 1991; Fernandez and Rodrik, 1991). High resistance to economic reforms from well-organized groups may lead to generalized protests and strikes. In response, the government may delay the adoption of reforms or launch an insufficient package of reforms, which, in turn, would have an insignificant impact on economic performance. This leads to the argument that resistance to reforms will be weaker when the distribution of adjustment costs is relatively equal and that economic reforms should be complemented with compensation mechanisms for workers affected by the reforms, such as job separation packages, early retirement programs, and unemployment benefits (Rama, 1995; Forteza and Rama, 2002).

Forteza and Rama (2002) find that labor regulations that determine the success or failure of structural reforms work through more political channels –proxied by unionization and government– , which are correlated with deeper recessions before adjustment and slower recoveries in the aftermath. On the other hand, economic aspects –measured by minimum wages or mandated benefits– do not seem to hinder growth. Finally, Besley and Burgess (2004) assess the role of labor market regulations in explaining the performance of the Indian manufacturing industry between 1958 and 1992. They find that regulations to protect workers (in the areas of collective bargaining and labor disputes) actually reduced growth and increased poverty.

2. The Data and Methodology

In this section, we discuss the labor regulation data used in our regression analysis and the estimation strategy pursued. First, we describe two recently developed databases on labor regulations and outcomes: the aggregate and individual measures proposed by Rama (1995) and Rama and Artecona (2002); and the indicators of labor market regulations gathered from labor codes by Djankov and others (2003a). Second, we outline the estimation techniques used to test the impact of labor regulations on long-term growth. Our preferred estimation technique is the generalized method of moments instrumental variables (GMM-IV) system estimator (Arellano and Bover, 1995; Blundell and Bond, 1998), which takes into account the unobserved country- and time-specific effects, as well as possible endogenous regressors, in a dynamic panel data model.

2.1 The Data

As mentioned above, we use two different databases on labor market regulation and outcomes. The Rama-Artecona database has information on a larger sample of countries (121), it has a panel dimension (five-year average observations spanning 1945 to 1999), and it allows us to distinguish between regulations on paper and in practice.[3] The database from Djankov and others (2003a), which we denote the Djankov-La Porta database, covers a smaller sample of countries (eighty-five) and contains only cross-sectional information. It specifically gathers information on three types of labor laws (employment, industrial relations, and social security) for the year 1997. Next, we further describe the main features of these databases.

The Rama-Artecona Database

Rama and Artecona (2002) have collected extensive information on labor market regulations and outcomes for 121 countries. They report the data in five-year-period averages from 1945–49 to 1995–99. In this database, we can distinguish between regulations on paper (or de jure regulations) and regulations in practice (or de facto regulations). De jure regulations are approximated by eight indicators of International Labor Organization (ILO) labor standards as ratified and stipulated by legal documents in several countries. These conventions contemplate universal legislation on issues such as child labor, compulsory labor, equal remuneration for male and female workers, equal opportunity, the right of collective bargaining, and organization in unions. De facto regulations and labor market outcomes are approximated by thirty-six indicators classified into six categories: labor force; employment and unemployment; wages and productivity; work conditions and benefits; trade unions and collective bargaining; and public sector employment.

3. See appendix A for the list of countries.

Here Rama and Artecona provide information on labor market regulations such as minimum wages, mandated benefits, nonwage costs, collective bargaining, and public employment, as well as labor market outcomes such as labor force, unemployment, earnings, and productivity.

Distinguishing between de jure and de facto regulations is crucial given that the enforcement of regulations and norms stipulated in labor codes is quite limited in developing countries.[4] We thus follow Rama (1995) and Forteza and Rama (2002) in defining aggregate indices of the overall extent of labor regulations in the economy. Our index of de jure regulation, which we denote $\mathbf{L}_0$, is measured as the cumulative number of ILO conventions ratified by a country's labor code over time. This index reflects not only the ideal regulatory framework from an institutionalist point of view (Freeman, 1993a), but also the thickness of national labor codes (Rama and Forteza, 2002). The index includes the ratification of the ILO conventions on the minimum age of employment (convention 138), forced or compulsory labor (convention 29), the abolition of forced labor (convention 105), equal remuneration for male and female workers (convention 100), discrimination with regard to equality of opportunity or conditions of employment on the basis of race, religion, sex, political opinion, or social origin (convention 111), the right of workers and employers to establish associations or organizations of their own (convention 87), and the right to bargain collectively (convention 98).

As mentioned, however, the extent of regulation in the labor market depends on the way these legal regulations are implemented and enforced. Therefore, we require an indicator that captures the degree of enforcement as opposed to the number of regulations. Rama and Artecona (2002) provide measures for regulations in the following four areas: minimum wages, mandated benefits, trade unions, and public sector employment. Unfortunately, no data are available on job separation costs for a large number of countries.[5] To evaluate the overall effect of labor reforms in these dimensions, we follow Rama (1995) and Forteza and Rama (2002) in constructing two aggregate indices of

4. Squire and Suthiwart-Narueput (1997) suggest that de jure regulations that appear to be more distortionary in developing countries could be the least enforced in practice.

5. Heckman and Pagés (2000) construct data on job separation costs for Latin America and find that these costs have a substantial impact on the level of employment in the region.

Table 1. Indicators of Labor Market Regulations

Category	*Aggregate index L1*	*Aggregate index L2*
Minimum wages	Ratio of minimum wages to labor costs per worker in the manufacturing sector	Ratio of minimum wages to income per capita
Mandated benefits	Social security contributions as a percentage of salaries	Number of days of maternity leave for a first child born without complications
Trade unions	Total trade union membership as a percentage of total labor force	Dummy: Ratification of ILO convention 87, which allows workers to organize
Government employment	Ratio of general government employment to total employment	Ratio of central government employment to total employment

de facto labor regulations. Both proposed indices include proxies for these four dimensions of labor regulations, as summarized in table 1.[6]

Both aggregate indices, $\mathbf{L}_1$ and $\mathbf{L}_2$, are the simple averages of the proxies in the four dimensions. We normalized all the labor regulation indicators so that these variables are comparable across countries. Specifically, their values fluctuate between 0 and 1, with higher values reflecting a higher degree of labor market regulation. Finally, the aggregate indices, $\mathbf{L}_1$ and $\mathbf{L}_2$, are computed for countries with information for at least two of the four dimensions of the analysis.

The Djankov-La Porta Database

Djankov and others (2003a) have collected data on labor regulation in eighty-five countries. They analyze three dimensions of the national labor codes: laws governing individual employment contracts (employment laws); laws regulating the adoption, bargaining, and enforcement of collective agreements, the organization of trade unions, and the industrial action by workers and employers (industrial and collective relations law); and laws governing the social response to needs and conditions that affect the quality of life, such as old age, disability, death, unemployment, and maternity (social security law).[7]

6. The higher degree of correlation between the different dimensions of the labor regulation index prevents us from including all the variables of the aggregate index in the same regression.

7. In contrast with the Rama-Artecona database, we only have cross-sectional information on these variables.

We first use the aggregate index of employment laws, which regulate aspects of the individual labor contract, terms of reference, and termination of the contract. This index covers the restrictions placed on alternative employment contracts, conditions of the employment contract, and job security. Next, we have the aggregate index of industrial relations laws, which protect workers from employers. These laws contemplate aspects of the worker-employer relationship such as collective bargaining, the participation of workers in management, and collective disputes (for example, strikes and lockouts). Finally, we have the aggregate index of social security laws covering the risk of old age, sickness, and unemployment. Since labor laws (rather than outcomes) are used to construct all these indices, they are closer in spirit to de jure labor rigidities than de facto implementation in Rama and Artecona (2002).

Growth and its Determinants

Our dependent variable is the growth rate of gross domestic product (GDP) per capita, and we obtain the data from the Penn World Table 6.1 gathered by Heston, Summers, and Aten (2002). Specifically, we use the real GDP per capita (chain index prices). We follow the vast existing empirical growth literature in choosing the determinants of long-run economic growth.[8] We include the initial GDP per capita (in logs) to test for transitional convergence. We also consider structural factors such as the level of secondary schooling from Barro and Lee (2000) as a proxy of human capital; credit to the private sector as a ratio to GDP to measure financial depth (Beck, Demirgüç-Kunt, and Levine, 2000); the ratio of real exports and imports to GDP as a measure of trade openness; and the Freedom House index of civil liberties as a proxy of governance. Data on the consumer price index (CPI) inflation rate and real exchange rate overvaluation are obtained from the World Bank's *World Development Indicators,* which proxy for stabilization policies. Finally, changes in the terms of trade (as a proxy for external shocks) are also taken from *World Development Indicators*.

8. The set of growth determinants follows the classification of Loayza, Fajnzylber and Calderón (2003).

2.2 The Empirical Framework

This subsection evaluates the role of labor market rigidities in long-term growth following the traditional empirical growth literature. Our regression framework is specified by the following system:

$$dy^{*}_{it,t-k} = \mu_i + \eta_t + \alpha y_{it-k} + X_{it}\beta \text{ and} \tag{1}$$

$$dy_{it,t-k} = dy^{*}_{it-k} + L_{it}\Gamma + \xi_{it} \ .$$

According to the first equation of system 1, the equilibrium growth rate of the economy in country i during the $[t, t-k]$ period, $dy^{*}_{it,t-k}$, is a function of the log of per capita output in the initial period $t-k$, y_{it-k}; a set of growth determinants for country i at time t described by the matrix $\mathbf{X}_{it}$; and unobserved country- and period-specific effects, μ_i and η_t, respectively. Our set of long-term growth determinants follows the work of Loayza, Fajnzylber, and Calderón (2003). The initial level of per capita output (in logs) is included to test for conditional convergence. We consider indicators of human capital, financial depth, trade openness, and governance as proxies for structural policies and institutions. The CPI inflation rate and the real exchange rate overvaluation are proxies for stabilization policies, and terms of trade shocks approximate external shocks.[9]

In the spirit of Rama (1995), our second equation in the system indicates that any deviation in long-term equilibrium growth may be explained by a set of variables that proxy for departures from competition in the labor markets, L_{it}. This matrix, $\mathbf{L}$, is our variable of interest; it may comprise different indicators that focus on specific policy or institutions in the labor markets, such as minimum wages, mandatory benefits, trade union membership, government employment, social security laws, and collective bargaining. We denote by

$$\left\{\ell^{k}_{it}\right\}_{k=1}^{K}$$

all the K indicators of labor market rigidities comprised in the matrix, $\mathbf{L}_{it}$. Unlike Rama (1995) and Forteza and Rama (2002), we do not

9. We follow the tradition of empirical cross-country and panel growth regression models in focusing on the ultimate policy, structural, and external determinants of factor accumulation and productivity growth. Hence, we exclude capital and any other direct factor of production.

assume that labor market policies and institutions are time-invariant, but rather expect that labor institutions may change over longer horizons. If any of the $\ell_{it}{}^{k}$ variables equals zero, labor markets are perfectly competitive. In contrast, larger values for any of these variables indicates greater deviation from perfect competition in the labor market. Negative values for the γ_k coefficients in the Γ matrix imply that the reduction of labor rigidities (that is, distortions that cause labor markets to depart from competitive equilibrium) may improve the growth rate in the long term.

Performing a regression analysis of equation (1) may raise additional empirical problems. Some of the $\ell_{it}{}^{k}$ variables are highly correlated with each other, thus leading to problems of multicollinearity. For example, the correlation between trade union membership and government employment is approximately 0.8, whereas mandated benefits and minimum wages have a correlation of 0.5. This problem of collinearity impedes the identification of the parameters of the Γ matrix.

We address the issues of collinearity among labor regulation indicators by aggregating the variables in the $\mathbf{L}_{it}$ matrix, using the same strategy as Rama (1995) and Forteza and Rama (2002). Before we aggregate them in a single index, we need to normalize them so as to express them in comparable units. We defined our labor market rigidity indicator above as $\ell_{it}{}^{k}$, for $k = 1,\dots, K$. Next, we define $\ell^{k}_{\min}$ and $\ell^{k}_{\max}$ as the minimum and maximum deviations from perfect competition that a country's labor market can achieve. We can thus specify our normalized labor market rigidity indicator as follows:

$$\tilde{\ell}_{it}{}^{k} = \frac{\ell_{it}{}^{k} - \ell^{k}{}_{\min}}{\ell^{k}{}_{\max} - \ell^{k}{}_{\min}} \ .$$

By construction, $\tilde{\ell}_{it}{}^{k}$ fluctuates between zero and one. We then define our aggregate measure of labor market rigidities as the average of J out of the K relevant labor market rigidities (where $J \le K$). In principle, this aggregate index also ranges from zero to one, but unless all of the labor market rigidities are perfectly correlated with each other, the actual range of variation across countries should be significantly narrower for the aggregate measures than for any of the individual indicators.

We use our aggregate index of labor market rigidities, $\ell_{it}{}^{A}$, to test the effects of the overall labor market rigidity on growth. We reformulate our growth equation in system 1 as

$$dy_{it,t-k} = dy^{*}{}_{it,t-k} + \gamma_A \ell_{it}{}^{A} + \xi_{it} \ . \qquad (2)$$

The sign and order of γ_A can be used to check the nature and magnitude of the impact of labor rigidities on growth. However, different labor market rigidities may have consequences of a different sign that cancel each other to some extent. Even if the estimate of the parameter γ_A turned out to be significant, its mere sign might not help identify the specific policies and institutions that need to be reformulated. We still need more information on the sign and order of magnitude of the γ_j parameters.

We are tempted to use equation (2) to test for the effects of particular labor market rigidities. If $\ell_{it}{}^{A}$ is replaced by $\tilde{\ell}_{it}{}^{k}$ in equation (2), the coefficient multiplying it captures not only the effects of the labor market regulation, k, but also (partly) those of all of the other missing rigidities. Since they are likely to be correlated with each other, the value obtained for γ_k might be reflecting the effects of these other rigidities. For example, let us assume that unionized labor does not affect growth, but minimum wages do, and that minimum wages tend to be higher in countries with larger labor unions (actually we find a correlation of 0.5 between these variables). If we include minimum wages in equation (2) instead of $\ell_{it}{}^{A}$, we obtain a significant estimate for this variable even though it should be statistically and economically irrelevant. This problem can be partially corrected by defining the complementary labor regulation variable, $\tilde{\ell}_{it}{}^{-k}$, as the average of the indicators that are different from k. This complementary variable can be used to control for all other labor market features, apart from $\tilde{\ell}_{it}{}^{k}$, by using the following model:

$$dy_{it,t-k} = dy^{*}{}_{it,t-k} + \gamma_k \tilde{\ell}_{it}{}^{k} + \gamma_{-k} \tilde{\ell}_{it}{}^{-k} + \xi_{it} \ , \tag{3}$$

with the coefficient γ_k capturing the effect of labor market regulation k on long-term growth.

2.3 Estimation Techniques

We first estimate the growth regression equation specified in equation (1) using pooled ordinary least squares (OLS). We then run regression again incorporating time dummies, given that we want to analyze differences in growth experiences across countries stemming from labor rigidities. Neither of these methods, however, controls for endogenous regressors. Forces that affect both labor rigidities and growth could be driving the correlation between the variables, and our estimates may be biased.

One way to tackle the problem of endogeneity is to instrument for labor rigidities. We follow Djankov and others (2003a) in choosing the appropriate instruments for our measures of labor institutions. According to these authors, three theories explain the choice of labor institutions: efficiency theory, political power theory, and legal theory. Of these, North (1981) considers that the choice of institutions is driven primarily by efficiency considerations. Different institutional arrangements (such as the reliance on market forces, contract and private litigation, and government regulation) may be appropriate in different circumstances. One version of efficiency theory focuses on the distinction between regulation and social insurance. Social insurance may be relatively more efficient than regulation in dealing with market failures in countries with a low social marginal cost of tax revenues, which presumably are the wealthy countries (Becker and Mulligan, 2000). Poor countries must regulate to protect workers from being fired or mistreated by employers, whereas rich countries provide unemployment insurance, sick leave, early retirement, and so on because they can raise taxes cheaply to finance such operations (Blanchard, 2002). A second version of efficiency theory argues the opposite. It holds that the principal cost of regulation, relative to other forms of social control of business, is its potential for abuse of regulated firms by the government and its officials. Labor regulations can be used to force firms to hire and keep excess labor, to empower unions friendly with the government, and so forth. Rich, well-governed countries thus have a comparative advantage at regulation relative to other forms of social control of business because their governments are less likely to abuse power.

Political power theories argue that institutions are designed to transfer resources from those out of political power to those in power and to entrench those in political power (Olson, 1993). Institutions are generally designed to be inefficient by political leaders aiming to help themselves and their favored groups. Regulations protecting workers are introduced by socialist, social-democratic, and generally leftist governments to benefit their political constituencies (Hicks, 1999). In addition, labor regulations are a response to pressure from trade unions, and they should thus be more extensive when unions are more powerful, regardless of which government is in charge. Dictatorships, which are less constrained than democratically elected governments, tend to have more redistributive laws and institutions. Constitutions, legislative constraints, and other forms of checks and balances are all conducive to fewer regulations (Djankov and others, 2002). Likewise, economies that are open to trade may be less likely than closed economies to introduce expensive regulations, because competition makes it less lucrative for governments to raise firms' regulatory costs (Ades and Di Tella, 1999).

With regard to legal theory, Djankov and others (2003b) argue that countries with different legal traditions use different social controls of business. Common law countries tend to rely on markets and contracts, civil law countries on regulation, and socialist countries on state ownership.[10] This implies that civil law countries and socialist law countries should regulate labor markets more extensively than common law countries. Common law countries may also have a less generous social security system since they rely on markets to provide insurance.

Our set of instruments for labor rigidity indicators is as follows. We use the log of GDP per capita to control for efficiency purposes. To test the political power theories, we use the index of institutionalized autocracy from the Polity IV codebook (Marshall and Jaggers, 2003) the leftist political orientation of the government and congress (Beck and others, 2001), and measures of trade openness. Finally, we test the legal theory by including dummy variables for countries with British common law and German civil code (La Porta and others, 1999).

Another way to tackle the endogeneity of labor rigidities is to use the GMM estimators developed by Arellano and Bover (1995) and Blundell and Bond (1998). This technique takes account of unobserved time effects through the inclusion of period-specific dummy variables, while country-specific effects are dealt with via differencing, given the dynamic nature of the regression. We also control for biases resulting from simultaneous or reverse causation. A more detailed reference to the GMM-IV techniques is presented in appendix B.

10. Common law emerged in England and is mostly characterized by the importance of decisionmaking by juries, independent judges, and judicial discretion as opposed to codes. Common law was transmitted to the British colonies, including Australia, Canada, New Zealand, India, Pakistan, the United States, and a number of countries in the Caribbean, East Africa, and Southeast Asia. Civil law evolved from Roman law in Western Europe and was incorporated into civil codes in France and Germany in the nineteenth century. It is characterized by less independent judiciaries, the relative unimportance of juries, and a greater role of both substantive and procedural codes as opposed to judicial discretion. French civil law was transplanted throughout Western Europe, including Belgium, Holland, Italy, Portugal, and Spain, and subsequently to the colonies in North and West Africa, Latin America, and parts in Asia. German codes became accepted in Germanic Western Europe, but were also transplanted to Japan and from there to China, Korea, and Taiwan. Socialist law was adopted in countries that came under the influence of the Soviet Union, while an indigenous Scandinavian legal tradition developed in Denmark, Finland, Iceland, Norway, and Sweden (Djankov and others, 2003).

3. Empirical Assessment

In this section we empirically evaluate whether labor market regulations have hindered long-term growth. We perform our regression analysis, first, on a cross-section sample of seventy-six countries with average figures for the 1970–2000 period and, second, on panel data for the same sample of countries with five-year averages over the same period. We use both the Rama-Artecona and Djankov-La Porta databases for the cross-section and only the former for the panel analysis.

We begin by presenting some basic statistics on the extent of labor market regulations and economic growth. We then perform a cross-sectional and panel data correlation analysis between growth and labor regulations. Next, we discuss the basic results of the growth regression in the cross-section of countries, followed by the panel data evidence on growth and labor regulations using different estimation techniques. Finally, we present our scorecard on the growth costs of labor regulations.

3.1 Basic Statistics

Table 2 reports the simple average of the growth rate in per capita GDP and different indicators of labor market regulation for a cross-section of countries during the 1970–2000 period. It includes both the simple averages for the Rama Artecona indicators of labor rigidity and the averages of the labor regulation indicators from the Djankov-La Porta database.

Based on the Rama-Artecona de jure index, we find that industrial countries are more regulated than developing countries (0.49 versus 0.25, respectively). Labor markets in Latin America are more regulated than the world sample (0.34 versus 0.30), whereas East Asia is less regulated than the world sample (0.09). Within the Latin American region, Chile has a similar number of regulations to the regional average, while Uruguay (not shown in the table) has the largest number of regulations (0.67). Both the Rama-Artecona de facto indices ($\mathbf{L}_1$ and $\mathbf{L}_2$) indicate that industrial countries exhibit a larger degree of labor market regulations than developing countries. If we use the $\mathbf{L}_2$ index of de facto regulations, Latin American labor markets are as regulated as labor markets in industrial economies. Chilean labor markets are less regulated than the Latin American average regardless of the aggregate index used.

The table also lists the components of the two aggregate indices of de facto labor regulations. Minimum wages, for example, are higher

Table 2. Basic Statistics for Labor Market Regulations and Economic Growth, 1970–2000

Simple averages across groups of countries

Variable	*Full sample*	*Industrial economies*	*Developing countries*	*East Asia*	*Latin America*	*Chile*
GDP per capita growth (percent)	1.60	2.20	1.40	4.30	0.90	2.40
Labor market rigidity[a]						
De jure index $\mathbf{L}_0$	0.30	0.49	0.25	0.09	0.34	0.33
De facto index $\mathbf{L}_1$	0.28	0.36	0.25	0.18	0.25	0.17
Minimum wage[b]	0.23	0.24	0.22	0.22	0.21	0.14
Social security contribution	0.37	0.45	0.35	0.26	0.35	0.40
Trade union membership	0.24	0.39	0.20	0.15	0.18	0.11
General government employment	0.27	0.39	0.22	0.16	0.25	0.05
De facto index $\mathbf{L}_2$	0.29	0.32	0.28	0.14	0.32	0.08
Minimum wage[c]	0.14	0.09	0.16	0.10	0.10	0.06
Maternity leave (no. days)	0.16	0.19	0.15	0.13	0.13	0.18
Ratification of ILO convention 87	0.59	0.79	0.54	0.17	0.78	0.03
Central government employment	0.16	0.19	0.16	0.11	0.21	0.03
De jure versus de facto						
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	–0.04	–0.12	–0.01	0.08	–0.09	–0.16
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	–0.02	–0.17	0.03	0.06	–0.02	–0.26
Labor regulation[d]						
Employment laws	1.53	1.36	1.60	1.39	1.79	1.46
Alternative employment contracts	0.56	0.58	0.56	0.57	0.55	0.58
Conditions of employment	0.62	0.49	0.67	0.52	0.73	0.58
Job security	0.35	0.28	0.37	0.30	0.50	0.31
Industrial (collective) relations law	1.25	1.22	1.26	1.12	1.44	1.18
Collective bargaining	0.51	0.46	0.53	0.37	0.68	0.78
Worker participation in management	0.23	0.32	0.20	0.27	0.15	0.00
Collective disputes	0.51	0.44	0.53	0.49	0.60	0.40
Social security laws	1.70	2.21	1.53	1.58	1.69	1.98
Old age, disability, and death benefits	0.57	0.68	0.53	0.56	0.53	0.46
Sickness and health benefits	0.65	0.75	0.62	0.69	0.74	0.79
Unemployment benefits	0.48	0.78	0.38	0.33	0.42	0.73

Source: Authors' calculations, based on data from Rama and Artecona (2002) and Djankov and others (2003a).
a. Indicators of labor market rigidity are from Rama and Artecona (2002).
b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.
c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.
d. Indicators of labor regulations are from Djankov and others (2003a).

(lower) among industrial countries than among developing countries when they are normalized by industrial wages (per capita income). Minimum wages in East Asia are below those in Latin American markets, with Argentina, Chile, and Peru achieving the lowest minimum wages in the region (when normalized by per capita income; again, only the case of Chile is shown in the table).

In the case of mandated benefits, social security contributions (normalized by total wages) are larger among industrial countries than among developing countries (0.45 versus 0.35, respectively). Chile's social security contributions (0.40) are larger than the averages for both the region (0.35) and East Asia (0.26). Industrial countries also have longer maternity leave than developing countries, with Chile again displaying a larger figure than the average in East Asia and Latin America. Trade union membership in developed economies is almost twice that in developing countries (0.39 versus 0.20, respectively). Trade union membership is lower in Latin America and East Asia than the mean for developing areas. The share of workers affiliated with trade unions is lower in Chile than in the group of Latin American countries, with Argentina and Brazil having the largest share of unionized workers. Finally, the size of public sector employment is larger in advanced economies than in developing countries; the difference is significantly larger when we use the general government (0.39 versus 0.22, respectively). Public employment in Chile is lower than average employment in both Latin America and East Asia, with the largest public employment in Latin America displayed by Argentina and Uruguay.

Finally, the table presents the simple average of the Djankov-La Porta indicators of labor regulation, which complement the measures of de jure regulations in the Rama-Artecona database. Regarding employment laws, developing countries are more regulated than industrial countries (1.60 versus 1.36), especially in the areas of job security and conditions of employment. The Chilean labor market is less regulated than the regional average (1.46 versus 1.79), as well as in job security and employment conditions. Similarly, developing countries are slightly more regulated than industrial countries in the area of industrial (collective) relations law (1.26 versus 1.22). Specifically, they are more regulated in the areas of collective bargaining and disputes and less regulated in the participation of workers in management. Argentina, Mexico, and Peru (not shown) have the most highly regulated labor markets in Latin America in the area of collective bargaining, followed by Chile and Colombia. Finally, workers in industrial countries are more protected with regard to social security than are workers in developing countries (2.21 versus 1.53); the largest difference is seen in unemployment benefits (0.78 versus 0.38).

Table 3 presents the evolution of labor regulations over the decades spanning the 1970–2000 period for different subsamples of countries. The aggregate index of de jure rigidities, $\mathbf{L}_0$, increased over the decades for all subgroups of countries. This implies that countries across the world ratified more ILO conventions over time. The extent of rigidities in practice decreased slightly among industrial countries in the 1990s relative to the 1980s, whereas it increased among developing countries. Chilean labor markets became more regulated in the 1990s, whether measured by the $\mathbf{L}_1$ or $\mathbf{L}_2$ index.

3.2 Correlation Analysis

Table 4 presents the correlation between economic growth and a wide array of labor regulation indicators for a cross-section of countries averaged over the 1970–2000 period. In the cross-correlation analysis between growth in per capita GDP and the indicators of labor market rigidity in the Rama-Artecona database, we find that growth and de facto rigidities ($\mathbf{L}_0$) are negatively correlated for the full sample (–0.12), with a stronger correlation among developing countries than industrial countries (–0.28 versus –0.12, respectively). The negative correlation between labor regulations and growth is strongest among East Asian countries (–0.54) and almost negligible in Latin America (–0.001).

The correlation between the $\mathbf{L}_1$ index of de facto labor regulations and economic growth is negative for the world sample (–0.06), as well as among industrial and developing countries (–0.24 and –0.12, respectively). The $\mathbf{L}_2$ index also yields a negative association between labor regulations and growth. In this case, the correlation is similar for both industrial and developing countries (fluctuating around 0.33). East Asia displays the strongest negative correlation (–0.83).

Economic growth is negatively associated with minimum wages among industrial countries (with a correlation above –0.30), and they are negatively associated among developing countries when normalized by per capita income (–0.20). The negative correlation between growth and mandated benefits is weak for the full sample of countries (–0.05 for social security contributions and –0.12 for maternity leave).[11] A larger share of trade union labor in the total labor force is associated with lower growth for developing and Latin American countries (–0.11 and –0.18, respectively). Finally, government employment has a positive correlation with growth among developing countries and a negative one among industrial economies and East Asia.

11. If we consider the contribution to social security, the correlation is positive and small for the group of industrial countries (0.06) and Latin America (0.09). Maternity leave has a negative correlation with growth for industrial and developing countries (–0.28 and –0.14, respectively), and a positive but negligible coefficient for Latin America.

Table 3. Basic Statistics for Labor Market Regulations and Economic Growth over the Decades[a]
Simple averages across groups of countries

	All countries			Industrial countries			Developing countries			Chile		
Variable	*1970s*	*1980s*	*1990s*	*1970s*	*1980s*	*1990s*	*1970s*	*1980s*	*1990s*	*1970s*	*1980s*	*1990s*
GDP per capita growth (percent)	2.36	1.23	1.42	2.49	2.19	2.12	2.32	0.94	1.20	1.20	1.27	4.78
De jure index $\mathbf{L}_0$	0.27	0.29	0.32	0.44	0.48	0.54	0.23	0.25	0.27	0.32	0.32	0.36
De facto index $\mathbf{L}_1$	0.27	0.27	0.28	0.36	0.37	0.36	0.24	0.25	0.26	0.15	0.17	0.20
Minimum wage[b]	0.23	0.22	0.23	0.25	0.23	0.22	0.22	0.21	0.23	0.12	0.12	0.19
Social security contribution	0.33	0.36	0.41	0.41	0.45	0.49	0.31	0.34	0.39	0.36	0.40	0.45
Trade union membership	0.24	0.25	0.23	0.39	0.41	0.37	0.19	0.21	0.19	0.09	0.09	0.13
General government employment	0.27	0.27	0.26	0.39	0.41	0.38	0.22	0.22	0.22	0.04	0.05	0.04
De facto index $\mathbf{L}_2$	0.28	0.29	0.30	0.31	0.32	0.31	0.27	0.27	0.29	0.06	0.06	0.11
Minimum wage[c]	0.14	0.14	0.13	0.10	0.09	0.09	0.17	0.16	0.15	0.05	0.06	0.08
Maternity leave (no. days)	0.14	0.15	0.17	0.18	0.19	0.20	0.13	0.14	0.16	0.15	0.15	0.23
Ratification of ILO convention 87	0.55	0.58	0.64	0.74	0.82	0.82	0.50	0.53	0.59	0.00	0.00	0.10
Central government employment	0.18	0.18	0.14	0.21	0.20	0.15	0.17	0.17	0.13	0.02	0.03	0.03
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.00	–0.02	–0.06	–0.06	–0.11	–0.18	0.01	0.00	–0.03	–0.16	–0.15	–0.16
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.01	–0.01	–0.04	–0.12	–0.16	–0.23	0.05	0.03	0.01	–0.26	–0.26	–0.26

Source: Authors' calculations, based on data from Rama and Artecona (2002) and Djankov and others (2003a).
a. Panel data of nonoverlapping five-year-average observations, 1970–2000. Indicators of labor market rigidity are from Rama and Artecona (2002).
b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.
c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.

Table 4. Cross-section Correlation Analysis between Labor Regulation and Economic Growth, 1970–2000

Variable	Full sample	Industrial economies	Developing countries	East Asia	Latin America
Labor market rigidity[a]					
De jure index $\mathbf{L}_0$	–0.12	–0.13	–0.28	0.00	–0.54
De facto index $\mathbf{L}_1$	–0.06	–0.24	–0.12	0.35	0.28
Minimum wage[b]	0.03	–0.32	0.11	0.51	–0.23
Social security contribution	–0.05	0.06	–0.11	0.09	0.18
Trade union membership	–0.04	0.01	–0.11	–0.18	0.50
General government employment	0.04	–0.31	0.00	0.15	0.08
De facto index $\mathbf{L}_2$	–0.31	–0.34	–0.33	0.11	–0.83
Minimum wage[c]	–0.23	–0.34	–0.20	0.56	–0.55
Maternity leave (no. days)	–0.12	–0.28	–0.14	0.02	0.22
Ratification of ILO convention 87	–0.31	–0.13	–0.37	–0.14	–0.52
Central government employment	0.23	–0.25	0.25	0.33	–0.15
De jure versus de facto					
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.07	–0.01	0.17	0.17	0.67
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	–0.11	–0.06	–0.05	0.08	–0.61
Labor regulation[d]					
Employment laws	–0.24	0.16	–0.28	–0.04	–0.14
Alternative employment contracts	–0.01	0.14	–0.04	0.13	0.21
Conditions of employment	–0.28	0.08	–0.30	–0.05	–0.07
Job security	–0.21	0.15	–0.23	–0.13	–0.42
Industrial (collective) relations law	–0.06	0.27	–0.11	–0.10	–0.28
Collective bargaining	–0.19	0.32	–0.25	–0.24	–0.35
Worker participation in management	0.14	0.20	0.13	0.05	0.04
Collective disputes	–0.09	0.07	–0.08	–0.09	–0.36
Social security laws	0.04	–0.20	–0.01	0.20	0.35
Old age, disability and death benefits	0.26	–0.23	0.28	0.24	0.06
Sickness and health benefits	0.05	–0.27	0.04	0.13	0.17
Unemployment benefits	–0.06	0.13	–0.14	0.10	0.45

Source: Authors' calculations, based on data from Rama and Artecona (2002) and Djankov and others (2003a).
a. Indicators of labor market rigidity are from Rama and Artecona (2002).
b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.
c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.
d. Indicators of labor regulations are from Djankov and others (2003a).

The table also shows the cross-section correlation analysis between economic growth and the indicators of labor market regulation from the Djankov-La Porta database. These variables describe the laws protecting workers in three main areas of the labor code: employment, industrial or collective relations, and social security. The aggregate index of employment laws is negatively associated with growth for the world sample (–0.24) and for developing countries (–0.28), although it is positive for industrial countries (0.16). Within the group of employment laws, the negative correlation for the full sample and for developing countries is strongest for conditions of employment (–0.28) and job security (–0.21).

Industrial relations (collective) laws, in turn, have a small and negative correlation with growth for the samples of Latin American and East Asian countries, whereas the correlation is positive for industrial countries. Laws on collective bargaining and on collective disputes have a negative correlation among developing areas, with the strongest correlation displayed for laws on collective bargaining (–0.25 for developing countries and –0.35 for East Asia). We find a positive association between growth and the participation of workers in management.

Finally, social security laws display a positive association with growth for the samples of Latin American and East Asian countries and a negative correlation for industrial countries. For the group of industrial countries, growth is negatively associated with laws contemplating old age, disability, and death benefits and with sickness and health benefits (–0.23 and –0.27, respectively), whereas there is a positive association between growth and unemployment benefits for the same group of countries (0.13). Developing countries showed a completely different correlation pattern: positive for old age and sickness benefits and negative for unemployment benefits.

Table 5 reports the results of our panel correlation analysis between economic growth and the labor regulation indicators in the Rama-Artecona database (the only one with a panel dimension). We present not only the panel correlation for the 1970–2000 period, but also the evolution of these correlation coefficients over the decades. In general, we find that de jure rigidities, $\mathbf{L}_0$, are negatively correlated with growth for all the samples. The correlation between growth and de facto rigidities is negative for industrial countries under both $\mathbf{L}_1$ and $\mathbf{L}_2$ and negative for developing countries under $\mathbf{L}_2$.

We also find that the degree of negative correlation between growth and $\mathbf{L}_0$ (de jure rigidities) increased in the 1990s relative to the 1980s for industrial countries (from –0.07 to –0.17), whereas it declined for developing countries over the same time period (from –0.25 to –0.10). Regarding the aggregate indices of de facto regulation, the negative correlation between $\mathbf{L}_1$ and growth decreased from –0.17 to –0.13 for industrial countries, while it remained constant for $\mathbf{L}_2$ at around –0.28. For developing countries, the correlation between $\mathbf{L}_1$ and growth became negative in the 1990s (–0.10) after being slightly positive in the 1980s (0.08), and it remained unchanged for $\mathbf{L}_2$ over the same period (–0.27).

Table 5. Panel Data Correlation Analysis between Labor Market Regulation and Economic Growth, 1970–2000[a]

Labor rigidity indicator	*All countries*				*Industrial countries*				*Developing countries*			
	Pooled	*1970s*	*1980s*	*1990s*	*Pooled*	*1970s*	*1980s*	*1990s*	*Pooled*	*1970s*	*1980s*	*1990s*
De jure index $\mathbf{L}_0$	–0.06	–0.05	–0.08	–0.03	–0.11	0.07	–0.07	–0.17	–0.15	–0.08	–0.25	–0.10
De facto index $\mathbf{L}_1$	0.04	0.14	0.12	–0.07	–0.15	–0.14	–0.17	–0.13	0.03	0.18	0.08	–0.10
Minimum wage[b]	0.07	0.02	0.09	0.10	–0.16	–0.13	–0.24	–0.15	0.14	0.06	0.18	0.17
Social security contribution	–0.01	0.09	–0.06	–0.01	0.03	0.23	–0.03	0.01	–0.04	0.08	–0.13	–0.03
Trade union membership	0.07	0.15	0.19	–0.09	0.00	–0.11	0.05	0.01	0.04	0.19	0.16	–0.16
General government employment	0.04	0.15	0.06	–0.05	–0.22	–0.36	–0.15	–0.20	0.02	0.26	–0.04	–0.09
De facto index $\mathbf{L}_2$	–0.20	–0.09	–0.25	–0.26	–0.18	0.08	–0.28	–0.28	–0.22	–0.11	–0.27	–0.27
Minimum wage[c]	–0.15	–0.21	–0.16	–0.10	–0.17	–0.05	–0.40	–0.09	–0.13	–0.23	–0.09	–0.08
Maternity leave (no. days)	–0.04	0.02	–0.04	–0.08	–0.17	–0.24	–0.06	–0.20	–0.05	0.05	–0.14	–0.08
Ratification of ILO convention 87	–0.18	–0.08	–0.25	–0.20	–0.07	0.24	–0.16	–0.20	–0.22	–0.11	–0.32	–0.23
Central government employment	0.09	0.15	0.03	0.07	–0.14	–0.25	–0.12	–0.15	0.12	0.23	0.02	0.10
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.09	0.12	0.18	–0.03	0.02	–0.18	–0.03	0.09	0.16	0.19	0.33	0.00
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	–0.09	0.03	–0.10	–0.18	–0.01	–0.08	–0.09	0.03	–0.06	0.04	–0.02	–0.18

Source: Authors' calculations, based on data from Rama and Artecona (2002).

a. Panel data for the 1970–2000 period are in five-year nonoverlapping observations. Indicators of labor market rigidity are from Rama and Artecona (2002).

b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.

c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.

Correlation Among Indicators of Labor Market Regulation

Here we briefly present the correlations between the different indicators of labor market regulation used in our regression analysis. First, we correlate labor indicators within the Rama-Artecona and Djankov-La Porta databases. We then correlate different indicators between the two databases. We find that countries with a higher degree of regulation on paper also display a higher degree of regulations in practice. This is reflected in the positive association between the index $\mathbf{L}_0$ and the aggregate indices of de facto rigidities, $\mathbf{L}_1$ and $\mathbf{L}_2$ (with correlation coefficients of 0.53 and 0.44). On the other hand, both aggregate indices, $\mathbf{L}_1$ and $\mathbf{L}_2$, are positively correlated (0.44).

Next, we analyze the correlation between each aggregate index of de facto regulations and their components. For the L_1 index, the proxies of trade unions and government employment have the highest correlation with the aggregate index (approximately 0.78), while minimum wages has the weakest correlation (0.44). In the case of the L_2 index, the trade union indicator displays the highest correlation with the aggregate index (0.92), while the correlation of the remaining dimensions fluctuates between 0.30 and 0.34. The proxies used in each dimension of the aggregate indices, $\mathbf{L}_1$ and $\mathbf{L}_2$, are positively correlated, with employment in general and the central government having a degree of correlation of 0.55. The correlation between minimum wage indicators is 0.35; between measures of mandated benefits, 0.29; and between trade union variables, 0.30.

We also report correlations for the labor regulation measures in the Djankov-La Porta database. We find that countries with higher regulation in employment laws also display a larger degree of regulations in industrial relations and social security laws. The positive correlation is strongest between employment laws and industrial relation laws (0.52) and weakest between employment and social security laws (0.10). We also find that the components of each aggregate index proposed by the Djankov-La Porta database are highly correlated with the aggregate index. For example, the aggregate index of social security laws is highly correlated with laws on sickness and health benefits (0.84) and unemployment benefits (0.89), while the aggregate index of employment laws is highly correlated with laws on job security (0.81) and employment conditions (0.79).

Finally, we evaluate the correlation of labor regulation indicators between databases. First, we find that the index of de jure regulations, $\mathbf{L}_0$, in the Rama-Artecona database is positively associated with the aggregate indices in the Djankov-La Porta database. The highest correlation is displayed between $\mathbf{L}_0$ and social security laws (0.46), while the lowest is between $\mathbf{L}_0$ and employment laws (0.16). Analogously, we find that either the $\mathbf{L}_1$ and $\mathbf{L}_2$ index of aggregate de facto regulations is positively correlated with the indices in the Djankov-La Porta data. The highest correlation is again displayed with social security laws (0.59 with the $\mathbf{L}_1$ index and 0.30 with the $\mathbf{L}_2$ index).

3.3 Cross-section Regression Analysis

This section discusses the results for the relation between labor market regulations and economic growth for a cross-section of countries. Our dependent variable is the annual average growth rate in GDP per capita over the 1970–2000 period. The explanatory variables are the log of per capita GDP in 1970, the average years of secondary schooling in 1970, the ratio of domestic credit to the private sector to GDP, the average annual inflation rate, the degree of openness, the average annual change in the terms of trade, the real exchange rate overvaluation, the index of civil liberties (as a proxy for governance), and our measures of labor regulations. For reasons of space we only report the coefficient of interest (namely, the labor regulation coefficient), its standard deviation, the coefficient of determination (R^2), and the number of observations.[12]

Results for the Rama-Artecona Labor Regulation Indicators

Table 6 reports the estimated coefficient of labor regulation measures in the Rama-Artecona database and its statistical significance for the sample of all countries and the samples of developing and industrial countries. We provide both OLS and IV estimations for these coefficients.[13] The OLS estimates indicate that aggregate measures of labor regulations—de jure and de facto—are negatively and significantly related to growth among industrial countries. The sample of developing countries, as well as the full sample of countries, yield a negative association (although statistically negligible) between both growth and de jure regulations and growth and the $\mathbf{L}_2$ index of de facto regulations. Also, while the R^2 coefficient fluctuates between 0.44 and 0.73 for the full sample of countries and for developing countries, it ranges from 0.82 to 0.91 for the sample of industrial countries.

12. The full regression results are available on request.

13. The coefficient estimates and standard errors of the OLS estimates are robust to autocorrelation and White heteroskedasticity (following White, 1980).

Turning to the components of these aggregate indices of de facto labor regulations, we find the following significant results. First, public employment (by either the central or the general government) as a share of total employment has a positive association with economic growth for the full sample and for developing countries. On the other hand, trade unions fully explain the negative and significant correlation between growth and the $\mathbf{L}_2$ index for these two samples. Second, the negative relation between growth and the index of de facto labor regulations—whether measured by $\mathbf{L}_1$ or $\mathbf{L}_2$—among industrial countries is mainly driven by minimum wages and mandated benefits, proxied by either social security contributions or days of maternity leave (that is, the economic dimension of labor market regulations, according to Forteza and Rama, 2002).

For the IV estimates presented in the table, we instrument for labor market regulations following Djankov and others (2003a), as mentioned earlier. We find that labor markets are more regulated on paper ($\mathbf{L}_0$) and in practice ($\mathbf{L}_1$ and $\mathbf{L}_2$) in richer countries and in countries with left-oriented governments. In contrast, the extent of regulation is lower in countries with common law tradition.[14]

We find, first, that de jure regulations have a negative impact on long-run growth for all samples, although it is statistically significant only for industrial countries. Hence, if regulations in the Spanish labor code (which was the most highly regulated in the OECD during the 1995–99 period) were reduced to the average levels (namely, those exhibited by Greece and Portugal), the country's growth rate would increase by 1 percentage point per year. Second, although the $\mathbf{L}_1$ and $\mathbf{L}_2$ indices of de facto regulations have a negative relation with growth in all samples, $\mathbf{L}_1$ exerts a negative and significant impact on growth among developing countries, whereas $\mathbf{L}_2$ has a negative impact on growth among industrial countries.[15] Finally, the minimum wage has a negative and robust relation in the IV estimations regarding the normalization factor and the samples of countries.

Economically speaking, our IV cross-sectional estimates suggest two key implications. First, if Sweden (the country with the highest degree of regulation in 1995–99 according to this index) reduced its labor rigidities to Switzerland's level (the representative country in

14. For the sake of brevity, we do not report the first-stage regression results; they are available on request.

15. The negative impact of $\mathbf{L}_1$ on growth is mainly attributed to minimum wages, mandated benefits, and public employment, while the negative impact of $\mathbf{L}_2$ on growth is explained by minimum wages and trade union membership.

Table 6. Cross-section Regression Analysis for Labor Market Regulations and Economic Growth[a]

Estimation method and labor indicator	*All countries* Coeff.	Std. dev.	R^2	No. obs.	*Industrial countries* Coeff.	Std. dev.	R^2	No. obs.	*Developing countries* Coeff.	Std. dev.	R^2	No. obs.
Least squares[b]												
De jure index $\mathbf{L}_0$	–0.005	0.01	0.49	76	–0.012**	0.00	0.88	22	–0.010	0.02	0.50	54
De facto index $\mathbf{L}_1$	0.019	0.01	0.47	75	–0.017**	0.01	0.86	22	0.055**	0.02	0.51	53
Minimum wage[c]	0.000	0.01	0.47	65	–0.007**	0.00	0.87	22	0.015	0.01	0.52	43
Social security contribution	0.003	0.01	0.55	53	–0.010**	0.00	0.91	18	–0.005	0.02	0.61	35
Trade union membership	0.004	0.01	0.55	53	0.006	0.00	0.92	18	–0.011	0.03	0.59	35
General government employment	0.018*	0.01	0.44	67	–0.006	0.01	0.86	22	0.038**	0.02	0.50	45
De facto index $\mathbf{L}_2$	–0.015	0.01	0.48	73	–0.022**	0.01	0.86	22	–0.013	0.01	0.50	51
Minimum wage[d]	–0.014	0.01	0.49	66	–0.016**	0.01	0.87	22	–0.027	0.02	0.53	44
Maternity leave (no. days)	–0.012	0.02	0.52	59	–0.016**	0.01	0.87	21	–0.070	0.07	0.56	38
Ratification of ILO convention 87	–0.009*	0.01	0.53	59	0.000	0.00	0.94	21	–0.018**	0.01	0.59	38
Central government employment	0.027*	0.01	0.55	66	–0.019**	0.01	0.88	21	0.051	0.02	0.58	45
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.013*	0.01	0.73	75	0.006*	0.00	0.82	22	0.032	0.01	0.51	53
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	–0.004	0.01	0.47	73	0.012**	0.00	0.85	22	–0.005	0.02	0.49	51
Instrumental variables[e]												
De jure index $\mathbf{L}_0$	–0.027	0.03	0.49	76	–0.026*	0.02	0.83	22	–0.042	0.04	0.51	54
De facto index $\mathbf{L}_1$	–0.042	0.03	0.48	75	–0.022	0.02	0.82	22	–0.077*	0.05	0.51	53
Minimum wage[c]	–0.081**	0.03	0.53	65	–0.043**	0.02	0.83	22	–0.102**	0.04	0.58	43
Social security contribution	–0.006	0.02	0.55	53	–0.041**	0.02	0.89	18	–0.013	0.03	0.61	35
Trade union membership	–0.049	0.04	0.57	53	0.018	0.02	0.91	18	–0.078*	0.04	0.62	35
General government employment	–0.037	0.04	0.43	67	0.086**	0.02	0.88	22	–0.120**	0.05	0.49	45
De facto index $\mathbf{L}_2$	0.013	0.04	0.47	73	–0.036*	0.02	0.83	22	–0.005	0.05	0.49	51
Minimum wage[d]	–0.165**	0.07	0.51	66	–0.071**	0.03	0.86	22	–0.096	0.10	0.53	44
Maternity leave (no. days)	–0.075	0.08	0.53	59	–0.044	0.03	0.84	21	–0.214	0.17	0.54	38
Ratification of ILO convention 87	0.026	0.03	0.52	59	–0.026**	0.01	0.87	21	–0.015	0.03	0.50	38
Central government employment	0.036	0.06	0.51	66	0.048**	0.02	0.87	21	0.007	0.08	0.52	45

Table 6. (continued)

Estimation method and labor indicator	*All countries*				*Industrial countries*				*Developing countries*			
	Coeff.	*Std. dev.*	R^2	*No. obs.*	*Coeff.*	*Std. dev.*	R^2	*No. obs.*	*Coeff.*	*Std. dev.*	R^2	*No. obs.*
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.022	0.05	0.46	75	0.064**	0.03	0.85	22	0.002	0.06	0.47	53
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.106**	0.05	0.52	73	0.042*	0.03	0.82	22	0.177**	0.07	0.55	51

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002) and Djankov and others (2003a).
a. We report the regression coefficient for the indicator of labor rigidity according to equation (1) in the text, based on an effective sample of seventy-six countries averaged over the 1970–2000 period. Our control variables are output per capita (in logs), secondary schooling, domestic credit to the private sector, trade openness, governance, inflation, real exchange rate overvaluation, terms-of-trade shocks, and the labor regulation indicator. Labor regulation data are from Rama and Artecona (2002). Full regression results and standard errors of the coefficients of the labor regulation variables are not reported for reasons of space, although they are available from the authors on request.
b. Standard errors are robust to autocorrelation and heteroskedasticity (White, 1980).
c. Minimum wages are normalized with the average labor cost in the manufacturing sectors.
d. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.
e. Our set of instruments for the labor indicators consists of the level of development, trade openness adjusted by geographic variables, leftist political orientation of the government, common law tradition, German civil code tradition, and institutionalized autocracy. The set of instruments was chosen from the existing literature, following Djankov and others (2003a).

the sample of advanced economies), then its growth rate would increase by 0.7 percentage points per year. Second, if Argentina (the country with the highest degree of labor regulations in Latin America in 1995–99) reduced its level of regulations to the regional average, its growth rate would increase by 1.2 percentage points. If Argentina reduced its degree of labor regulation to the East Asian average, its growth rate would increase by 1.6 percentage points.

Results for the Djankov-La Porta Labor Regulation Indicators

Table 7 provides the coefficient estimates for labor regulation indicators in the Djankov-La Porta database, for which we run the same experiments as in the previous table.[16] The least squares estimates of the labor regulation coefficients for different samples of countries show that the aggregate index of employment laws has a negative and significant relation with growth among industrial countries, which is explained by laws on employment conditions. The index of social security laws has a positive and significant relation with growth in the world sample and the sample of developing countries, although the quantitative relevance of this estimated relation seems to be non significant.

With regard to the estimated IV coefficients for the labor regulation indicators, all three aggregate types of labor laws (namely, employment, industrial relations, and social security) have a negligible impact on growth for the world sample and for developing countries. However, all three aggregate indices have a negative and significant impact on growth among industrial countries. In the case of employment laws, for example, we find that if Portugal (the country with the strictest regulations in 1997) were to reduce its regulations to the level of Austria (a country with the average level of regulations), its growth rate would increase by 0.6 percentage points. An analogous decline in job security (from the countries with the highest levels to the average) might improve growth by almost 3 percentage points. In the case of industrial relations laws, Portugal could improve its growth rate by 0.6 percentage points if it reduced its degree of regulation to the average levels in the region (for example, that of the Netherlands). Finally, if social security regulations were lessened in Denmark and Sweden (that is, the countries with the region's most extensive regulations) to a

16. The full specification is analogous to that used in the cross-section analysis for the Rama-Artecona database and is available on request.

level on par with the regional average (for example, Switzerland and Italy), their growth rate would increase by 0.6 percentage points.

3.4 Panel Data Regression Analysis

We now present the panel data estimates of the relation between labor market rigidities and economic growth. We use panel data on seventy-six countries with nonoverlapping five-year-average observations for the 1970–2000 period. Here, we report three types of estimators: least-squares-based estimators (pooled OLS, least squares with time effects, and the within-group estimator); instrumental variables estimators, in which we instrument for labor market regulations following the strategy outlined earlier, both with pooled and time / country effects IV estimators; and generalized method of moments estimators (Arellano and Bond, 1991; Arellano and Bover, 1995), in which we control for unobserved country and time effects and the possibility of endogenous regressors and in which we use both internal instruments (that is, lagged levels of the variables in our regression framework) and external instruments (that is, exogenous variables that determine the choice of labor institutions and regulations in the country). For reasons of space, we only briefly explain the OLS and IV results and then focus our discussion on the GMM result, which is our preferred estimation method.

Panel Results from Least-square-based Estimators

Table 8 contains the regression results for the estimated coefficients of the wide array of labor regulation indicators using different least-squares-based techniques (namely, pooled OLS, least squares with time effects, and least squares with country dummy variables) applied to the full sample, the industrial country sample, and the developing country sample. Two caveats apply: these estimation techniques do not address the possibility of endogenous regressors, and taking into account unobserved country effects through country dummy variables (as in the within-group estimator) in a dynamic panel data model leads to inconsistent estimates.

Our pooled OLS estimates indicate that both the $\mathbf{L}_0$ index of de jure rigidities and the L_2 index of de facto rigidities have a negative and significant relationship with economic growth. The impact of $\mathbf{L_1}$ on growth is negative and significant only for industrial countries. These results for the indices of de facto regulations hold for the time-effects

Table 7. Cross-section Regression Analysis for Labor Market Regulations and Economic Growth[a]

Estimation method and labor indicator	*All countries*				*Industrial countries*				*Developing countries*			
	Coeff.	*Std. dev.*	R^2	*No. obs.*	*Coeff.*	*Std. dev.*	R^2	*No. obs.*	*Coeff.*	*Std. dev.*	R^2	*No. obs.*
Least squares[b]												
Employment laws	0.001	0.00	0.50	58	–0.004*	0.00	0.84	21	–0.002	0.01	0.53	37
Alternative employment contracts	0.009	0.01	0.51	58	0.005	0.00	0.87	21	0.007	0.01	0.54	37
Conditions of employment	0.003	0.01	0.50	58	–0.009**	0.00	0.85	21	0.008	0.01	0.54	37
Job security	–0.010	0.01	0.52	58	–0.007	0.01	0.84	21	–0.017*	0.01	0.56	37
Industrial (collective) relations law	0.001	0.00	0.50	58	–0.003*	0.00	0.84	21	–0.003	0.01	0.53	37
Collective bargaining	0.000	0.01	0.50	58	0.001	0.00	0.85	21	–0.004	0.01	0.53	37
Worker participation in management	0.003	0.00	0.50	58	–0.004	0.00	0.84	21	0.001	0.01	0.54	37
Collective disputes	–0.003	0.01	0.50	58	–0.010	0.01	0.85	21	–0.007	0.01	0.54	37
Social security laws	0.008**	0.00	0.55	58	0.002	0.00	0.81	21	0.007**	0.00	0.56	37
Old age, disability, and death benefits	0.016	0.01	0.55	58	0.000	0.01	0.81	21	0.030*	0.02	0.58	37
Sickness and health benefits	0.003	0.01	0.56	58	0.004	0.00	0.81	21	0.002	0.01	0.57	37
Unemployment benefits	0.013**	0.01	0.56	58	–0.001	0.01	0.81	21	0.010	0.01	0.56	37
Instrumental variables[c]												
Employment laws	0.004	0.01	0.51	58	–0.007*	0.00	0.85	21	0.002	0.01	0.53	37
Alternative employment contracts	0.027	0.03	0.51	58	0.044**	0.02	0.87	21	–0.035	0.08	0.54	37
Conditions of employment	0.018	0.02	0.51	58	–0.029**	0.01	0.86	21	0.089	0.09	0.55	37
Job security	–0.040	0.06	0.51	58	–0.073**	0.02	0.88	21	–0.009	0.09	0.53	37
Industrial (collective) relations law	0.004	0.01	0.51	58	–0.006**	0.00	0.86	21	0.005	0.01	0.54	37
Collective bargaining	0.042**	0.02	0.54	58	–0.010	0.01	0.86	21	0.061**	0.02	0.59	37
Worker participation in management	–0.030	0.02	0.54	58	–0.018**	0.01	0.86	21	–0.032	0.02	0.57	37
Collective disputes	0.047	0.03	0.52	58	0.018*	0.01	0.87	21	0.025	0.03	0.55	37

Table 7. (continued)

Estimation method and labor indicator	All countries				Industrial countries				Developing countries			
	Coeff.	Std. dev.	R^2	No. obs.	Coeff.	Std. dev.	R^2	No. obs.	Coeff.	Std. dev.	R^2	No. obs.
Social security laws												
Old age, disability, and death benefits	–0.034	0.07	0.54	58	–0.131**	0.04	0.88	21	–0.088	0.10	0.54	37
Sickness and health benefits	–0.040	0.04	0.52	58	–0.031**	0.01	0.88	21	–0.066	0.06	0.56	37
Unemployment benefits	0.033	0.03	0.51	58	–0.043**	0.01	0.90	21	0.056	0.04	0.56	37

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002) and Djankov and others (2003a).
a. We report the regression coefficient for the indicator of labor rigidity according to equation (1) in the text, based on an effective sample of seventy-six countries averaged over the 1970–2000 period. Our control variables are output per capita (in logs), secondary schooling, domestic credit to the private sector, trade openness, governance, inflation, real exchange rate overvaluation, terms-of-trade shocks, and the labor regulation indicator. Labor regulation data are from Djankov and others (2003a). Full regression results and standard errors of the coefficients of the labor regulation variables are not reported for reasons of space, although they are available from the authors on request.
b. Standard errors are robust to autocorrelation and heteroskedasticity (White, 1980).
c. Our set of instruments for the labor indicators consists of the level of development, trade openness adjusted by geographic variables, leftist political orientation of the government, common law tradition, German civil code tradition, and institutionalized autocracy. The set of instruments was chosen from the existing literature, following Djankov and others (2003a).

Table 8. Panel Data Regression Analysis for Labor Market Regulations and Economic Growth: Least Squares[a]

Estimation method and labor indicator	Full sample Coeff.	Std. dev.	R^2	No. obs.	Industrial countries Coeff.	Std. dev.	R^2	No. obs.	Developing countries Coeff.	Std. dev.	R^2	No. obs.
Pooled estimator												
De jure index $\mathbf{L}_0$	–0.010*	0.01	0.20	382	–0.014**	0.01	0.29	111	–0.021*	0.01	0.20	271
De facto index $\mathbf{L}_1$	0.003	0.01	0.19	399	–0.015*	0.01	0.29	120	0.020	0.02	0.19	279
Minimum wage[b]	0.006	0.01	0.22	349	–0.008*	0.00	0.30	120	0.022*	0.01	0.24	229
Social security contribution	0.001	0.01	0.21	355	–0.006	0.01	0.30	105	0.000	0.01	0.21	250
Trade union membership	–0.001	0.01	0.22	366	0.005	0.01	0.34	119	–0.016	0.02	0.23	247
General government employment	–0.006	0.01	0.19	333	–0.002	0.01	0.29	120	0.004	0.01	0.20	213
De facto index $\mathbf{L}_2$	–0.024**	0.01	0.21	393	–0.024*	0.01	0.30	120	–0.025**	0.01	0.21	273
Minimum wage[c]	–0.011	0.01	0.24	358	–0.021**	0.01	0.31	120	–0.005	0.01	0.25	238
Maternity leave (no. days)	0.000	0.01	0.21	364	0.009	0.01	0.32	117	–0.018	0.04	0.21	247
Ratification of ILO convention 87	–0.012*	0.01	0.22	387	0.019**	0.01	0.33	120	–0.024**	0.01	0.23	267
Central government employment	0.011	0.01	0.22	335	–0.016*	0.01	0.34	119	0.029*	0.02	0.24	216
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.012*	0.01	0.20	377	0.008	0.01	0.27	111	0.031**	0.01	0.21	266
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	–0.002	0.01	0.19	370	0.008	0.01	0.27	111	0.001	0.01	0.20	259
Time-effects estimator												
De jure index $\mathbf{L}_0$	–0.007	0.01	0.24	382	–0.011**	0.01	0.46	111	–0.015	0.01	0.24	271
De facto index $\mathbf{L}_1$	–0.001	0.01	0.24	399	–0.013*	0.01	0.46	120	0.012	0.02	0.23	279
Minimum wage[b]	0.005	0.01	0.29	349	–0.008*	0.01	0.47	120	0.019*	0.01	0.30	229
Social security contribution	0.003	0.01	0.27	355	–0.004	0.01	0.47	105	0.004	0.01	0.26	250
Trade union membership	–0.007	0.01	0.28	366	0.007	0.01	0.50	119	–0.028*	0.02	0.28	247
General government employment	–0.007	0.01	0.24	333	–0.003	0.01	0.46	120	0.001	0.01	0.24	213
De facto index $\mathbf{L}_2$	–0.021**	0.01	0.25	393	–0.024**	0.01	0.47	120	–0.022**	0.01	0.25	273
Minimum wage[c]	0.001	0.01	0.30	358	–0.021**	0.01	0.48	120	0.012	0.02	0.30	238
Maternity leave (no. days)	0.002	0.02	0.26	364	0.010	0.01	0.49	117	0.001	0.04	0.25	247
Ratification of ILO convention 87	–0.012*	0.01	0.27	387	0.019**	0.01	0.50	120	–0.024**	0.01	0.28	267
Central government employment	0.006	0.01	0.27	335	–0.017*	0.01	0.51	119	0.024	0.02	0.28	216

Table 8. (continued)

Estimation method and labor indicator	Full sample Coeff.	Std. dev.	R^2	No. obs.	Industrial countries Coeff.	Std. dev.	R^2	No. obs.	Developing countries Coeff.	Std. dev.	R^2	No. obs.
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.007	0.01	0.24	377	0.007	0.01	0.44	111	0.022*	0.01	0.25	266
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	–0.004	0.01	0.24	370	0.007	0.01	0.44	111	–0.002	0.01	0.25	259
Country-effects estimator												
De jure index $\mathbf{L}_0$	0.017	0.03	0.54	382	–0.033	0.03	0.57	111	0.017	0.04	0.55	271
De facto index $\mathbf{L}_1$	–0.013	0.03	0.54	399	–0.010	0.03	0.52	120	–0.014	0.04	0.55	279
Minimum wage[b]	0.014	0.02	0.57	349	–0.032*	0.02	0.53	120	0.024	0.02	0.58	229
Social security contribution	0.030	0.02	0.57	355	0.044*	0.03	0.54	105	0.025	0.03	0.57	250
Trade union membership	0.001	0.02	0.55	366	0.020	0.02	0.52	119	–0.003	0.03	0.56	247
General government employment	–0.040**	0.02	0.59	333	–0.034*	0.02	0.54	120	–0.053*	0.03	0.60	213
De facto index $\mathbf{L}_2$	–0.010	0.03	0.54	393	–0.025	0.03	0.52	120	0.003	0.04	0.55	273
Minimum wage[c]	–0.049	0.03	0.56	358	–0.074*	0.05	0.53	120	–0.057	0.05	0.57	238
Maternity leave (no. days)	0.051	0.04	0.57	364	0.012	0.04	0.53	117	0.099	0.08	0.58	247
Ratification of ILO convention 87	0.022*	0.01	0.56	387	0.019	0.02	0.53	120	0.024	0.02	0.57	267
Central government employment	–0.024	0.02	0.59	335	–0.011	0.02	0.51	119	–0.049	0.05	0.60	216
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	–0.019	0.02	0.54	377	0.016	0.03	0.57	111	–0.013	0.03	0.55	266
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	–0.005	0.02	0.55	370	0.000	0.03	0.57	111	0.003	0.03	0.56	259

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002) and Djankov and others (2003a).
a. We report the regression coefficient for the indicator of labor rigidity according to equation (1) in the text, based on an effective sample of seventy-six countries averaged over the 1970–2000 period. The estimation method is least squares. The dependent variable is the growth rate in per capita GDP. Our control variables are output per capita (in logs), secondary schooling, domestic credit to the private sector, trade openness, governance, inflation, real exchange rate overvaluation, terms-of-trade shocks, and the labor regulation indicator. Labor regulation data are from Rama and Artecona (2002). Full regression results and standard errors of the coefficients of the labor regulation variables are not reported for reasons of space, although they are available on request. Asymptotic standard errors robust to general cross-section and time-series heteroskedasticity are reported.
b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.
c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.

estimation. In the case of $\mathbf{L}_2$ index, the negative association between growth and labor regulations among industrial countries is explained by minimum wages and public employment, whereas it is explained by trade union membership for developing countries. When we account for country effects, we find that the aggregate indices of labor regulations—whether on paper or in practice—have no significant relation with growth for all samples.[17] Here, only the minimum wage has a negative and significant association with growth among developing countries, while general government employment has a negative and significant link with growth regardless of the sample.

Panel Results from Instrumental Variables Estimators

In table 9, we present the coefficient estimates for a large set of labor market regulation indicators using IV techniques. We performed pooled IV regressions, IV with time effects, and IV with country effects on the full sample of countries, as well as on the samples of industrial and developing countries. Our set of instruments for labor regulation indicators includes per capita output (in logs), trade openness, the government's orientation to the left, common law tradition, German civil code tradition, and institutionalized autocracy. The main results for the first-stage panel regressions are that rich countries and countries with a left-leaning political orientation have a higher propensity to impose labor rigidities and regulations than poor or conservative countries. Also, fewer regulations would be imposed in more open countries, in countries with common law tradition, and in less autocratic governments.

Our pooled and time-effects IV estimates yield similar qualitative results. The $\mathbf{L}_0$ index of de jure regulations and the $\mathbf{L}_1$ index of de facto regulations have a negative and significant impact on growth for the full sample of countries and the sample of developing nations, and all the components of the $\mathbf{L}_1$ index have a negative and significant effect on growth.[18] The $\mathbf{L}_2$ index of regulations, however, has a negative

17. This estimation is consistent in a dynamic panel data setting model only if the time dimension is very large (Nickell, 1981). These results should thus be interpreted very cautiously.

18. Based on our IV estimates with time effects, we find that if regulations in Mexican labor markets (which have the highest adjusted degree of labor regulations using the $\mathbf{L}_1$ index) declined to average Latin American levels (such as those of Colombia and Paraguay), the country's growth rate would increase by 1.1 percentage points. If labor regulations in Mexico declined to average East Asian levels, the gains in economic growth would be even higher (approximately 1.8 percentage points). In this latter case, the growth effects of reducing the extent of regulations are larger for minimum wages (2.4 percentage points) and for public sector employment (approximately 3.0 percentage points).

(though insignificant) impact on growth in all samples. Finally, our country-effects estimates yield an insignificant statistical effect of both de jure and de facto labor regulations on economic growth, although the minimum wage remains negative and significant when we account for country-effects. As mentioned earlier, these results should be taken with caution since they do not properly account for the presence of unobserved country-specific effects.

The GMM-IV System Estimator

Having characterized the link between economic growth and labor regulations using some conventional panel data estimation techniques, we now use the GMM-IV system estimator for dynamic panel data proposed by Arellano and Bover (1995) and Blundell and Bond (1998). The reasons behind the application of this methodology are threefold: we need an estimator that deals properly with the dynamic nature of our model; we need to account for unobserved country-specific effects within the framework of a dynamic panel data model; and we need to control for the possibility of endogenous regressors. One of the advantages of this estimation technique is that we can compute some specification tests to confirm whether our growth regressions are valid for statistical inference. Further statistical details on the estimation technique are included in appendix B.

Table 10 presents the regression results of our growth equation using the GMM-IV system estimator. The main difference with respect to the IV estimator used above is that we use not only the economic, legal, and political determinants of labor regulations, but also internal instruments (that is, lagged levels or differences of the explanatory variables) to account for the endogenous explanatory variables. Our instruments are valid according to the Sargan test, and we reject the possibility that the error terms display high-order serial correlation.[19] Among the main results for our control variables we find evidence of convergence for the full sample of countries. We also find that growth is enhanced by larger stocks of human capital, better governance, lower inflation, and real exchange rate overvaluation, as well as positive terms-of-trade shocks. Coefficient estimates of credit to the private sector and openness either are not robust or display an unexpected sign (see table 10). In the following paragraphs, we evaluate the significance of the impact of our variable of interest, that is, the effect of labor market regulations.

19. By construction, the error process should always exhibit first-order linear correlation (Arellano and Bover, 1995).

Table 9. Panel Data Regression Analysis for Labor Market Regulations and Economic Growth: Instrumental Variables[a]

Estimation method and labor indicator	*Full sample*				*Industrial countries*				*Developing countries*			
	Coeff.	*Std. dev.*	R^2	*No. obs.*	*Coeff.*	*Std. dev.*	R^2	*No. obs.*	*Coeff.*	*Std. dev.*	R^2	*No. obs.*
Pooled estimator												
De jure index $\mathbf{L}_0$	–0.034**	0.02	0.21	381	–0.005	0.02	0.26	111	–0.052**	0.03	0.21	270
De facto index $\mathbf{L}_1$	–0.064**	0.02	0.21	398	0.023	0.02	0.29	120	–0.136**	0.04	0.23	278
Minimum wage[b]	–0.078**	0.02	0.24	349	0.023	0.02	0.29	120	–0.178**	0.04	0.28	229
Social security contribution	–0.032**	0.01	0.22	355	–0.019	0.02	0.30	105	–0.048**	0.02	0.23	250
Trade union membership	–0.068**	0.03	0.23	366	0.088**	0.02	0.39	119	–0.139**	0.04	0.25	247
General government employment	–0.092**	0.03	0.21	333	0.068**	0.02	0.34	120	–0.274**	0.05	0.28	213
De facto index $\mathbf{L}_2$	–0.036	0.02	0.20	393	–0.011	0.03	0.28	120	–0.056*	0.03	0.20	273
Minimum wage[c]	–0.099**	0.05	0.25	358	–0.030	0.06	0.30	120	–0.146**	0.06	0.27	238
Maternity leave (no. days)	–0.093*	0.05	0.22	364	0.115**	0.05	0.35	117	–0.188**	0.07	0.23	247
Ratification of ILO convention 87	–0.008	0.01	0.21	387	–0.008	0.01	0.30	120	–0.011	0.01	0.22	267
Central government employment	–0.087	0.06	0.22	335	0.143**	0.05	0.37	119	–0.251**	0.10	0.26	216
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.052*	0.03	0.20	376	0.080*	0.05	0.30	111	0.051	0.04	0.21	265
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.106**	0.03	0.22	370	0.003	0.03	0.26	111	0.198**	0.05	0.24	259
Time-effects estimator												
De jure index $\mathbf{L}_0$	–0.030*	0.02	0.25	381	–0.007	0.02	0.44	111	–0.045*	0.03	0.25	270
De facto index $\mathbf{L}_1$	–0.064**	0.02	0.26	398	0.014	0.02	0.45	120	–0.131**	0.04	0.27	278
Minimum wage[b]	–0.075**	0.02	0.30	349	0.014	0.02	0.45	120	–0.166**	0.04	0.33	229
Social security contribution	–0.029*	0.02	0.27	355	0.004	0.02	0.47	105	–0.043**	0.02	0.27	250
Trade union membership	–0.074**	0.03	0.29	366	0.074**	0.03	0.53	119	–0.143**	0.05	0.30	247
General government employment	–0.096**	0.03	0.26	333	0.059**	0.02	0.49	120	–0.281**	0.06	0.32	213
De facto index $\mathbf{L}_2$	–0.028	0.03	0.24	393	–0.008	0.02	0.45	120	–0.040	0.04	0.24	273
Minimum wage[c]	–0.092**	0.05	0.31	358	–0.193**	0.09	0.47	120	–0.128*	0.07	0.31	238
Maternity leave (no. days)	–0.103*	0.05	0.27	364	0.093	0.05	0.51	117	–0.176**	0.08	0.27	247
Ratification of ILO convention 87	–0.005	0.01	0.27	387	–0.006	0.01	0.47	120	–0.005	0.02	0.26	267
Central government employment	–0.098*	0.06	0.28	335	–0.035*	0.02	0.53	119	–0.253**	0.09	0.30	216

Table 9. (continued)

Estimation method and labor indicator	*Full sample*				*Industrial countries*				*Developing countries*			
	Coeff.	*Std. dev.*	R^2	*No. obs.*	*Coeff.*	*Std. dev.*	R^2	*No. obs.*	*Coeff.*	*Std. dev.*	R^2	*No. obs.*
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.041	0.03	0.24	376	0.065**	0.03	0.46	111	0.029	0.04	0.25	265
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.102**	0.03	0.26	370	0.013	0.03	0.44	111	0.189**	0.05	0.28	259
Country-effects estimator												
De jure index $\mathbf{L}_0$	–0.088	0.08	0.55	381	–0.069	0.08	0.57	111	–0.137	0.11	0.56	270
De facto index $\mathbf{L}_1$	–0.049	0.07	0.55	398	0.030	0.06	0.52	120	–0.146	0.11	0.56	278
Minimum wage[b]	–0.108*	0.07	0.57	349	–1.150**	0.19	0.66	120	–0.185*	012	0.58	229
Social security contribution	–0.082	0.07	0.57	355	–0.086	0.41	0.52	105	–0.158*	0.10	0.58	250
Trade union membership	–0.020	0.05	0.56	366	1.379**	0.27	0.62	119	–0.096	0.09	0.57	247
General government employment	0.000	0.08	0.58	333	0.706**	0.19	0.58	120	–0.117	0.14	0.60	213
De facto index $\mathbf{L}_2$	–0.093	0.14	0.54	393	–0.070	0.20	0.52	120	–0.057	0.18	0.55	273
Minimum wage[c]	–0.376**	0.16	0.57	358	–0.306**	0.15	0.54	120	–0.477**	0.24	0.58	238
Maternity leave (no. days)	–0.147	0.13	0.57	364	0.054	0.13	0.53	117	–0.342*	0.21	0.58	247
Ratification of ILO convention 87	–0.066*	0.04	0.56	387	–0.070	0.05	0.53	120	–0.055	0.05	0.57	267
Central government employment	0.252*	0.14	0.59	335	0.559**	0.13	0.59	119	0.057	0.22	0.59	216
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.250**	0.12	0.56	376	0.701**	0.14	0.67	111	0.137	0.15	0.56	265
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.199*	0.11	0.56	370	0.163*	0.11	0.58	111	0.299*	0.16	0.57	259

* Statistically significant at the 10 percent level.

** Statistically significant at the 5 percent level.

Source: Authors' calculations, based on data from Rama and Artecona (2002) and Djankov and others (2003a).

a. We report the regression coefficient for the indicator of labor rigidity according to equation (1) in the text, based on an effective sample of seventy-six countries averaged over the 1970–2000 period. The estimation method is instrumental variables. The dependent variable is the growth rate in per capita GDP. Our control variables are output per capita (in logs), secondary schooling, domestic credit to the private sector, trade openness, governance, inflation, real exchange rate overvaluation, terms-of-trade shocks, and the labor regulation indicator. Labor regulation data are from Rama and Artecona (2002). Full regression results and standard errors of the coefficients of the labor regulation variables are not reported for reasons of space, although they are available on request. Asymptotic standard errors robust to general cross-section and time-series heteroskedasticity are reported.

b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.

c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.

Table 10. Panel Data Regression Analysis for Labor Market Regulation and Economic Growth: GMM-IV Estimations[a]

	Full sample			Industrial countries			Developing countries		
Explanatory variable	L_0 *index*	L_1 *index*	L_2 *index*	L_0 *index*	L_1 *index*	L_2 *index*	L_0 *index*	L_1 *index*	L_2 *index*
Constant	0.139**	0.185**	0.178**	2.572**	1.492*	1.678*	0.191**	0.173**	0.123**
	(0.06)	(0.02)	(0.02)	(0.60)	(0.96)	(1.07)	(0.04)	(0.04)	(0.04)
Output per capita (logs)	−0.008	−0.016**	−0.008**	−0.069**	−0.031**	−0.058**	−0.018**	−0.012**	−0.002
	(0.01)	(0.00)	(0.00)	(0.03)	(0.01)	(0.02)	(0.00)	(0.01)	(0.01)
Secondary schooling	0.021*	0.031**	0.018**	0.008	0.049*	0.103*	0.025**	0.024**	0.016**
	(0.01)	(0.00)	(0.00)	(0.04)	(0.03)	(0.06)	(0.01)	(0.01)	(0.01)
Credit to private sector	−0.008	−0.004*	−0.004*	−0.001	0.008	−0.012	−0.007**	−0.003	−0.003
	(0.01)	(0.00)	(0.00)	(0.01)	(0.02)	(0.02)	(0.00)	(0.00)	(0.00)
Inflation	−0.021**	−0.020**	−0.021**	−0.366**	−0.316**	−0.231**	−0.018**	−0.019**	−0.024**
	(0.01)	(0.00)	(0.00)	(0.10)	(0.10)	(0.11)	(0.00)	(0.00)	(0.00)
Openness	0.001	−0.008**	−0.009**	0.002	0.029	0.039	−0.004	−0.001	−0.001
	(0.01)	(0.00)	(0.00)	(0.02)	(0.03)	(0.03)	(0.01)	(0.01)	(0.01)
Terms-of-trade shocks	0.066*	0.058**	0.062**	0.185**	0.141**	0.354**	0.052**	0.037	0.044*
	(0.04)	(0.02)	(0.02)	(0.02)	(0.03)	(0.18)	(0.02)	(0.02)	(0.02)
Real exchange rate overvaluation	−0.006	−0.012**	−0.011**	−0.041*	0.034	−0.086*	−0.005	−0.011**	−0.003
	(0.01)	(0.00)	(0.00)	(0.03)	(0.04)	(0.05)	(0.00)	(0.00)	(0.00)
Governance	0.005**	0.004**	0.003**	0.003	−0.002	−0.003	0.008**	0.005**	0.003*
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)
Labor regulation indicator	−0.032	−0.006	−0.036**	−0.026	−0.154*	−0.133*	−0.043*	0.009	−0.040*
	(0.02)	(0.01)	(0.01)	(0.03)	(0.09)	(0.08)	(0.02)	(0.04)	(0.02)
Summary statistic									
No. countries	71	70	69	20	19	20	51	50	49
No. observations	220	238	235	64	72	73	156	165	162
R^2	0.18	0.19	0.27	0.47	0.36	0.51	0.14	0.23	0.29
Specification test (*p* value)									
Sargan test	0.63	0.76	0.68	0.50	0.30	0.41	0.89	0.92	0.92
Second-order correlation	0.16	0.76	0.44	0.45	0.59	0.63	0.31	0.75	0.27

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002).
a. The dependent variable is the growth rate in per capita GDP. The estimation method is the GMM-IV system estimator (Arellano and Bond, 1995; Blundell and Bond, 1998). We use the full sample of countries averaged, with nonoverlapping five-year observations over the 1970–2000 period. Labor regulation data are from Rama and Artecona (2002). Asymptotic standard errors robust to general cross-section and time-series heteroskedasticity are reported.

First, the $\mathbf{L}_0$ index of de jure regulations has a negative and significant relation with economic growth for all samples of countries, but the impact of deregulation is statistically significant only for the sample of developing countries. This result implies that a simplification of national labor codes may promote growth in developing countries. For example, if the index of de jure labor regulations for a representative developing country declined by one standard deviation (0.16), then its growth rate would increase from the regression sample mean of 1.2 percent to 1.9 percent. Also, if labor market regulations in Argentina (the developing country with the highest value for $\mathbf{L}_0$ in 1995–99) were relaxed to levels exhibited by the average developing country (say, the Philippines or Honduras), its growth rate would increase by 0.8 percentage points. Even so, while labor market deregulation might be effective at reducing de jure regulations, it might not reduce regulations in practice (Forteza and Rama, 2002).

The coefficient estimate of the $\mathbf{L}_1$ index of de facto labor regulations is negative and significant in our regression analysis only for the sample of industrial countries. Economically speaking, if the representative industrial country reduced its degree of labor regulation by one standard deviation (0.14), its growth rate would increase from the regression sample mean of 2.0 percent to 4.0 percent. However, serious efforts to deregulate labor markets in industrial countries would be required to achieve growth effects of this magnitude.[20] For developed economies, we find that all the components of the $\mathbf{L}_1$ index have a negative coefficient estimate, although one (general government employment) is not statistically significant (see table 11). If the level of market regulations in Sweden, for example, were to decline to the average level exhibited by the industrial countries, the growth rate would increase by 0.1 percentage point if the reduction is in minimum wages and by 0.8 percentage points if the decline is in social security contributions or trade union membership.

Finally, we find a negative and significant coefficient estimate for the $\mathbf{L}_2$ index of de facto labor regulations regardless of the sample of countries evaluated. From our coefficient estimates, we find that a one standard deviation decrease in the index $\mathbf{L}_2$ for industrial countries (0.1) would increase their growth rate by 1.3 percentage points, whereas an analogous decline for developing countries (0.15) would raise their growth rate by 0.6 percentage points. The negative impact

20. The level of regulations displayed by the average industrial economy over the 1990s is similar to that exhibited in the 1970s (see the average of the aggregate $\mathbf{L}_1$ index over decades in table 4).

Table 11. Labor Market Regulations and Long-term Growth: Sensitivity Analysis for GMM-IV Estimates[a]

	Full sample				*Industrial countries*				*Developing countries*			
Labor indicator	*Coeff.*	*Std. dev.*	R^2	*No. obs.*	*Coeff.*	*Std. dev.*	R^2	*No. obs.*	*Coeff.*	*Std. dev.*	R^2	*No. obs.*
De jure index $\mathbf{L}_0$	–0.032	0.02	0.18	220	–0.026	0.03	0.47	64	–0.043*	0.02	0.14	156
De facto index $\mathbf{L}_1$	–0.006	0.01	0.19	238	–0.154*	0.09	0.36	72	0.009	0.04	0.23	165
Minimum wage[b]	0.031**	0.01	0.19	210	–0.013*	0.01	0.50	73	0.018	0.04	0.22	137
Social security contribution	–0.007	0.01	0.19	212	–0.077**	0.03	0.50	65	–0.007	0.02	0.25	147
Trade union membership	–0.010	0.02	0.25	218	–0.116*	0.07	0.45	73	–0.049	0.04	0.27	145
General government employment	0.022*	0.01	0.20	193	–0.086	0.09	0.47	73	0.026	0.03	0.22	120
De facto index $\mathbf{L}_2$	–0.036**	0.01	0.27	235	–0.133*	0.08	0.51	73	–0.040*	0.02	0.29	162
Minimum wage[c]	–0.013*	0.01	0.31	217	–0.028*	0.02	0.53	73	–0.031*	0.02	0.32	144
Maternity leave (no. days)	0.008	0.02	0.24	215	0.021	0.05	0.48	70	0.023	0.05	0.26	145
Ratification of ILO convention 87	–0.028**	0.00	0.28	232	–0.080	0.08	0.53	73	–0.043*	0.02	0.13	159
Central government employment	–0.043**	0.02	0.28	195	–0.025*	0.01	0.55	73	–0.025	0.07	0.25	122
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.038**	0.01	0.16	217	0.013	0.01	0.45	64	0.073**	0.02	0.12	153
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.008	0.01	0.20	214	0.028	0.03	0.48	64	0.021*	0.01	0.16	150

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002).
a. The dependent variable is the growth rate in per capita GDP. The estimation method is the GMM-IV system estimator (Arellano and Bond, 1995; Blundell and Bond, 1998). We use the full panel data of countries, with nonoverlapping five-year observations over the 1970–2000 period. Labor regulation data are from Rama and Artecona (2002). Asymptotic standard errors robust to general cross-section and time-series heteroskedasticity are reported.
b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.
c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.

of a higher degree of labor regulation in industrial economies (as proxied by higher values of $\mathbf{L}_2$) is mainly driven by the negative and significant growth effects of higher minimum wages and higher general government employment (see table 11). The other dimensions (maternity leave and trade union membership) have no statistically significant impact. Here, an increase in the growth rate of industrial countries of 0.3 percentage points would be achieved by a one-standard-deviation reduction of either. In developing economies, the negative impact of increased labor market regulations minimum wages (0.12) or public employment (0.11). on the growth rate may be attributed to higher minimum wages and larger trade unions. A one-standard-deviation cut in minimum wages (0.17) may increase the growth rate of a developing country by 0.5 percentage point, whereas an analogous decline in the role of the trade unions may raise the growth rate of the economy by 2.0 percentage points (see table 11).

3.5 A Scorecard on the Growth Costs of Labor Regulations

To assess the growth costs of labor market regulations, we constructed a scorecard based on the seven panel estimation techniques applied to the data. Using our seven different sets of estimated coefficients, we input the value of –1 (+1) to a negative (positive) and significant coefficient estimate, and 0 to an insignificant coefficient. Table 12 reports the proportion of negative and positive significant coefficients.[21]

We obtain four main stylized facts from our scorecard. First, thicker labor codes (as proxied by the index of de jure regulations) seem to be negatively associated with economic growth for both the full sample of countries and the sample of developing countries. Second, the aggregate index of de facto labor regulations—whether the $\mathbf{L}_1$ or $\mathbf{L}_2$ index—has a weak negative relation with growth for the full sample of countries and the sample of industrial economies. The $\mathbf{L}_2$ index of de facto labor regulations has a negative relation with growth only for the sample of developing countries. Third, when the minimum wage is expressed as a ratio of the average labor cost in the manufacturing sector, it has a robust negative relation with economic growth among industrial countries.

21. A scorecard that assigns a higher value to estimation techniques that give higher points to econometric techniques that deal with unobserved country and time effects and endogeneity yields similar results.

Table 12. Scorecard of Labor Regulations and Economic Growth[a]

Labor regulation indicator	*All countries*	*Industrial countries*	*Developing countries*
De jure index $\mathbf{L}_0$	–0.6	–0.3	–0.6
De facto index $\mathbf{L}_1$	–0.3	–0.4	–0.3
Minimum wage[b]	–0.3	–1.0	–0.1
Social security contribution	–0.3	0.0	–0.4
Trade union membership	–0.3	0.3	–0.4
General government employment	–0.3	0.3	–0.4
De facto index $\mathbf{L}_2$	–0.4	–0.4	–0.6
Minimum wage[c]	–0.6	–0.9	–0.6
Maternity leave (no. days)	–0.3	0.3	–0.4
Ratification of ILO convention 87	–0.4	0.3	–0.4
Central government employment	–0.1	–0.3	–0.1
De jure versus de facto			
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.6	0.4	0.4
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.4	0.1	0.6

a. Based on seven different panel data estimations, we assigned a value of –1 (+1) to a negative (positive) and significant coefficient estimate, and 0 to an insignificant coefficient. Here, we report the proportion of negative and positive significant coefficients.
b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.
c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.

When expressed as a ratio of income per capita, it seems to have an adverse impact on growth regardless of the sample of countries used. Finally, the rest of the categories—namely, mandated benefits, trade union membership, and public employment—have a negative but weak relation with growth for both the world sample and the sample of developing countries.

Our measures of labor regulation enforcement—measured here as the difference between de facto and de jure regulations—have a positive relation with growth. This relation is more robust for the full sample of countries when we use the $\mathbf{L}_1$ index, and for developing countries when we use the $\mathbf{L}_2$ index.

4. Conclusions

This paper has assessed whether labor market regulations represent an obstacle for long-term growth. For this analysis, we used two recently developed databases on labor regulations and outcomes: Rama and Artecona (2002), which contains data on labor regulations on paper and in practice for 121 countries and is organized in five-year observations from 1945–49 to 1995–99; and Djankov and others (2003a), which analyzes the labor codes for a cross-section of eighty-five countries.

We followed the empirical literature on growth in performing our regression analysis on two levels. First, we reported the cross-sectional regression results using least squares and instrumental variables. To instrument for labor regulations, we followed Djankov and others (2003a) in the selection of our instruments (that is, the level of development, leftist political orientation of the government, trade openness, common law tradition, German civil code tradition, and institutionalized autocracy). Next, we reported the panel data regression results using three different types of estimators: least-squares-based estimators, including pooled OLS, least squares with time effects, and least squares with country dummies (fixed-effects estimator); IV estimators using pooled IV and IV with time- and country-effects; and the generalized method of moments (GMM) estimators for dynamic panel data models developed by Arellano and Bond (1991), Arellano and Bover (1995), and Blundell and Bond (1998). Here, we appropriately controlled for the presence of unobserved country effects in a dynamic panel data model, and we accounted for endogenous regressors with both external and internal instruments.

Our main findings are as follows. First, our cross-sectional analysis finds that thicker labor codes (or de jure regulations) have an adverse impact on long-run growth only among industrial economies. The impact of de facto regulations—as proxied by $\mathbf{L}_1$ and $\mathbf{L}_2$—is mixed. While we find a negative and significant relationship between growth and the $\mathbf{L}_2$ index in industrial countries, we find a negative and significant relationship between growth and the $\mathbf{L}_1$ index in developing countries. In addition, we find that all three types of labor laws described by Djankov and others (2003a) (namely, employment, industrial relations, and social security) have a significant impact on growth only among industrial countries.

Second, our GMM-IV system panel data estimates suggest that less-regulated labor codes may foster growth among developing countries. Economically speaking, if the $\mathbf{L}_0$ index declines by one standard deviation, the growth rate of a developing country should increase by 0.7 percentage points. One should be very cautious about this result, however, since simplifying labor codes does not guarantee an improvement in the ability to enforce these laws.

Third, the $\mathbf{L}_1$ index of de facto rigidities has a negative and significant relationship only for industrial economies. Our estimates suggest that a one-standard-deviation decrease in the $\mathbf{L}_1$ index may increase the growth rate of advanced economies by 2 percent. These growth effects, however, entail a significant effort to deregulate labor markets among industrial economies, especially considering that most European countries have made only marginal changes in their labor market institutions.

Fourth, a high degree of labor regulation (as proxied by high values in our $\mathbf{L}_2$ index) has an adverse and significant impact on growth in both industrial and developing countries. We find that a one-standard-deviation decline in the $\mathbf{L}_2$ index developing countries (industrial economies) would increase their growth rate by 0.6 (1.3) percentage points.

Fifth, the adverse growth effects of labor regulations among developing countries might be explained by the significant negative growth effects of minimum wages and trade unions. If minimum wages were to decline by one standard deviation, the growth rate in developing countries would increase by 0.5 percentage points; the growth rate would increase by 2.0 percentage points if an analogous decline were experienced by the role of trade unions. To achieve these growth effects, however, developing countries would have to undertake a very strong effort toward labor market deregulation.

Finally, a scorecard of our panel data estimates suggests that thicker codes are negatively related to growth for both the full sample of countries and developing countries. Also, the impact of the aggregate indices of de facto regulations is negative, but weak. The minimum wage is the only variable with a robust negative relation with growth.

Appendix A
Sample of Countries

- Industrial countries (twenty-two countries): Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Luxembourg, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, United Kingdom, and United States.
- Latin America and the Caribbean (twenty-one countries): Argentina, Bahamas, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Guatemala, Honduras, Jamaica, Mexico, Nicaragua, Panama, Paraguay, Peru, Trinidad and Tobago, Uruguay, and Venezuela.
- East Asia and the Pacific (twelve countries): China, Hong Kong, Indonesia, Korea, Malaysia, Mongolia, Papua New Guinea, Philippines, Singapore, Taiwan, Thailand, and Vietnam.
- Eastern Europe and Central Asia (seventeen countries): Belarus, Bulgaria, Croatia, Czech Republic, Estonia, Hungary, Kazakhstan, Kyrgyz Republic, Latvia, Lithuania, Poland, Romania, Russian Federation, Slovak Republic, Slovenia, Ukraine, and Yugoslavia.
- Middle East and North Africa (twenty-one countries): Algeria, Bahrain, Cyprus, Egypt, Iran, Iraq, Israel, Jordan, Kuwait, Lebanon, Libya, Malta, Morocco, Oman, Qatar, Saudi Arabia, Syria, Tunisia, Turkey, United Arab Emirates, and Yemen.
- South Asia (five countries): Bangladesh, India, Nepal, Pakistan, and Sri Lanka.
- Sub-Saharan Africa (twenty-three countries): Botswana, Burkina Faso, Côte d'Ivoire, Ethiopia, Ghana, Guinea, Guinea-Bissau, Kenya, Lesotho, Madagascar, Mali, Mauritania, Mauritius, Niger, Nigeria, Rwanda, Senegal, Sierra Leone, South Africa, Tanzania, Uganda, Zambia, and Zimbabwe.

Appendix B
Estimation Methodology: The GMM-IV System Estimator

The estimation of a growth regression using a panel data set of countries across the world poses some challenges.[22] First, there are some unobserved country- and time-specific effects. We can account for the presence of time effects by including time-specific dummy variables in our regression, but the common methods used to account for country effects (that is, within-group estimators) are inappropriate given the dynamic nature of the regression equation. Second, most explanatory variables are likely to be jointly endogenous with economic growth, and we therefore need to control for the biases resulting from simultaneous or reverse causation. The main objective of this appendix is to outline the econometric methodology we use to control for country-specific effects and joint endogeneity in a dynamic model of panel data.

We use the generalized method of moments (GMM) estimators developed by Holtz-Eakin, Newey, and Rosen (1988), Arellano and Bond (1991), and Arellano and Bover (1995) for dynamic models of panel data. These estimators take advantage of the panel data set; they are based, first, on differencing regressions or instruments to control for unobserved effects and, second, on the use of previous observations of the explanatory variables as instruments (which are called internal instruments).

After accounting for time-specific effects, we can rewrite equation (1) as follows:

$$y_{it} = \alpha\, y_{i,t-1} + \beta' \mathbf{X}_{it} + \eta_i + \varepsilon_{it} \ . \tag{B.1}$$

To eliminate the country-specific effect, we take first-differences of equation (B.1):

$$y_{it} - y_{i,t-1} = \alpha\left(y_{i,t-1} - y_{i,t-2}\right) + \beta'\left(\mathbf{X}_{it} - \mathbf{X}_{i,t-1}\right) + \left(\varepsilon_{it} - \varepsilon_{i,t-1}\right) . \tag{B.2}$$

The use of instruments is necessary to deal with the likely endogeneity of the explanatory variables and the problem that, by construction, the new error term, $\varepsilon_{it} - \varepsilon_{i,t-1}$, is correlated with the lagged dependent variable, $y_{i,t-1} - y_{i,t-2}$. The instruments consist of

22. This appendix draws heavily on Loayza, Fajnzylber, and Calderón (2003).

previous observations of the explanatory and lagged dependent variables. Given that it relies on past values as instruments, this method only allows current and future values of the explanatory variables to be affected by the error term. Therefore, while relaxing the common assumption of strict exogeneity, our instrumental variable method does not allow the **X** variables to be fully endogenous.

Under the assumptions that the error term, ε, is not serially correlated and that the explanatory variables, **X**, are weakly exogenous (that is, the explanatory variables are assumed to be uncorrelated with future realizations of the error term), the GMM dynamic panel estimator uses the following moment conditions:

$$E\left[y_{i,t-s}\left(\varepsilon_{it}-\varepsilon_{it-1}\right)\right]=0 \text{ and} \tag{B.3}$$

$$E\left[\mathbf{X}_{i,t-s}\left(\varepsilon_{it}-\varepsilon_{it-1}\right)\right]=0 , \tag{B.4}$$

for $s \geq 2$; $t = 3,\ldots, T$. The GMM estimator based on these conditions is known as the difference estimator. Notwithstanding its advantages with respect to simpler panel data estimators, the difference estimator has important statistical shortcomings. Alonso-Borrego and Arellano (1996) and Blundell and Bond (1998) show that when the explanatory variables are persistent over time, lagged levels of these variables are weak instruments for the regression equation in differences. Instrument weakness influences the asymptotic and small-sample performance of the difference estimator. Asymptotically, the variance of the coefficients rises. Monte Carlo experiments show that the weakness of the instruments can produce biased coefficients in small samples.[23]

To reduce the potential biases and imprecision associated with the usual difference estimator, we use a new estimator that combines in a system the regression in differences with the regression in levels (developed in Arellano and Bover, 1995; Blundell and Bond, 1998). The instruments for the regression in differences are the same as above. The instruments for the regression in levels are the lagged differences of the corresponding variables. These are appropriate instruments under the following additional assumption: although there may be correlation

23 An additional problem with the simple difference estimator relates to measurement error: differencing may exacerbate the bias as a result of errors in variables by decreasing the signal-to-noise ratio (see Griliches and Hausman, 1986).

between the levels of the right-hand-side variables and the country-specific effect in equation (B.1), there is no correlation between the differences of these variables and the country-specific effect. This assumption results from the following stationarity property:

$$E\left(y_{i,t+p}\,\eta_i\right)=E\left(y_{i,t+q}\,\eta_i\right) \text{ and} \tag{B.5}$$

$$E\left(\mathbf{X}_{i,t+p}\,\eta_i\right)=E\left(\mathbf{X}_{i,t+q}\,\eta_i\right),$$

for all p and q. The additional moment conditions for the second part of the system (the regression in levels) are as follows:[24]

$$E\left[\left(y_{i,t-1}-y_{i,t-2}\right)\left(\eta_i+\varepsilon_{it}\right)\right]=0 \text{ and} \tag{B.6}$$

$$E\left[\left(\mathbf{X}_{i,t-1}-\mathbf{X}_{i,t-2}\right)\left(\eta_i+\varepsilon_{it}\right)\right]=0\ . \tag{B.7}$$

Using the moment conditions presented in equations (B.3), (B.4), (B.6), and (B.7), we employ a generalized method of moments (GMM) procedure to generate consistent estimates of the parameters of interest and their asymptotic variance-covariance (Arellano and Bond, 1991; Arellano and Bover, 1995). These are given by the following formulas:

$$\hat{\theta}=\left(\bar{\mathbf{X}}'\mathbf{Z}\hat{\Omega}^{-1}\mathbf{Z}'\bar{\mathbf{X}}\right)^{-1}\bar{\mathbf{X}}'\mathbf{Z}\hat{\Omega}^{-1}\mathbf{Z}'\bar{\mathbf{y}} \text{ and} \tag{B.8}$$

$$\text{AVAR}\left(\hat{\theta}\right)=\left(\bar{\mathbf{X}}'\mathbf{Z}\hat{\Omega}^{-1}\mathbf{Z}'\bar{\mathbf{X}}\right)^{-1}, \tag{B.9}$$

where θ is the vector of parameters of interest (α, β); $\bar{\mathbf{y}}$ is the dependent variable stacked first in differences and then in levels; $\bar{\mathbf{X}}$ is the explanatory-variable matrix, including the lagged dependent variable (y_{t-1}, $\mathbf{X}$) stacked first in differences and then in levels; $\mathbf{Z}$ is the matrix

24. Given that lagged levels are used as instruments in the differences specification, only the most recent difference is used as an instrument in the levels specification. Using other lagged differences would result in redundant moment conditions (see Arellano and Bover, 1995).

of instruments derived from the moment conditions; and is a consistent estimate of the variance-covariance matrix of the moment $\hat{\Omega}$ conditions.[25]

The consistency of the GMM estimators depends on whether lagged values of the explanatory variables are valid instruments in the growth regression. We address this issue by considering two specification tests suggested by Arellano and Bond (1991) and Arellano and Bover (1995). The first is a Sargan test of overidentifying restrictions, which tests the overall validity of the instruments by analyzing the sample analog of the moment conditions used in the estimation process. Failure to reject the null hypothesis gives support to the model. The second test examines the null hypothesis that the error term, ε_{it}, is not serially correlated. As in the case of the Sargan test, the model specification is supported when the null hypothesis is not rejected. In the system specification, we test whether the differenced error term (that is, the residual of the regression in differences) is second-order serially correlated. First-order serial correlation of the differenced error term is expected even if the original error term (in levels) is uncorrelated, unless the latter follows a random walk. Second-order serial correlation of the differenced residual indicates that the original error term is serially correlated and follows a moving average process of at least order one. This would reject the appropriateness of the proposed instruments (and would call for higher-order lags to be used as instruments).

25. In practice, Arellano and Bond (1991) suggest the following two-step procedure to obtain consistent and efficient GMM estimates. First, assume that the residuals, e_{it}, are independent and homoskedastic both across countries and over time. This assumption corresponds to a specific weighting matrix that is used to produce first-step coefficient estimates. Then, construct a consistent estimate of the variance-covariance matrix of the moment conditions with the residuals obtained in the first step, and use this matrix to reestimate the parameters of interest (that is, second-step estimates). Asymptotically, the second-step estimates are superior to the first-step ones insofar as efficiency is concerned.

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Labor Market Regulations and Income Inequality: Evidence for a Panel of Countries

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The fact that labor market regulations are at the cornerstone of the debate on economic policy and political economy in many countries shows that changes in regulations may have nontrivial effects. At the very least, they have different consequences for protected and unprotected groups. They may also face interesting tradeoffs, specifically regarding efficiency and equity. In this paper, we empirically study one particular ingredient of this type of tradeoff, namely, the effect of labor regulations on income distribution.

For that purpose, we present evidence on the impact of labor regulations on income inequality using two recently published databases on labor institutions (or de jure regulations) and outcomes (or de facto regulations) (Rama and Artecona, 2002; Djankov and others, 2003). We consider other country characteristics that may affect income distribution, including income level and growth, education, and the structure of the economy. We use a battery of cross-section and panel data analysis techniques to evaluate the robustness of the results. In particular, we use cross-section ordinary least squares (OLS), pooled OLS, OLS with time and country fixed effects, cross-section instrumental variables (IV), IV with time and country fixed effects, and generalized method of moments (GMM) estimators. The sample

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we consider covers 121 countries over the 1970–2000 period. We focus on two groups: the total sample and the sample of developing countries.

This paper is closely related to Calderón and Chong (in this volume) and should, in certain dimensions, be taken as its complement. To begin with, it is based on the same data sets (except for inequality) and considers similar estimation techniques. Taken together, the papers allow the reader to determine whether the tradeoff mentioned above exists and to evaluate its relative importance.

We report four main findings. First, de jure regulations (that is, what the labor codes prescribe) do not improve income inequality. Our results using the Rama-Artecona database are not robust, but in a few cases, they indicate that regulations worsen income distribution. When we consider the Djankov–La Porta data set, we find that regulations on employment and industrial relations (although not on social security) do have a negative effect on income distribution.

Second, compliance with labor regulations, measured as the ratio between a de facto index and a de jure index, has a positive effect on income distribution. Since this result cannot be explained by summing up the individual effects of each index separately, it may capture institutional development rather than labor market considerations.

Third, de facto regulations are weakly associated with a better income distribution. This result could be due to endogeneity of labor regulations. When we control for this problem, the effect frequently is not different from zero, although in some cases these regulations improve income distribution.

Finally, aside from the endogeneity problem, these mixed results are partly explained by the fact that the results can differ markedly across specific de facto regulations. In this regard, the most robust results are the following: minimum wages, especially when measured as a percentage of per capita income, worsen income inequality; trade union membership (as a percentage of the labor force) has a positive effect on income distribution, although its effect on the poorest 20 percent is smaller and less robust than for the middle class; government employment at the general level (less so at the central level) has a positive effect on income distribution, but no effect on the poorest quintile; days of maternity leave have a positive effect on income distribution; and neither the ratification of the International Labor Organization's convention 87 nor social security contributions has a robust effect on income inequality across estimation methods and samples.

The paper is organized as follows. Section 1 presents a brief literature review on the impact of labor market regulations on income inequality. Section 2 reviews the data sets and the methodology we use. Section 3 presents the results of the different estimation techniques. Section 4 discusses our overall results and concludes.

1. Literature Review

In a seminal paper, Kuznets (1955) argues that the relation between income inequality and the level of development follows an inverted-U-shaped curve. Inequality rises in the face of economic expansion during the initial stages of development, and it declines afterwards. The relation stipulated by Kuznets has recently been simulated successfully within a general equilibrium framework (Galor and Tsiddon, 1996). Recent evidence also shows that unemployment is one of the major sources of inequality (Jenkins, 1995, 1996) and labor market policies are a potential instrument for reducing inequality (Rama, 2001a).

Saint-Paul (1999) claims that labor markets institutions across the world usually consist of tax systems or other transfer mechanisms that shift resources from the working to the nonworking population. These institutions include unemployment benefits, employment protection laws, and active employment policies by the government. Some analysts argue that these institutions are necessary to protect workers from bad outcomes and unexpected shocks (Blanchard, in this volume). In general, labor market institutions are supposed to help achieve socially desirable redistributive goals (Emerson and Dramais, 1988; Rama, 2001a, 2003). In this context, labor market policies may be an effective tool for reducing income inequality, but there is increasing debate on the benefits of labor policies such as minimum wages, mandated benefits, collective bargaining, job security, and public sector employment in developing countries (Rama, 2001a, 2003).

Regarding the imposition of minimum wages, Saint-Paul (1994) argues that it may have an adverse effect on the income distribution. Minimum wages redistribute income from skilled to unskilled labor, as well as from the poorest to the lower-middle quintiles by generating unemployment.[1] Microeconomic studies suggest that the impact

1. Saint-Paul (1994) claims that minimum wages create unemployment among unskilled workers and reduce the income of skilled workers, thus lowering output. In addition, the impact of minimum wages on inequality is affected by other forms of labor rigidities. For example, income is shared equally among unskilled workers in a world with high job turnover, so minimum wages have a small impact on inequality among the unskilled.

of minimum wages on income inequality is small in many developing countries (Maloney and Nuñez, 2001). On the other hand, Rama (2001b) analyzes the doubling of minimum wages (in real terms) in Indonesia in the early 1990s. He finds that the elasticity of average to minimum wages was approximately 10 percent over this period, and the doubling of minimum wages was associated with a slight decline in total wage employment and a substantial increase in unemployment among small enterprises. Trade union membership, in turn, seems to guarantee a higher wage for union members relative to nonmembers. However, union wage premiums in developing countries is smaller than among industrial countries. This finding may be due to the role of trade unions in keeping wage rates invariant in periods of economic adjustment (Nelson, 1991).

Rama (2001a) reviews studies on the impact of public sector employment on income inequality. Public sector wages could have a significant effect on private sector wages in countries with a small formal sector, as in sub-Saharan Africa, (Rama, 2000).

The impact of separation costs on employment and income distribution depends on the tightness of job security regulations. Fallon and Lucas (1991) find that very strict regulations on job security depressed labor demand in India and Zimbabwe. Separation costs—in the form of mandatory severance payments—may also reduce the level of employment (Heckman and Pagés, 2000).

Rama (2003) analyzes the impact of labor market interventions on indicators of income inequality after controlling for some of their determinants.[2] He shows that social security programs help reduce income inequality. Collective bargaining is less effective at improving the income distribution: its impact is statistically significant only for the share of the second richest quintile of the population. The core conventions supported by the International Labor Organization (ILO) seem ineffective for reducing inequality.[3] In summary, countries pushing to adopt ILO labor standards, raise minimum wages, or expand government employment may not generate any significant effect on inequality.

Finally, Vanhoudt (1997) analyzes the impact of labor market policies on income inequality in member countries of the Organization

2. Rama (2003) includes the following as determinants of income inequality: educational attainment, civil liberties, and financial development.

3. According to Rama, the core ILO conventions are those addressing the abolition of forced labor, the effective elimination of child labor, nondiscrimination in the workplace, and freedom of association and the right to collective bargaining.

for Economic Cooperation and Development (OECD). He finds that the Gini coefficient is not affected by labor market policies, although other measures of inequality are. Specifically, he finds that active labor market policies—such as expenditures for public employment services, labor market training, and subsidized employment—improve the income share of the bottom quintiles of the population and reduce the income gap between the top and bottom quintiles. Passive labor markets—that is, income compensation schemes—have a negligible impact.

2. The Data and Methodology

This section describes the database used in our regression analysis, as well as our estimation strategy. Since our discussion draws heavily on Calderón and Chong (in this volume), we present only a brief description of both the data and the methodology used.

2.1 The Data

We use two recently developed databases on labor regulations to test whether labor regulations have been an effective tool for reducing income inequality: namely, the Rama-Artecona database (Rama and Artecona, 2002) and the Djankov–La Porta database (Djankov and others, 2003).[4]

The Rama-Artecona Database

Rama and Artecona (2002) collect data on labor market regulations and outcomes in 121 countries over the period 1945–99. The data is organized in five-year averages and distinguishes between de jure regulation and de facto regulation. De jure regulation is approximated by the number of ILO standards ratified by the national labor laws.[5] De facto regulation is approximated by information on categories such as

4. This subsection draws heavily on Calderón and Chong (in this volume).

5. The ratified conventions included in this index encompass universal legislation on issues such as child labor, compulsory labor, equal remuneration for male and female workers, equal opportunity, the right of collective bargaining, and organization in unions.

minimum wages, conditions of work and benefits, trade union membership and collective bargaining, and public sector employment. The distinction between de jure and de facto regulation is very important because the ability of developing countries to enforce the regulations stipulated in their labor laws is quite limited (Squire and Suthiwart-Narueput, 1997).

We define aggregate indices of the overall extent of labor regulations in the economy following the strategy pursued by Rama (1995) and Forteza and Rama (2002). We define an index of de jure regulation, $\mathbf{L}_0$, as the cumulative number of ILO conventions ratified by a country over time. This index reflects the ideal regulatory framework of the country from an institutionalist point of view (Freeman, 1993a, 1993b), while it also captures the thickness of the labor code (Forteza and Rama, 2002). The $\mathbf{L}_0$ index includes the ratification of ILO conventions on the minimum age of employment, compulsory labor, the abolition of forced labor, equal remuneration for men and women, the right to collective bargaining, and the discrimination on equality of opportunity or conditions of employment on the basis of race, religion, sex, political opinion, or social origin. However, the number of existing regulations does not give us information on a country's ability to implement and enforce these regulations. For this reason, we require an index that reflects the practical extent of labor regulations instead of their number.

Rama (1995) constructs an aggregate index of de facto regulations using information on the following four categories: minimum wages, mandated benefits, trade union membership, and public sector employment. Unfortunately, data on job separation costs are only available for a quite limited sample of countries.[6] Following Rama (1995) and Forteza and Rama (2002), we construct two aggregate indices of de facto labor regulations, both of which include different proxies for these four dimensions. The first aggregate index of labor de facto regulations, $\mathbf{L}_1$, includes the simple average of the ratio of the minimum wage to unit labor costs in the manufacturing sector; social security contributions as a percentage of salaries; total trade union membership as a percentage of total labor force; and the share of general government employment in total employment. The second aggregate index of de facto regulations, $\mathbf{L}_2$, comprises the simple average of the ratio of

6. Heckman and Pagés (2000) construct data on job separation costs for Latin America; they find that these costs have a substantial impact on the level of employment in the region.

minimum wage to income per capita; the number of days of maternity leave for a first child born without complications; the ratification of ILO convention 87, which allows workers to establish organizations; and the ratio of central government employment to total employment.

To make all these variables comparable across countries, we normalize all the labor market regulation indicators in such a way that their values fluctuate between zero and one, with one representing the highest practical extent of labor regulation and zero the lowest. The aggregate indices of de facto regulation, $\mathbf{L}_1$ and $\mathbf{L}_2$, are computed for countries with at least two of the four dimensions involved in the analysis.

The Djankov-La Porta Database

Djankov and others (2003) evaluate the degree of labor market regulation in the labor codes of eighty-five countries. Their sample thus represents a cross-section of labor regulation indices for a broad sample of countries. Since these measures are extracted from labor codes, they are closer in spirit to de jure labor rigidities than de facto enforcement.

These measures focus on three types of labor laws: employment laws, industrial relations laws, and social security laws. Employment laws contemplate the laws governing individual employment contracts in the economy. This type of law specifically regulates aspects of individual labor contracts, terms of reference, and contract termination. It covers the restrictions placed on alternative employment contracts, the conditions of the employment contract, and job security.

Industrial relations laws regulate the adoption, bargaining, and enforcement of collective agreements, the unionization of workers, and industrial actions by workers and employers. These laws capture aspects of the worker-employer relationship, such as collective bargaining, the participation of workers in the company's management, and the resolution of collective disputes (such as strikes and lockouts).

Finally, social security laws contemplate the social response to quality-of-life conditions and requirements. They protect workers against the risk of disability, sickness, and unemployment.

Income Inequality and its Determinants

The dependent variable in our regression analysis is the Gini coefficient. Our main source of data is information gathered by Deininger and Squire (1996), but this source only covers through 1995. For the final five years, we extrapolate data for income shares and the Gini coefficient for the countries present in the analysis of Milanovic (2002a, 2002b). For countries that are absent in Milanovic's papers, we generate information on the Gini coefficient based on the coefficient of variation of income and income's linear correlation with ranks, as in Milanovic (1997). We also use the income shares of the top, bottom, and middle quintiles of the population. This allows us to analyze the robustness of our results to changes in the dependent variable and to assess the impact of labor market policies on the income of the poor.

Our choice of the set of determinants of income inequality follows the empirical literature on income distribution (Milanovic, 2000; Gradstein, Milanovic, and Ying, 2001; Chong, 2002; Clarke, Xu, and Zou, 2003). We include the log of the level of GDP per capita and its squared value. This variable is obtained from the Penn World Table 6.1 (Heston, Summers, and Aten, 2002). The squared specification of the GDP per capita allows us to test for the presence of the Kuznets curve (that is, whether income inequality rises in the early stages of development and declines in later stages). We also consider indicators of education, such as the level of secondary schooling (Barro and Lee, 2000), and of financial depth, such as the ratio of credit to the private sector to GDP (Beck, Demirgüç-Kunt, and Levine, 2000). The number of physicians per 1,000 people is included as a proxy for improvements in the health sector. Macroeconomic instability is proxied by the consumer price index (CPI) inflation rate, and the size of the modern sector is calculated as the share of industry and services in the economy's total value added.

2.2 The Methodology

Our main goal is to assess the impact of labor regulations on income distribution by running the following regression:

$$y_{it} = \mu_i + \eta_t + X_{it}\beta + L_{it}\Gamma + \xi_{it} \quad . \tag{1}$$

According to this equation, income inequality in country i during period t (y_{it}) depends on a set of determinants described by the matrix $\mathbf{X}_{it}$, as well as on unobserved country- and period-specific effects (μ_i and η_t, respectively). Our set of long-term growth determinants follows the work of Milanovic (2000), Gradstein, Milanovic, and Ying (2001), and Chong (2002). The determinants of income inequality include (in logs) the initial level of per capita output and per capita output squared, human capital, financial depth, health, inflation, and the size of the modern sector (manufacturing and services).

Our income inequality regression framework also includes a set of variables that captures the extent of regulations in the labor market, as represented by the matrix $\mathbf{L}_{it}$ in equation (1). This matrix includes different indicators that focus on specific policy or institutions in the labor market, such as minimum wages, mandatory benefits, trade union membership, government employment, social security laws, and collective bargaining. The matrix $\mathbf{L}_{it}$ consists of a series of K labor regulations,

$$\left\{\ell_{it}^{\,k}\right\}_{k=1}^{K}.$$

The larger the values of these variables, the more regulated are the labor markets. We do not assume that labor regulations and outcomes are time invariant, but rather expect them to change over longer horizons.

We normalize these variables in such a way that they are equal to one (zero) if labor markets are fully regulated (deregulated).[7] If our dependent variable is the Gini coefficient, a negative estimate for the parameters in the Γ matrix implies that deregulating labor markets may enhance the distribution of income.

We encounter additional problems when we attempt to run a regression of equation 1, in that some variables in the $\mathbf{L}_{it}$ matrix may be highly correlated with each other. Trade unions and public employment display the highest correlation at 0.8, while mandated benefits and minimum wages have a correlation of 0.5. In this case, we may be unable to identify the parameters of the Γ matrix. To address this issue, we create aggregate indices of labor market regulations, as in Rama (1995) and Forteza and Rama (2002). We compute a simple average of the normalized values of our labor regulation indicators as

7. The variables are not all expressed in comparable units, so we need to normalize them before we can aggregate them. We defined our labor market rigidity indicator above as $\ell_{it}^{\,k}$, for $k = 1,\ldots, K$. Next, we define $\ell_{t\,\max}^{\,k}$ and $\ell_{t\,\min}^{\,k}$ as the closest and farthest a country can get to perfect competition in the labor markets. We then define our normalized labor market rigidity indicator as

$$\tilde{\ell}_{it}^{\,k} = \frac{\ell_{it}^{\,k} - \ell^{k}_{\min}}{\ell^{k}_{\max} - \ell^{k}_{\min}}.$$

described above.[8] We then use the aggregate index of regulations in the labor market, ℓ_{it}^{A}, to test the overall effects of labor market regulation on income inequality. We reformulate our income inequality regression in equation (1) as follows:

$$y_{it} = \mu_i + \eta_t + X_{it}\beta + \gamma_A \ell_{it}^{A} + \xi_{it} \ . \tag{2}$$

The nature and magnitude of the overall impact of labor market regulations on income inequality is captured by the sign and size of γ_A. However, individual regulations may have different consequences that cancel each other to some extent in the aggregate. One of the shortcomings of this approach is that a significant parameter estimate for γ_A may not help identify the specific regulations that need to be reformulated. Consequently, we still need to estimate the individual effect of different regulations, as captured by the γ_j parameters.

If we replace the aggregate index, ℓ_{it}^{A}, in equation (2) by one of our individual measures of labor market regulations, the coefficient estimate will be biased due to omitted variables. That is, the coefficient of the individual regulation will capture the effects of the labor market rigidity, *k*, as well as some of the effects of all of the other missing rigidities. Since the different rigidities are likely to be correlated with each other, the value obtained for γ_k might reflect the effects of these other rigidities. We can partially solve this problem by defining complementary labor market regulations, $\tilde{\ell}_{it}^{-k}$, as the average of the indicators that are different from *k*. This complementary variable can be used to control for all other labor market features, apart from itself, based on the following model:

$$y_{it} = \mu_i + \eta_t + X_{it}\beta + \gamma_k \tilde{\ell}_{it}^{k} + \gamma_{-k} \tilde{\ell}_{it}^{-k} + \xi_{it} \ , \tag{3}$$

where the coefficient γ_k captures the effect of labor market rigidity, *k*, on long-term growth.

The Estimation Strategy

We estimate our regression equation on two dimensions: cross-section and panel data.[9] Our cross-section regressions are estimated

8. In principle, we compute the average of *J* out of the *K* relevant labor market rigidities (where $J \leq K$). Our aggregate index takes values between zero and one, but unless all of the labor market rigidities are perfectly correlated with each other, the actual range of variation across countries should be significantly narrower for the aggregate measures than for any of the individual indicators.

9. Here again, we draw heavily on Calderón and Chong (in this volume).

using least squares with robust standard errors (White, 1980). We then use an IV estimator in which we control for the endogeneity of labor market regulations using a set of instruments outlined by Djankov and others (2003). We discuss the outline of the IV strategy when we analyze the panel data techniques.

For the panel estimation of equations (2) and (3), we first use a series of three least-squares-based estimators: the pooled OLS estimator, which is the simplest regression technique given that we do not account for either unobserved effects or endogeneity; the time-effects estimator (that is, least squares with time dummies), through which we can explain differences in income inequality across country stemming from differences in the extent of labor market regulations; and the within-group or country-effects estimator (that is, least squares with country dummies), with which we analyze the movement of income inequality indicators in a country to changes in its labor market regulations. We complement these least-squares-based estimation techniques with methods that control for endogenous regressors. We thus present several estimators from the instrumental variables family.

Because it is very likely that labor regulations are partly endogenous, we focus our final analysis on techniques that account for the endogeneity problems. We tackle this issue using two different strategies. Our first strategy is based on IV techniques in which we select external instruments for labor regulations. We present pooled IV estimates, IV with time effects, and IV with country effects. This set of instruments follows the literature on the choice of labor regulations, as outlined by Djankov and others (2003). According to Djankov and associates, the choice of labor regulations across countries is explained by efficiency considerations, political power theories, and legal theories.

North (1981) claims that a set of regulations is usually chosen based on an efficiency criterion. Efficiency theory focuses on the distinction between regulation and social insurance. Some economists argue that social insurance may be an efficient way to deal with market failures in countries with lower social marginal cost of tax revenues—in other words, in richer countries (Becker and Mulligan, 2000). Poor countries regulate to protect workers from being mistreated by employers, while rich countries provide unemployment insurance, sick leave, and early retirement since they can raise taxes cheaply to finance such operations. Efficiency theory may argue the opposite, however. Government officials may use labor regulations to force firms to hire and keep excess labor or to empower unions that are friendly with the government. In this case, countries with good governance have a comparative advantage at regulation relative to other forms of social control of business.

According to political power theories, institutions are designed to transfer resources from those without political power to those with power (Olson, 1993). Institutions are thus designed to be inefficient by political leaders aiming to help themselves and their favored groups. Political power theorists argue that regulations protecting workers are introduced by socialist, social-democratic, and more generally leftist governments to benefit their political constituencies (Hicks, 1999). In addition, labor regulations are a response to pressure from trade unions, and the degree of regulations should be higher when unions are more powerful. Dictatorships are less constrained than democratically elected governments, so they will have more redistributive laws and institutions. Constitutions, legislative constraints, and other forms of checks and balances are all conducive to fewer regulations (Djankov and others, 2003). Likewise, open economies may find it expensive to introduce regulations, since competition makes it less lucrative for governments to raise firms' regulatory costs (Ades and Di Tella, 1999).

Finally, legal theories suggest that the legal tradition is at the root of the way countries control economic activities (Djankov and others, 2003). Common law countries tend to rely on markets and contracts, civil law countries on regulation, and socialist countries on state ownership.[10] This implies that civil law countries and socialist law countries should regulate labor markets more extensively than common law countries. Common law countries may also have a less generous social security system since they rely on markets to provide insurance.

10. Common law emerged in England and is mostly characterized by the importance of decisionmaking by juries, independent judges, and judicial discretion as opposed to codes. Common law was transmitted to the British colonies, including Australia, Canada, India, New Zealand, Pakistan, the United States, and a number of countries in the Caribbean, East Africa, and Southeast Asia. Civil law evolved from Roman law in Western Europe and was incorporated into civil codes in France and Germany in the nineteenth century. It is characterized by less independent judiciaries, the relative unimportance of juries, and a greater role of both substantive and procedural codes as opposed to judicial discretion. French civil law was transplanted throughout Western Europe, including Belgium, Holland, Italy, Portugal, and Spain, and subsequently to the colonies in North and West Africa, Latin America, and parts in Asia. German codes became accepted in Germanic Western Europe, but were also transplanted to Japan and from there to China, Korea, and Taiwan. Socialist law was adopted in countries that came under the influence of the Soviet Union, while a Scandinavian legal tradition developed in Denmark, Finland, Iceland, Norway, and Sweden (Djankov and others, 2003).

Our set of instruments reflects these different theories on the processes affecting labor regulation. To capture efficiency effects, we use the log of GDP per capita. To test political power theories, we analyze the significance of the index of institutionalized autocracy from the Polity IV codebook (Marshall and Jaggers, 2003), the leftist political orientation of the government and congress (Beck and others, 2001), and measures of trade openness. Finally, we include dummy variables for countries with British common law and the German civil code to test legal theories (La Porta and others, 1999).

Our second strategy for tackling the endogeneity of labor rigidities is to use the GMM estimators developed by Arellano and Bover (1995) and Blundell and Bond (1998). This technique takes into account the presence of unobserved period- and country-specific effects. Time effects are accounted for by the inclusion of period-specific dummy variables, whereas country-specific effects are dealt with via differencing, given the dynamic nature of the regression. We also control for biases resulting from simultaneous or reverse causation. A more detailed reference to the GMM-IV techniques is presented in appendix B in Calderón and Chong (in this volume).

3. Empirical Assessment

This section present our empirical assessment of the link between income inequality and labor market regulation. We gather data for a sample of 121 countries over the 1970–2000 period (see appendix A for a list of the countries). We present some basic statistics on income inequality and labor regulations, as well as the correlation analysis. We then perform the regression analysis. Our assessment is undertaken along two dimensions: a cross-section analysis over the 1970–2000 period and a panel data analysis of nonoverlapping five-year-average observations over the same period.

3.1 Basic Statistics

Table 1 reports simple averages of income inequality and the indicators of labor regulation across the world for a cross-section of countries over the 1970–2000 period. First, we find that the distribution of income is more egalitarian among industrial nations (with an average

Gini coefficient of 0.32) than among developing countries (0.41). Income distribution in Latin America is more unequal, on average, than among the whole set of developing countries in our sample. Second, labor codes in industrial countries (as proxied by the $\mathbf{L}_0$ index of labor market rigidity in the Rama-Artecona data set) contain more regulations (that is, ILO standards) than developing countries. Third, industrial countries have a higher ability to enforce regulations than developing countries (as displayed by the $\mathbf{L}_1$ and $\mathbf{L}_2$ indices in the Rama-Artecona data set). Latin American countries have an even lower enforcement capability. Among the component variables in the aggregate $\mathbf{L}_1$ and $\mathbf{L}_2$ indices (not shown in the table), the ratio of minimum wages to income per capita is larger in developing economies

Table 1. Basic Statistics for Labor Market Regulation and Income Inequality, 1970–2000[a]
Average across groups of countries

Variable	*All countries*	*Industrial economies*	*Developing countries*	*East Asia*	*Latin America*	*Chile*
Income distribution[b]						
Gini coefficient (0–1)	0.39	0.32	0.41	0.39	0.48	0.53
Income shares by quintile (percent)						
Top quintile	46.4	39.3	48.9	46.8	55.0	61.6
Second quintile	67.5	62.6	69.3	68.3	74.7	77.4
Third quintile	15.5	17.8	14.8	15.0	13.0	12.0
Fourth quintile	16.9	19.6	16.0	16.7	12.2	10.6
Bottom quintile	6.3	7.0	6.0	6.1	4.2	3.9
Labor market rigidity[c]						
De jure index $\mathbf{L}_0$	0.30	0.49	0.25	0.09	0.34	0.33
De facto index $\mathbf{L}_1$	0.28	0.36	0.25	0.18	0.25	0.17
De facto index $\mathbf{L}_2$	0.29	0.32	0.28	0.14	0.32	0.08
De jure versus de facto						
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	−0.04	−0.12	−0.01	0.08	−0.09	−0.16
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	−0.02	−0.17	0.03	0.06	−0.02	−0.26
Labor regulation[d]						
Employment laws	1.53	1.36	1.60	1.39	1.79	1.46
Industrial (collective) relations law	1.25	1.22	1.26	1.12	1.44	1.18
Social security laws	1.70	2.21	1.53	1.58	1.69	1.98

Source: Authors' calculations, based on data from Deininger and Squire (1996); Milanovic (2000); Rama and Artecona (2002); Djankov and others (2003).
a. Based on a cross-section sample of 121 countries for the period 1970–2000. All variables are normalized. For the mean of the different subcategories of the aggregate indices of labor institutions, see Calderon and Chong (in this volume).
b. Indicators of income distribution are from Deininger and Squire (1996) and Milanovic (2000).
c. Indicators of labor market rigidity are from Rama and Artecona (2002).
d. Indicators of labor regulations are from Djankov and others (2003).

than in industrial countries, while social security contributions as a percentage of workers' salaries, trade union membership, and public sector employment (proxied by employment in the central or general government) are larger in industrial countries than in developing nations.

Finally, using the Djankov-La Porta data set of labor regulations, we find that labor codes in developing countries contain more regulations regarding employment laws and industrial (collective) relations laws than do labor codes in industrial countries. Latin American countries, in particular, appear to have a high degree of regulations. On the other hand, labor codes in industrial countries contain more benefits in their social security laws. Further analysis of the components of the different aggregate indices of laws protecting workers (not shown in the table) indicates that regulations on the conditions of employment are significantly larger among developing nations than among industrial countries; industrial countries have more regulations regarding the participation of workers in management than developing countries, although the latter group has more regulations on collective bargaining and collective disputes; and workers in industrial countries are more protected than those in developing countries in terms of the benefits stipulated in their social security laws, especially in the area of unemployment benefits.

In table 2, we present the evolution of the sample averages by decade over the 1970–2000 period. Our panel statistics are reported for all of countries, and for the sample of developing, Latin American countries and Chile. We find that income inequality decreased over the period regardless of the sample of countries evaluated. Gini coefficients decreased (from 0.40 in the 1970s to 0.38 in the 1990s), the income shares of the top quintiles decreased, and the income shares of middle and bottom quintiles increased. Second, labor codes incorporated more ILO standards over time. Specifically, the $\mathbf{L}_0$ index increased from 0.27 in the 1970s to 0.32 in the 1990s for the full sample of countries. Third, the enforcement of labor regulations also increased, on average, over time for the full sample of countries (whether we use the aggregate $\mathbf{L}_1$ or $\mathbf{L}_2$ index). Finally, a closer look into the components of the aggregate $\mathbf{L}_1$ and $\mathbf{L}_2$ indexes (not shown in the table) yields the following result: the increase in the aggregate $\mathbf{L}_1$ and $\mathbf{L}_2$ indices among developing nations is explained by upward trends in minimum wages and social security contributions.

Table 2. Basic Statistics for Labor Market Regulation and Income Inequality over the Decades[a]

	All countries			*Developing countries*			*Latin America*			*Chile*		
Variable	*1970s*	*1980s*	*1990s*	*1970s*	*1980s*	*1990s*	*1970s*	*1980s*	*1990s*	*1970s*	*1980s*	*1990s*
Income distribution												
Gini coefficient (0–1)	0.40	0.39	0.38	0.43	0.41	0.40	0.49	0.48	0.47	0.49	0.56	0.56
Income shares by quintile (percent)												
Top quintile	47.4	46.3	45.7	50.4	48.8	47.9	53.9	55.3	56.0	59.7	64.5	60.6
Top two quintiles	68.4	67.5	66.9	70.4	69.2	68.4	75.0	74.4	74.8	75.0	78.3	78.9
Middle quintile	15.2	15.6	15.8	14.2	14.8	15.2	12.3	13.4	13.3	13.8	11.6	10.9
Bottom two quintiles	16.4	16.9	17.3	15.4	16.1	16.4	12.7	12.1	11.9	11.2	10.1	10.2
Bottom quintile	6.1	6.3	6.5	5.8	6.1	6.2	4.4	4.2	4.1	4.2	3.9	3.7
Labor market rigidity												
De jure index $\mathbf{L}_0$	0.27	0.29	0.32	0.23	0.25	0.27	0.30	0.34	0.39	0.32	0.32	0.36
De facto index $\mathbf{L}_1$	0.27	0.27	0.28	0.24	0.25	0.26	0.24	0.26	0.24	0.15	0.17	0.20
De facto index $\mathbf{L}_2$	0.28	0.29	0.30	0.27	0.27	0.29	0.31	0.33	0.32	0.06	0.06	0.11
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.00	−0.02	−0.06	0.01	0.00	−0.03	−0.06	−0.08	−0.14	−0.16	−0.15	−0.16
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.01	−0.01	−0.04	0.05	0.03	0.01	0.01	−0.01	−0.07	−0.26	−0.26	−0.26

Source: Authors' calculations, based on data from Deininger and Squire (1996); Milanovic (2000); and Rama and Artecona (2002).

a. Based on panel data of a sample of 121 countries for the period 1970–2000, in nonoverlapping five-year-average observations. All variables are normalized. For the mean of the different subcategories of the aggregate indices of labor institutions, see Calderon and Chong (in this volume). Indicators of income distribution are from Deininger and Squire (1996) and Milanovic (2000); indicators of labor market rigidity are from Rama and Artecona (2002).

3.2 Correlation Analysis

Table 3 presents the cross-section correlation analysis of income inequality and labor regulation indicators for the full sample of countries and for developing countries.[11] For the sake of robustness, we use not only different sets of labor market rigidity indicators, but also different measures of income inequality (namely, Gini coefficients and income shares). We first present the cross-section correlation between inequality and the labor market rigidity indicators in the Rama-Artecona data set. In general, we find that de jure labor regulation (as proxied by the $\mathbf{L}_0$ index) and de facto labor regulation (as proxied by the aggregate $\mathbf{L}_1$ and $\mathbf{L}_2$ indices) have a negative association with the Gini coefficient for the full sample of countries. All three labor regulation indices have a negative correlation with the income shares of the top quintiles of the population and a positive association with the income shares of the middle and bottom quintiles (see table 3). In particular, the aggregate $\mathbf{L}_1$ index of de facto rigidities has a larger negative correlation with the Gini coefficient than the $\mathbf{L}_2$ index (–0.46 versus –0.12).

A further look in the correlation between income inequality (as proxied by the Gini coefficient) and the aggregate indices of labor regulation yields two important results. First, minimum wages and trade union membership in the $\mathbf{L}_1$ index display the largest correlation with the Gini coefficient (approximately –0.5). Second, trade union membership and public sector employment in the $\mathbf{L}_2$ index exhibit the largest negative association with the Gini coefficient (with a correlation coefficient of approximately –0.1). This preliminary evidence suggests that the countries with greater labor regulations (independently of whether they are de jure or de facto) tend to display lower levels of income inequality.

Table 3 also presents the cross-section correlation between income inequality and the labor regulation indicators in the Djankov–La Porta data set. We find that the aggregate index of employment laws (as well as the different subindices) are positively correlated with the Gini coefficient, with the largest positive correlation displayed by regulations on job security. We also find a negative association between the index of industrial relations laws and the Gini coefficient that is mainly driven by worker participation in management.

11. For reasons of space, we comment only on the results for the full sample of countries. Where necessary, we point out some differences in the correlation analysis between industrial and developing countries.

Table 3. Cross-section Correlation Analysis for Labor Market Regulation and Income Inequality, 1970–2000[a]

	Full sample of countries						Developing countries					
		Income quintile						Income quintile				
Variable	*Gini coeff.*	*Top*	*Top two*	*Middle*	*Bottom two*	*Bottom*	*Gini coeff.*	*Top*	*Top two*	*Middle*	*Bottom two*	*Bottom*
Labor market rigidity[b]												
De jure index $\mathbf{L}_0$	−0.28	−0.23	−0.25	0.29	0.20	0.17	−0.08	0.02	−0.05	0.15	−0.02	0.02
De facto index $\mathbf{L}_1$	−0.46	−0.44	−0.44	0.36	0.43	0.36	−0.44	−0.39	−0.43	0.34	0.42	0.37
Minimum wage[c]	−0.49	−0.47	−0.43	0.34	0.44	0.36	−0.48	−0.42	−0.42	0.29	0.44	0.38
Social security contribution	−0.08	−0.15	−0.13	0.15	0.10	0.08	−0.17	−0.28	−0.27	0.25	0.26	0.28
Trade union membership	−0.48	−0.46	−0.43	0.32	0.44	0.37	−0.42	−0.36	−0.36	0.24	0.38	0.33
General government employment	−0.41	−0.40	−0.38	0.34	0.36	0.27	−0.40	−0.33	−0.35	0.28	0.35	0.30
De facto index $\mathbf{L}_2$	−0.12	−0.14	−0.13	0.16	0.09	0.04	−0.08	−0.08	−0.08	0.13	0.04	0.00
Minimum wage[d]	−0.06	−0.08	−0.05	0.09	0.02	−0.04	0.00	0.00	0.03	0.03	−0.06	−0.11
Maternity leave (no. days)	0.24	0.06	0.06	−0.06	−0.06	−0.02	0.18	−0.05	−0.05	0.04	0.06	0.11
Ratification of ILO convention 87	−0.10	−0.12	−0.10	0.16	0.05	0.00	0.01	0.00	0.00	0.09	−0.06	−0.10
Central government employment	−0.09	0.02	0.03	−0.06	−0.01	−0.07	−0.03	0.18	0.18	−0.17	−0.16	−0.18
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	−0.09	−0.14	−0.11	0.01	0.16	0.12	−0.36	−0.43	−0.39	0.20	0.46	0.36
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.15	0.09	0.12	−0.15	−0.08	−0.11	−0.06	−0.16	−0.10	0.03	0.14	0.05
Labor regulation[e]												
Employment laws	0.10	0.09	0.06	−0.08	−0.05	−0.03	0.07	−0.03	0.01	0.00	−0.01	−0.08
Alternative employment contracts	0.07	0.02	0.03	−0.02	−0.03	−0.05	0.17	0.09	0.12	−0.09	−0.12	−0.16
Conditions of employment	0.05	0.06	0.01	−0.02	0.00	0.05	−0.10	−0.17	−0.16	0.16	0.14	0.13
Job security	0.10	0.10	0.10	−0.13	−0.08	−0.09	0.10	0.03	0.07	−0.07	−0.07	−0.14

Table 3. (continued)

	Full sample of countries						Developing countries					
		Income quintile						Income quintile				
Variable	*Gini coeff.*	*Top*	*Top two*	*Middle*	*Bottom two*	*Bottom*	*Gini coeff.*	*Top*	*Top two*	*Middle*	*Bottom two*	*Bottom*
Industrial (collective) relations law	–0.01	0.03	0.01	0.03	–0.04	–0.01	0.03	0.03	0.02	0.08	–0.07	–0.05
Collective bargaining	0.11	0.13	0.11	–0.07	–0.12	–0.10	0.13	0.12	0.11	–0.07	–0.13	–0.14
Worker participation in management	–0.23	–0.23	–0.17	0.12	0.18	0.16	–0.12	–0.17	–0.11	0.11	0.09	0.12
Collective disputes	0.14	0.23	0.11	0.02	–0.19	–0.13	0.05	0.14	0.02	0.14	–0.12	–0.09
Social security laws	–0.38	–0.36	–0.35	0.39	0.29	0.19	–0.27	–0.21	–0.24	0.27	0.20	0.14
Old age, disability, and death benefits	–0.23	–0.31	–0.25	0.29	0.20	0.07	–0.10	–0.15	–0.12	0.12	0.11	0.02
Sickness and health benefits	–0.17	–0.11	–0.15	0.22	0.10	0.05	–0.10	–0.03	–0.10	0.17	0.05	0.02
Unemployment benefits	–0.47	–0.45	–0.42	0.41	0.37	0.28	–0.36	–0.31	–0.31	0.29	0.28	0.22

Source: Authors' calculations, based on data from Rama and Artecona (2002) and Djankov and others (2003).
a. Based on a cross-section sample of 121 countries for the 1970–2000 period. All labor indicators are normalized as specified in the paper.
b. Indicators of labor market rigidity are from Rama and Artecona (2002).
c. Minimum wages are normalized with the average labor cost in the manufacturing sectors.
d. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.
e. Indicators of labor regulations are from Djankov and others (2003).

The other two components of the aggregate industrial relations index (namely, collective bargaining and collective disputes) exhibit a positive correlation with income inequality. Finally, we find a negative association between social security laws and the Gini coefficient; this is the largest negative coefficient among the aggregate indices, at –0.38. Of the different benefits covered by social security laws, unemployment benefits display the largest negative correlation with the Gini coefficient (–0,47), while sickness and health benefits display the smallest correlation (–0.17). In summary, countries with more egalitarian distribution usually offer a better social security environment (with a legal framework that entails more benefits on old age, sickness, and unemployment than in other countries).

Table 4 reports the panel data correlation analysis between the Gini coefficient and the different indicators of labor market regulations from the Rama-Artecona database. We find that most of our indicators (both aggregate indices and individual categories) have an unconditional negative correlation with income inequality. The correlation coefficient between $\mathbf{L}_0$ and the Gini coefficient is –0.32, while the correlation between $\mathbf{L}_1$ and income inequality is higher than between $\mathbf{L}_2$ and income inequality (–0.47 and –0.20, respectively).[12]

The table also shows the evolution of the correlation between these variables over decades. The correlation between income inequality and de jure labor regulations (the $\mathbf{L}_0$ index) is negative in all decades, although it decreases from –0.34 in the 1970s to –0.30 in the 1990s. In the case of de facto regulations (as proxied by the aggregate $\mathbf{L}_1$ and $\mathbf{L}_2$ indices), the correlations decreased in the 1980s relative to the 1970s, but they then increased in the 1990s (although very slightly for $\mathbf{L}_1$). Finally, regulations on minimum wages (whether normalized by industrial wages or per capita income) are positively associated with income inequality for industrial countries (not shown in the table). For developing countries, we find a positive correlation only for minimum wages normalized by per capita income. Of course, we need to control for other determinants of inequality and possible reverse causation before we can properly conclude whether labor regulations affect inequality.

12. The largest negative correlation among the categories of the aggregate $\mathbf{L}_1$ index is with trade union membership (–0.50), followed by general government employment (–0.36) and social security contribution (–0.30). The smallest correlation is exhibited by minimum wages (–0.10). Days of maternity leave and trade union membership (as proxied by the ratification of ILO convention 87) show a negative correlation with the Gini coefficient among the $\mathbf{L}_2$ components (–0.31 and –0.18, respectively), while minimum wages and central government employment display a positive correlation (0.16 and 0.03, respectively).

Table 4. Panel Data Correlation Analysis for Labor Market Regulation and Income Inequality (Gini coefficient), 1970–2000[a]

	Full sample of countries				*Developing countries*			
Labor rigidity indicator	*1970-2000*	*1970s*	*1980s*	*1990s*	*1970-2000*	*1970s*	*1980s*	*1990s*
De jure index $\mathbf{L}_0$	−0.32	−0.34	−0.32	−0.30	−0.13	−0.18	−0.11	−0.11
De facto index $\mathbf{L}_1$	−0.47	−0.51	−0.45	−0.45	−0.43	−0.46	−0.39	−0.45
Minimum wage[b]	−0.10	−0.13	−0.06	−0.11	−0.18	−0.21	−0.10	−0.25
Social security contribution	−0.30	−0.23	−0.30	−0.34	−0.24	−0.17	−0.22	−0.28
Trade union membership	−0.50	−0.60	−0.50	−0.44	−0.46	−0.56	−0.45	−0.40
General government employment	−0.36	−0.38	−0.33	−0.38	−0.25	−0.23	−0.17	−0.35
De facto index $\mathbf{L}_2$	−0.20	−0.24	−0.15	−0.20	−0.17	−0.19	−0.07	−0.21
Minimum wage[c]	0.16	0.17	0.18	0.16	0.07	0.09	0.11	0.06
Maternity leave (no. days)	−0.31	−0.36	−0.32	−0.29	−0.34	−0.41	−0.34	−0.33
Ratification of ILO convention 87	−0.18	−0.19	−0.15	−0.17	−0.08	−0.09	−0.04	−0.10
Central government employment	0.03	0.09	0.07	−0.09	0.15	0.28	0.17	−0.00
De jure versus de facto								
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	−0.02	0.03	−0.01	−0.08	−0.26	−0.15	−0.26	−0.36
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.17	0.19	0.23	0.12	−0.05	−0.01	0.02	−0.13

Source: Authors' calculations, based on data from Rama and Artecona (2002).

a. Based on a panel data sample of 121 countries for the 1970–2000 period, in five-year nonoverlapping observations. The income inequality indicator is the Gini coefficient (0–1); indicators of labor market rigidity are from Rama and Artecona (2002).

b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.

c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.

3.3 Cross-section Regression Analysis

We first analyze the impact of labor regulations on income inequality for our cross-section of 121 countries over the 1970–2000 period. We start with our cross-section OLS estimates and then instrument for labor regulation in our simple IV estimates, where our dependent variable is the Gini coefficient. Table 5 presents the OLS and IV estimates including the three indicators of labor market regulations at the same time, both for the full sample and developing countries.[13] In this table, we report the coefficients, their standard errors and the R squared.[14] Tables B1 and B2 in appendix B present a similar exercise, but including only one labor market indicator at a time.

De jure regulations (the $\mathbf{L}_0$ index) do not seem to have a significant relationship with income inequality, regardless of the sample and estimation technique used. The $\mathbf{L}_1$ index of de facto regulations has a negative coefficient that is significant only for the OLS regression for developing countries. The $\mathbf{L}_2$ index has no significant association to the Gini coefficient. Using our IV estimates, we find that the following variables have a robust negative impact on the Gini coefficient across samples: the share of unionized labor, the share of general government employment, and the ratio of minimum wages to per capita income. Based on their estimated coefficients in table 5, we infer that a one-standard-deviation increase in trade union membership and public employment would reduce the Gini coefficient (0–1) by 0.094 and 0.082, respectively, while an analogous increase in the ratio of minimum wages to per capita income would increase income inequality by 0.15 over the thirty-year period. Finally, improving the ratio between $\mathbf{L}_1$ and $\mathbf{L}_0$, which serves as a measure of compliance, significantly improves income inequality in both samples.

13. Following the strategy applied by Calderón and Chong (in this volume), we base our choice of instruments for the labor market rigidity indicators on the literature summarized in Djankov and others (2003). Our main findings are that labor markets are more regulated in wealthy countries and in countries with a left-oriented government, while they are less regulated in countries with common law (British legal tradition). In addition, wealthy countries, more open countries, and countries with a British legal tradition have fewer labor regulations (proxied by employment laws, industrial relations laws, and social security laws). For the sake of brevity, we do not report the first-stage regression results; they are available on request.

14. The income inequality regression includes the following explanatory variables: output per capita (in logs), output per capita squared, secondary schooling, liquid liabilities, inflation, size of the modern sector, number of physicians per 1,000 people, and the different indicators of labor regulation. A full report of the regression results is available on request.

Table 5 also reports our results for the Djankov–La Porta indicators of labor regulations. We find that the aggregate index of employment laws has a positive and significant relationship with the Gini coefficient, regardless of the sample and estimation technique used. This positive relationship is mainly explained by regulations on alternative employment contracts. Industrial relations laws have a positive association with inequality, although it is only significant when we use IV. This effect on inequality is attributed to regulations on collective bargaining and collective disputes. Finally, social security laws also have a positive relation with inequality, which is significant only in the OLS estimations and is mainly attributed to the significance of regulations on sickness and health benefits. Economically speaking, a one-standard-deviation increase in the aggregate index of employment laws and industrial relations laws would increase the Gini coefficient by 0.02 over the thirty-year period (that is, the coefficient moves from an average of 0.39 for the full sample of countries to 0.37). An analogous increase in the regulations of both collective bargaining and disputes has a stronger negative impact on the distribution of income: the Gini coefficient increases by 0.04 and 0.10, respectively, over the thirty-year period.

3.4 Panel Data Regression Analysis

After performing our cross-section regression analysis, we evaluate the relation between labor market regulations and income inequality using a panel data set of nonoverlapping five-year observations for the 1970–2000 period. We take advantage of the additional dimension (that is, the time dimension) to draw some inferences on the impact of labor market regulations on income inequality with robust panel data estimation techniques.

Simple Techniques

We start by characterizing the relation between labor market regulations and income inequality using simpler techniques such as pooled OLS and OLS with time and country fixed effects. The pooled OLS does not take into account unobserved specific effects and endogeneity of the regressors. While the first problem can be accounted for by using time and country fixed effects, the second one is solved by including instrumental variables. We report estimates using IV with and without time and country fixed effects. In the next subsection, we

Table 5. Cross-country Regression Analysis for Labor Market Regulation and Income Inequality[a] (Dependent variable: Gini coefficient)

	Full sample of countries						*Developing countries*					
	Least squares			*Instrumental variables*			*Least squares*			*Instrumental variables*		
Variable	*Coefficient*	*Std. dev.*	R^2	*Coefficient*	*Std. dev.*	R^2	*Coefficient*	*Std. dev.*	R^2	*Coefficient*	*Std. dev.*	R^2
Labor market rigidity[b]												
De jure index $\mathbf{L}_0$	0.040	0.07	0.41	−0.008	0.18	0.41	0.084	0.10	0.23	0.047	0.21	0.22
De facto index $\mathbf{L}_1$	−0.123*	0.07	0.42	−0.125	0.23	0.41	−0.215*	0.11	0.24	−0.055	0.30	0.21
Minimum wage[c]	0.059	0.05	0.53	0.265	0.24	0.54	0.018	0.06	0.38	0.351	0.28	0.40
Social security contribution	−0.071*	0.04	0.42	0.107	0.16	0.42	−0.038	0.07	0.23	0.176	0.19	0.22
Trade union membership	−0.077	0.06	0.42	−0.421**	0.21	0.44	−0.144*	0.09	0.25	−0.399*	0.26	0.25
General government employment	−0.083*	0.05	0.44	−0.444*	0.25	0.44	−0.186**	0.08	0.30	−0.787**	0.38	0.29
De facto index $\mathbf{L}_2$	0.026	0.08	0.41	0.128	0.28	0.41	0.047	0.09	0.22	0.258	0.36	0.23
Minimum wage[d]	0.130	0.10	0.51	1.011**	0.33	0.54	0.118	0.11	0.36	1.623**	0.49	0.42
Maternity leave (no. days)	−0.023	0.08	0.41	−0.466	0.36	0.43	−0.138*	0.09	0.24	−1.372*	0.71	0.28
Ratification of ILO convention 87	−0.004	0.02	0.41	0.031	0.10	0.41	0.011	0.03	0.22	0.066	0.13	0.22
Central government employment	−0.069	0.09	0.39	−0.120	0.22	0.39	−0.109	0.10	0.20	0.078	0.37	0.21
De jure versus de facto												
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	−0.077*	0.05	0.42	−0.495**	0.23	0.44	−0.152*	0.08	0.25	−0.582**	0.29	0.26
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	−0.013	0.08	0.40	0.134	0.33	0.41	−0.014	0.11	0.22	0.107	0.41	0.22
Labor regulation[e]												
Employment laws (Djankov-La Porta)	0.054**	0.02	0.48	0.092*	0.05	0.48	0.084**	0.03	0.35	0.151**	0.06	0.36
Alternative employment contracts	0.105*	0.06	0.49	0.239	0.22	0.47	0.175**	0.08	0.38	0.479	0.34	0.34
Conditions of employment	0.046	0.06	0.48	0.185	0.13	0.50	0.062	0.10	0.35	0.282*	0.16	0.38
Job security	0.001	0.05	0.49	0.098	0.12	0.50	0.022	0.05	0.37	0.181	0.15	0.39

Table 5. (continued)

	Full sample of countries						*Developing countries*					
	Least squares			*Instrumental variables*			*Least squares*			*Instrumental variables*		
Variable	*Coefficient*	*Std. dev.*	R^2	*Coefficient*	*Std. dev.*	R^2	*Coefficient*	*Std. dev.*	R^2	*Coefficient*	*Std. dev.*	R^2
Industrial (collective) relations law	0.022	0.02	0.45	0.058*	0.03	0.48	0.031	0.03	0.26	0.096**	0.04	0.34
Collective bargaining	0.049	0.04	0.46	0.152*	0.08	0.50	0.071	0.05	0.28	0.234**	0.11	0.36
Worker participation in management	–0.021	0.03	0.48	–0.173	0.15	0.52	–0.012	0.04	0.29	0.064	0.15	0.29
Collective disputes	0.098*	0.06	0.47	0.602**	0.25	0.54	0.075	0.10	0.27	0.342**	0.15	0.38
Social security laws	0.043**	0.02	0.48	0.062	0.06	0.46	0.058**	0.03	0.34	0.107	0.08	0.29
Old age, disability, and death benefits	0.052	0.07	0.48	0.208	0.49	0.48	0.023	0.10	0.34	0.639	0.63	0.36
Sickness and health benefits	0.077**	0.04	0.50	0.277*	0.17	0.47	0.094**	0.04	0.37	0.208*	0.12	0.30
Unemployment benefits	–0.005	0.04	0.50	–0.103	0.20	0.47	0.014	0.04	0.36	0.014	0.21	0.37

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002) and Djankov and others (2003).
a. We report the regression coefficient for the indicator of labor rigidity according to equations 2 and 3 in the text. The dependent variable is the Gini coefficient (0–1). Our control variables are output per capita (in logs), output per capita squared, secondary schooling, liquid liabilities, inflation, size of the modern sector, physicians per 1,000 people, and the labor regulation indicator. Our set of instruments for the labor indicators consists of the level of development, trade openness adjusted by geographic variables, leftist political orientation of the government, British legal origin, German legal origin, and institutionalized autocracy; this set of instruments was chosen based on the existing literature, following Djankov and others (2003). Standard errors are robust to autocorrelation and heteroskedasticity (White, 1980). Full regression results and standard errors of the coefficients of the labor regulation variables are not reported for reasons of space, but they are available from the authors on request.
b. Indicators of labor market rigidity are from Rama and Artecona (2002).
c. Minimum wages are normalized with the average labor cost in the manufacturing sectors.
d. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.
e. Indicators of labor regulations are from Djankov and others (2003).

present estimates using the GMM-IV system estimator developed by Arellano and Bover (1995) and Blundell and Bond (1998), which takes unobserved effects and endogeneity into account using both internal and exogenous instruments for the labor regulation indicators. Since the latter method is our preferred estimation technique, we emphasize these estimates in our discussion of the results.

Tables 6 and 7 report the panel data estimation results for pooled and time effects, using both OLS and IV country techniques, respectively.

We focus our discussion on the IV estimates since, in principle, they tackle the endogeneity problem (see table 7). We find that de jure labor regulations have no significant relation with income inequality in almost all cases. However, $\mathbf{L}_0$ has a negative and significant impact on inequality for the world sample using our country-effects estimator. The $\mathbf{L}_1$ index has a negative and significant impact on inequality in developing countries when using country-effects estimator, while the $\mathbf{L}_2$ index has no significant impact on income distribution regardless of the sample. If we look at the components of the $\mathbf{L}_1$ index, the share of unionized labor and the size of public employment seem to drive down inequality among developing countries. When we analyze the components of the $\mathbf{L}_2$ index, we find that maternity leave and public employment have a negative and significant effect on the Gini coefficient in developing countries.

Our extensive regression analysis using OLS and IV estimates (pooled and with time and country fixed effects) of income inequality and aggregate de jure and de facto labor regulations indices ($\mathbf{L}_0$, $\mathbf{L}_1$, and $\mathbf{L}_2$) is presented in tables B3 through B8 in appendix B. Our specification includes other explanatory variables such as per capita output (in logs), per capita output squared, secondary schooling, liquid liabilities (as a percentage to GDP), the number of physicians per 1,000 people, the CPI inflation rate, and the size of the modern sector.[15]

15. In general, we find a nonlinear relation between income inequality and output per capita that is consistent with the Kuznets curve hypothesis (an inverted-U-shaped curve for the Gini coefficient). We also find that countries with a relatively more equal income distribution also have a greater stock of human capital, deeper financial systems, better health systems, lower macroeconomic instability, and a larger agricultural sector than do countries in which the income distribution is more skewed (see tables B3 through B8 in appendix B for further detail).

The GMM-IV System Estimator

The previous section used simple panel data techniques to characterize the relation between income inequality and labor market regulations. In this section, we introduce the GMM-IV system estimator proposed by Arellano and Bover (1995) and Blundell and Bond (1998). The GMM-IV system estimator is our preferred estimator for two main reasons. First, it accounts for unobserved country-specific effects that may bias our estimates. Specifically, we incorporate time dummies to control for the presence of time effects, and we take care of the country-specific effects by expressing our equation in differences. Second, the estimator controls for the possibility of endogenous regressors. We use both internal instruments (that is, lagged levels as instruments for the differences and lagged differences as instruments for the levels) and exogenous instruments for labor regulations suggested by theory (namely, legal and institutional variables). To confirm the validity of our income inequality regressions, we compute the Sargan test of overidentifying restrictions, which tests the validity of the moment conditions that we set up to perform the IV regressions, and tests of higher-order serial correlation.[16] These specification tests validate our regressions for statistical inference: our instruments are valid according to the Sargan test, and we reject the possibility of our errors displaying high-order serial correlation.

Before we discuss our results on the variable of interest (namely, labor market regulations), we briefly comment on the coefficient estimates for the other explanatory variables. First, we find evidence in favor of the Kuznets hypothesis that income inequality increases in the early stages of development and then decreases in later stages. On average, the turning point for GDP (in logs) in the full sample of countries is 8.1 (approximately the initial level of GDP per capita in Morocco during the 1996–2000 period), whereas the mean in the regression sample is 8.6 (Colombia during the same period). Second, a larger stock of human capital (as proxied by a larger enrollment rate in secondary education or a larger number of physicians per 1,000 people) may help reduce income inequality. Deeper financial systems also drive inequality down. Income inequality increases if the country features high inflation or a large modern sector, although the

16. By construction, our error term displays first-order serial correlation. For more technical details on the estimation technique, see Calderón y Chong (in this volume).

Table 6. Panel Data Regression Analysis for Labor Market Regulations and Income Inequality: Ordinary Least Squares[a] (Dependent variable: Gini coefficient)

Estimation method and labor regulation indicator	*Full sample*			*Developing countries*		
	Coefficient	*Std. dev.*	R^2	*Coefficient*	*Std. dev.*	R^2
Pooled estimator						
De jure index $\mathbf{L}_0$	0.022	0.03	0.38	0.067*	0.04	0.27
De facto index $\mathbf{L}_1$	–0.174**	0.03	0.41	–0.248**	0.05	0.30
Minimum wage[b]	–0.014	0.02	0.48	–0.063**	0.03	0.36
Social security contribution	–0.038**	0.02	0.39	–0.030	0.03	0.29
Trade union membership	–0.087**	0.03	0.42	–0.112**	0.04	0.31
General government employment	–0.049**	0.02	0.45	–0.076**	0.03	0.36
De facto index $\mathbf{L}_2$	–0.065**	0.03	0.38	–0.053	0.04	0.26
Minimum wage[c]	0.041	0.05	0.44	0.027	0.05	0.31
Maternity leave (no. days)	–0.090**	0.03	0.39	–0.121**	0.03	0.26
Ratification of ILO convention 87	–0.015*	0.01	0.38	–0.008	0.01	0.26
Central government employment	–0.024	0.03	0.43	–0.014	0.04	0.32
De jure versus de facto						
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	–0.084**	0.02	0.39	–0.142**	0.03	0.29
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	–0.059*	0.03	0.39	–0.083**	0.04	0.28
Time-effects estimator						
De jure index $\mathbf{L}_0$	0.024	0.03	0.41	0.055	0.04	0.30
De facto index $\mathbf{L}_1$	–0.159**	0.04	0.43	–0.231**	0.06	0.32
Minimum wage[b]	–0.017	0.02	0.49	–0.063**	0.03	0.37
Social security contribution	–0.043*	0.02	0.41	–0.037	0.03	0.32
Trade union membership	–0.064**	0.03	0.44	–0.084**	0.04	0.34
General government employment	–0.032	0.03	0.48	–0.055*	0.03	0.40
De facto index $\mathbf{L}_2$	–0.061**	0.03	0.41	–0.054*	0.03	0.29
Minimum wage[c]	0.023	0.04	0.47	0.011	0.05	0.33
Maternity leave (no. days)	–0.089**	0.04	0.41	–0.126**	0.06	0.30
Ratification of ILO convention 87	–0.014*	0.01	0.41	–0.010	0.01	0.29
Central government employment	–0.012	0.03	0.45	0.002	0.04	0.35

Table 6. (continued)

Estimation method and labor regulation indicator	*Full sample* Coefficient	*Std. dev.*	R^2	*Developing countries* Coefficient	*Std. dev.*	R^2
De jure versus de facto						
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	–0.076**	0.03	0.42	–0.124**	0.04	0.32
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	–0.055**	0.03	0.42	–0.074**	0.03	0.32
Country-effects estimator						
De jure index $\mathbf{L}_0$	–0.110*	0.06	0.91	–0.030	0.13	0.91
De facto index $\mathbf{L}_1$	0.162**	0.06	0.91	–0.360**	0.12	0.91
Minimum wage[b]	0.043	0.03	0.90	–0.269**	0.13	0.90
Social security contribution	0.083**	0.04	0.91	–0.357**	0.13	0.91
Trade union membership	0.071**	0.03	0.91	–0.318**	0.09	0.91
General government employment	–0.032	0.03	0.91	–0.462**	0.14	0.92
De facto index $\mathbf{L}_2$	0.126**	0.05	0.90	–0.364	0.27	0.90
Minimum wage[c]	–0.075	0.07	0.90	0.706**	0.32	0.91
Maternity leave (no. days)	0.128**	0.04	0.91	–0.677**	0.26	0.91
Ratification of ILO convention 87	0.039**	0.02	0.90	–0.056	0.07	0.90
Central government employment	–0.003	0.04	0.91	0.125**	0.05	0.91
De jure versus de facto						
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.190**	0.04	0.91	–0.489**	0.18	0.91
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.149**	0.04	0.91	–0.083	0.17	0.91

* Statistically significant at the 10 percent level.

** Statistically significant at the 5 percent level.

Source: Authors' calculations, based on data from Rama and Artecona (2002).

a. We report the regression coefficient for the indicator of labor rigidity according to equations (2) and (3) in the text. The dependent variable is the Gini coefficient (0–1). Our control variables are output per capita (in logs), output per capita squared, secondary schooling, liquid liabilities, inflation, size of the modern sector, physicians per 1,000 people, and the labor regulation indicator. Our set of instruments for the labor indicators consists of the level of development, trade openness adjusted by geographic variables, leftist political orientation of the government, British legal origin, German legal origin, and institutionalized autocracy; this set of instruments was chosen based on the existing literature, following Djankov and others (2003). Standard errors are robust to autocorrelation and heteroskedasticity (White, 1980). Full regression results and standard errors of the coefficients of the labor regulation variables are not reported for reasons of space, but they are available from the authors on request.

b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.

c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.

Table 7. Panel Data Regression Analysis for Labor Market Regulations and Income Inequality: Instrumental Variables[a] (Dependent variable: Gini coefficient)

Estimation method and labor regulation indicator	*Full sample*			*Developing countries*		
	Coefficient	*Std. dev.*	R^2	*Coefficient*	*Std. dev.*	R^2
Pooled estimator						
De jure index $\mathbf{L}_0$	0.033	0.07	0.40	0.102	0.09	0.30
De facto index $\mathbf{L}_1$	–0.103	0.08	0.39	0.023	0.12	0.28
Minimum wage[b]	0.075	0.10	0.49	0.105	0.12	0.38
Social security contribution	0.122**	0.06	0.39	0.151**	0.08	0.29
Trade union membership	–0.547**	0.14	0.42	–0.534**	0.17	0.31
General government employment	–0.368**	0.13	0.44	–0.681**	0.21	0.35
De facto index $\mathbf{L}_2$	0.091	0.10	0.39	0.181	0.13	0.29
Minimum wage[c]	0.282*	0.16	0.47	0.465**	0.21	0.37
Maternity leave (no. days)	–0.841**	0.29	0.41	–0.645*	0.36	0.30
Ratification of ILO convention 87	0.087*	0.05	0.40	0.051	0.05	0.29
Central government employment	–0.297**	0.13	0.43	–0.352*	0.21	0.33
De jure versus de facto						
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	–0.347**	0.11	0.41	–0.513**	0.14	0.32
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.051	0.12	0.40	–0.010	0.16	0.29
Time-effects estimator						
De jure index $\mathbf{L}_0$	0.015	0.07	0.42	0.055	0.08	0.33
De facto index $\mathbf{L}_1$	–0.130	0.10	0.42	–0.028	0.13	0.31
Minimum wage[b]	0.045	0.11	0.51	0.137	0.27	0.38
Social security contribution	0.108*	0.06	0.42	0.119*	0.07	0.32
Trade union membership	–0.557**	0.15	0.45	–0.539**	0.18	0.35
General government employment	–0.443**	0.13	0.48	–0.661**	0.21	0.39
De facto index $\mathbf{L}_2$	0.059	0.10	0.41	0.109	0.13	0.31
Minimum wage[c]	0.323*	0.17	0.49	0.430*	0.23	0.39
Maternity leave (no. days)	–0.880**	0.29	0.43	–0.761**	0.36	0.33
Ratification of ILO convention 87	0.106*	0.06	0.42	0.066	0.11	0.31
Central government employment	–0.391**	0.15	0.46	–0.400*	0.24	0.35

Table 7. (continued)

Estimation method and labor regulation indicator	*Full sample*			*Developing countries*		
	Coefficient	*Std. dev.*	R^2	*Coefficient*	*Std. dev.*	R^2
De jure versus de facto						
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	–0.352**	0.11	0.44	–0.440**	0.15	0.35
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.050	0.12	0.43	0.028	0.16	0.34
Country-effects estimator						
De jure index $\mathbf{L}_0$	–0.154**	0.07	0.89	0.007	0.15	0.89
De facto index $\mathbf{L}_1$	0.160**	0.07	0.89	–0.498**	0.16	0.90
Minimum wage[b]	0.054	0.04	0.88	–0.434**	0.18	0.88
Social security contribution	0.100**	0.05	0.89	–0.417**	0.15	0.89
Trade union membership	0.047	0.04	0.89	–0.449**	0.12	0.90
General government employment	–0.031	0.04	0.89	–0.738**	0.20	0.90
De facto index $\mathbf{L}_2$	0.159**	0.05	0.89	–0.143	0.31	0.88
Minimum wage[c]	–0.087	0.08	0.88	0.719*	0.39	0.88
Maternity leave (no. days)	0.158**	0.05	0.90	–0.826**	0.31	0.89
Ratification of ILO convention 87	0.043**	0.02	0.89	0.336**	0.13	0.89
Central government employment	0.125*	0.07	0.89	–0.895**	0.24	0.90
De jure versus de facto						
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	0.198**	0.05	0.90	–0.576**	0.20	0.90
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	0.170**	0.04	0.90	–0.077	0.20	0.89

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002).
a. We report the regression coefficient for the indicator of labor rigidity according to equations 2 and 3 in the text. The dependent variable is the Gini coefficient (0–1). Our control variables are output per capita (in logs), output per capita squared, secondary schooling, liquid liabilities, inflation, size of the modern sector, physicians per 1,000 people, and the labor regulation indicator. Our set of instruments for the labor indicators consists of the level of development, trade openness adjusted by geographic variables, leftist political orientation of the government, British legal origin, German legal origin, and institutionalized autocracy; this set of instruments was chosen based on the existing literature, following Djankov and others (2003). Standard errors are robust to autocorrelation and heteroskedasticity (White, 1980). Full regression results and standard errors of the coefficients of the labor regulation variables are not reported for reasons of space, but they are available from the authors on request.
b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.
c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.

Table 8. GMM-IV Panel Data Regression Analysis for Labor Market Regulations and Income Inequality: GMM-IV[a] (Dependent variable: Gini coefficient)

	Full sample of countries			*Developing countries*		
Explanatory variable	L_0	L_1	L_2	L_0	L_1	L_2
Constant	−0.548**	−0.865**	−1.115**	−0.580**	−1.130**	−2.134**
	(0.28)	(0.15)	(0.21)	(0.66)	(0.71)	(0.63)
Output per capita (logs)	0.214**	0.333**	0.364**	0.193**	0.377**	0.604**
	(0.07)	(0.04)	(0.05)	(0.17)	(0.19)	(0.16)
Output per capita squared	−0.013**	−0.020**	−0.023**	−0.012	−0.022**	−0.036**
	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)
Economic growth	−0.450**	−0.515**	−0.612**	−0.438**	−0.482**	−0.618**
	(0.06)	(0.05)	(0.04)	(0.07)	(0.08)	(0.11)
Secondary schooling	−0.018**	−0.008**	−0.019**	−0.035**	−0.058**	−0.040**
	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)
Liquid liabilities	−0.015**	−0.039**	−0.057**	−0.045**	−0.024	−0.077**
	(0.01)	(0.01)	(0.01)	(0.02)	(0.02)	(0.01)
Physicians per 1,000 people	−2.867**	0.556	−0.908**	−4.733**	0.451	−1.101
	(0.38)	(0.54)	(0.38)	(0.85)	(1.25)	(1.20)
Inflation rate	−0.002	−0.011**	−0.008	−0.005	−0.016*	−0.015*
	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)
Modern sector	0.201**	0.047	0.257**	0.351**	0.136	0.240**
	(0.04)	(0.04)	(0.05)	(0.14)	(0.13)	(0.11)
Labor rigidity	0.046**	−0.289**	−0.222**	0.103**	−0.291**	−0.205**
	(0.02)	(0.02)	(0.02)	(0.05)	(0.07)	(0.06)
Summary statistic						
No. countries	65	65	65	52	51	51
No. observations	182	199	200	146	156	157
R^2	0.42	0.38	0.42	0.34	0.31	0.29
Turning point	7.96	8.23	8.08	8.16	8.55	8.42

Table 8. (continued)

Explanatory variable	*Full sample of countries* L_0	L_1	L_2	*Developing countries* L_0	L_1	L_2
Specification tests (*p* values)						
Sargan test	0.85	0.70	0.86	0.85	0.80	0.86
Second-order correlation	0.71	0.99	0.91	0.63	0.96	0.91

* Statistically significant at the 10 percent level.

** Statistically significant at the 5 percent level.

Source: Authors' calculations, based on data from Rama and Artecona (2002).

a. The dependent variable is the Gini coefficient (0–1). The estimation method is the GMM-IV system estimator (Arellano and Bover, 1995), based on a panel data set of 121 countries over the 1970–2000 period, with nonoverlapping five-year observations. Asymptotic standard errors robust to general cross-section and time-series heteroskedasticity are reported in parentheses.

coefficient estimate of inflation is not robust (see table 8 for additional details).

We now turn to the effect of labor market regulations on income inequality. First, we find that de jure regulations, as proxied by the $\mathbf{L}_0$ index, have a positive and significant impact on the Gini coefficient for both the full sample of countries and the sample of developing countries. Hence, income inequality worsens with the adoption of an increasing number of ILO standards. A one-standard-deviation increase in the $\mathbf{L}_0$ index (namely, 0.21 for the full sample of countries) would reduce the Gini coefficient by 0.01, while an analogous increase in the $\mathbf{L}_0$ index for developing countries (that is, 0.18) would raise the Gini coefficient by 0.025. This increase in the $\mathbf{L}_0$ index (again, 0.21 for the full sample of countries) is much larger than the average increase observed from 1976–80 to 1996–2000 (0.06). Only Brazil, Finland, Spain, and Uruguay experienced such a large change over that period (that is, an increase of approximately 0.21 in the normalized number of ILO standards between 1976–80 and 1996–2000). We should take this result with caution, however, since reducing the number of regulations contained in the labor codes does not necessarily enhance the enforcement abilities of the regulators.

In contrast with our results for de jure regulations, we find that both the $\mathbf{L}_1$ and $\mathbf{L}_2$ indices of labor de facto regulations have a negative and significant coefficient estimate for the full sample of countries and the sample of developing countries. Labor market regulations should thus reduce income inequality in countries with a solid capability to enforce the law. A one-standard-deviation increase in the $\mathbf{L}_1$ index (or 0.13) may reduce income inequality by 0.037, while an analogous increase in the $\mathbf{L}_2$ index (or 0.15) may reduce the Gini coefficient by 0.033. An analogous increase in the extent of de facto regulations would cause a decline in the Gini coefficient between 0.028 (when $\mathbf{L}_1$ declines) and 0.032 (when $\mathbf{L}_2$ declines).[17]

Tables 9 and 10 report—for the full sample and the sample of developing countries, respectively—the sensitivity analysis of our coefficient estimates of labor regulations to changes in the indicator of labor regulation used in the regression (here we use the different

17. From 1976–80 to 1996–2000, the $\mathbf{L}_1$ index increased more than one standard deviation in Bangladesh, Jordan, and South Africa, whereas it decreased one standard deviation or more in Australia, Bulgaria, Israel, Syria, and the United Kingdom. The $\mathbf{L}_2$ index increased at least one standard deviation in Bangladesh, Romania, Turkey, and Venezuela, while it decreased one standard deviation or more in Bahrain, Niger, and New Zealand.

components of the aggregate indices used in table 8) and to changes in the proxy of income inequality used as our dependent variable. In addition to the Gini coefficient, we use the income share of selected quintiles of the population.

We first analyze the impact of the different individual measures of labor market regulations on the Gini coefficient. The negative impact of the $\mathbf{L}_1$ index on income inequality for the full sample of countries is mainly attributed to a negative and significant impact of social security contributions, trade union membership, and government employment. A one-standard-deviation increase in social security contributions reduces the Gini coefficient by 0.008, whereas analogous increases in trade union membership and public employment generate a decline in the Gini coefficient of 0.028 and 0.01, respectively. In the case of the negative impact of the $\mathbf{L}_2$ index, we find negative and significant effects on income inequality from maternity leave and trade union membership (as proxied by the ratification of the ILO convention on organized labor). A one-standard-deviation increase in mandated benefits (as proxied by a one-standard-deviation increase in the days of maternity leave) may reduce the Gini coefficient by 0.01. When we restrict our regression analysis to developing countries, mandated benefits—that is, social security contributions—drive the redistributive impact of the $\mathbf{L}_1$ index, whereas maternity leave and trade union membership drive the redistributive effects of the $\mathbf{L}_2$ index. The impact of a one-standard-deviation increase in mandated benefits among developing nations generates a reduction in the Gini coefficient of 0.012 regardless of the proxy used.

Next we analyze the impact of the different aggregate indices on the incomes share of the top, middle, and bottom quintiles of the population. Our index of de jure regulations, $\mathbf{L}_0$, has a positive but insignificant impact on the income shares of the top quintiles, but it has a negative and significant impact on the income share of the middle class (as proxied by the income share of the middle quintile) and the poor (as proxied by the share of the bottom quintile). A one-standard-deviation increase in the (normalized) number of ILO standards ratified would reduce the income share of the middle and bottom quintiles by 0.005 and 0.003, respectively. For the sample of developing countries, de jure regulations have a positive and significant relation with the income share of the top two quintiles and a negative and significant relationship with the middle and bottom quintiles. A one-standard-deviation increase in the $\mathbf{L}_0$ index raises the income share of the top two quintiles by 0.03, and it reduces the income share of the middle and bottom quintiles by 0.015 and 0.008, respectively.

Table 9. GMM-IV Panel Regressions for Labor Market Regulations and Income Inequality: Full Sample[a] (Dependent variables: alternative measures of income inequality)

			Population quintile										
	Gini		*Top quintile*		*Top two quintiles*		*Middle quintile*		*Bottom two quintiles*		*Bottom quintile*		
Labor market rigidity indicator	*Coeff.*	R^2	*Coeff.*	R^2	*Coeff.*	R^2	*Coeff.*	R^2	*Coeff.*	R^2	*Coeff.*	R^2	*No. obs.*
De jure index $\mathbf{L}_0$	0.046**	0.42	0.007	0.43	0.060	0.44	−0.026**	0.42	0.011	0.36	−0.017*	0.26	182
	(0.02)		(0.10)		(0.04)		(0.01)		(0.03)		(0.01)		
De facto index $\mathbf{L}_1$	−0.289**	0.38	−0.283**	0.40	−0.158**	0.41	0.041**	0.42	0.100**	0.31	0.036**	0.22	199
	(0.02)		(0.04)		(0.03)		(0.02)		(0.02)		(0.01)		
Minimum wage[b]	0.030	0.49	0.026	0.47	0.020	0.49	0.007	0.44	−0.032	0.41	−0.016*	0.32	198
	(0.02)		(0.03)		(0.04)		(0.01)		(0.02)		(0.01)		
Social security contribution	−0.038*	0.35	−0.034	0.38	−0.038**	0.38	0.014**	0.38	0.034**	0.29	0.019*	0.23	171
	(0.02)		(0.03)		(0.02)		(0.01)		(0.01)		(0.01)		
Trade union membership	−0.140**	0.42	0.037	0.37	0.027	0.37	−0.015	0.41	−0.034**	0.25	−0.021**	0.16	194
	(0.02)		(0.03)		(0.03)		(0.03)		(0.01)		(0.01)		
General government employment	−0.092**	0.48	−0.050*	0.43	−0.097**	0.43	0.012	0.46	0.056**	0.32	0.023	0.21	174
	(0.02)		(0.03)		(0.04)		(0.01)		(0.01)		(0.02)		
De facto index $\mathbf{L}_2$	−0.222**	0.42	−0.170**	0.42	−0.070**	0.41	0.016	0.41	0.040*	0.32	0.019**	0.21	200
	(0.02)		(0.03)		(0.03)		(0.03)		(0.02)		(0.01)		
Minimum wage[c]	−0.041	0.50	−0.024	0.52	−0.033	0.50	−0.007	0.43	0.001	0.44	−0.023	0.33	199
	(0.08)		(0.12)		(0.04)		(0.03)		(0.05)		(0.04)		
Maternity leave (no. days)	−0.049**	0.41	−0.121**	0.41	−0.085**	0.42	0.025**	0.43	0.061**	0.33	0.029**	0.22	175
	(0.02)		(0.04)		(0.03)		(0.01)		(0.03)		(0.01)		
Ratification of ILO convention 87	−0.018*	0.41	−0.017*	0.36	0.003	0.38	0.001	0.40	0.001	0.30	−0.005**	0.19	200
	(0.01)		(0.01)		(0.01)		(0.00)		(0.01)		(0.00)		
Central government employment	−0.048	0.45	−0.063	0.45	−0.007	0.45	0.005	0.45	0.013	0.34	−0.008	0.24	174
	(0.08)		(0.05)		(0.03)		(0.01)		(0.02)		(0.02)		

Table 9. (continued)

	Population quintile												
	Gini		*Top quintile*		*Top two quintiles*		*Middle quintile*		*Bottom two quintiles*		*Bottom quintile*		
Labor market rigidity indicator	*Coeff.*	*R^2*	*Coeff.*	*R^2*	*Coeff.*	*R^2*	*Coeff.*	*R^2*	*Coeff.*	*R^2*	*Coeff.*	*R^2*	*No. obs.*
De jure versus de facto													
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	−0.175**	0.35	−0.169**	0.43	−0.132**	0.41	0.074**	0.40	0.072**	0.33	0.037**	0.24	180
	(0.02)		(0.03)		(0.03)		(0.01)		(0.02)		(0.01)		
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	−0.065**	0.46	−0.114**	0.45	−0.085	0.46	0.001	0.46	0.048*	0.35	0.018*	0.24	181
	(0.02)		(0.04)		(0.06)		(0.01)		(0.03)		(0.01)		

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002).
a. The dependent variable is the Gini coefficient (0–1). The estimation method is the GMM-IV system estimator (Arellano and Bover, 1995), based on a panel data set of 121 countries over the 1970–2000 period, with nonoverlapping five-year observations. Indicators of labor market rigidity are from Rama and Artecona (2002). Asymptotic standard errors robust to general cross-section and time-series heteroskedasticity are reported in parentheses.
b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.
c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.

Table 10. GMM-IV Panel Regressions for Labor Market Regulations and Income Inequality: Developing Countries[a] (Dependent variables: alternative measures of income inequality)

			Population quintile										
	Gini		*Top quintile*		*Top two quintiles*		*Middle quintile*		*Bottom two quintiles*		*Bottom quintile*		
Labor market rigidity indicator	*Coeff.*	R^2	*Coeff.*	R^2	*Coeff.*	R^2	*Coeff.*	R^2	*Coeff.*	R^2	*Coeff.*	R^2	*No. obs.*
De jure index $\mathbf{L}_0$	0.103** (0.05)	0.34	0.140 (0.10)	0.21	0.170** (0.05)	0.33	−0.083** (0.02)	0.30	−0.099** (0.03)	0.28	−0.043** (0.02)	0.30	146
De facto index $\mathbf{L}_1$	−0.291** (0.07)	0.31	−0.415** (0.12)	0.23	−0.235** (0.06)	0.27	0.101** (0.04)	0.18	0.152** (0.05)	0.24	0.077** (0.02)	0.23	156
Minimum wage[b]	−0.071 (0.06)	0.34	−0.093 (0.06)	0.30	−0.057 (0.05)	0.36	−0.062** (0.03)	0.20	0.048 (0.04)	0.38	0.017 (0.02)	0.35	128
Social security contribution	−0.053* (0.03)	0.29	−0.094** (0.05)	0.20	−0.084* (0.05)	0.26	0.029 (0.02)	0.21	0.047* (0.03)	0.23	0.022** (0.01)	0.24	149
Trade union membership	−0.058 (0.04)	0.34	0.140** (0.06)	0.16	0.096** (0.04)	0.21	−0.037* (0.02)	0.18	−0.088 (0.09)	0.16	−0.023 (0.02)	0.17	151
General government employment	−0.063 (0.04)	0.35	−0.047 (0.07)	0.25	−0.022 (0.05)	0.33	0.027 (0.02)	0.34	−0.016 (0.04)	0.28	−0.006 (0.01)	0.27	131
De facto index $\mathbf{L}_2$	−0.205** (0.06)	0.29	−0.063 (0.05)	0.18	−0.029 (0.04)	0.29	0.040** (0.01)	0.24	0.017 (0.05)	0.24	0.005 (0.01)	0.24	157
Minimum wage[c]	0.038 (0.19)	0.34	0.239 (0.22)	0.26	0.185 (0.15)	0.37	−0.127* (0.08)	0.19	−0.115 (0.10)	0.37	−0.103** (0.04)	0.37	132
Maternity leave (no. days)	−0.104** (0.03)	0.29	−0.144** (0.04)	0.24	−0.110** (0.03)	0.33	0.035** (0.01)	0.26	0.090** (0.02)	0.29	0.033** (0.01)	0.30	147
Ratification of ILO convention 87	−0.028** (0.01)	0.34	0.014 (0.02)	0.17	0.017 (0.01)	0.29	0.003 (0.01)	0.25	−0.007 (0.01)	0.25	−0.007** (0.00)	0.24	157
Central government employment	0.053 (0.09)	0.33	−0.035 (0.07)	0.17	−0.040 (0.05)	0.31	0.004 (0.03)	0.30	0.023 (0.03)	0.21	−0.017 (0.03)	0.25	131

Table 10. (continued)

	Population quintile												
	Gini		*Top quintile*		*Top two quintiles*		*Middle quintile*		*Bottom two quintiles*		*Bottom quintile*		
Labor market rigidity indicator	*Coeff.*	R^2	*Coeff.*	R^2	*Coeff.*	R^2	*Coeff.*	R^2	*Coeff.*	R^2	*Coeff.*	R^2	*No. obs.*
De jure versus de facto													
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	−0.323**	0.30	−0.244**	0.26	−0.242**	0.35	0.059**	0.32	0.127**	0.29	0.059**	0.30	144
	(0.05)		(0.06)		(0.04)		(0.01)		(0.04)		(0.01)		
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	−0.163**	0.38	−0.249**	0.23	−0.178**	0.32	0.048**	0.29	0.101**	0.26	0.044**	0.25	145
	(0.02)		(0.05)		(0.02)		(0.01)		(0.02)		(0.01)		

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002).
a. The dependent variable is the Gini coefficient (0–1). The estimation method is the GMM-IV system estimator (Arellano and Bover, 1995), based on a panel data set of 121 countries over the 1970–2000 period, with nonoverlapping five-year observations. Indicators of labor market rigidity are from Rama and Artecona (2002). Asymptotic standard errors robust to general cross-section and time-series heteroskedasticity are reported in parentheses.
b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.
c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.

The $\mathbf{L}_1$ index has a negative and significant impact on the top shares and a positive and significant effect on the middle and bottom shares. Social security contribution is the dimension that reduces the income share of the top quintiles and increases the income share of the middle quintile. Specifically, a one-standard-deviation increase in the social security contributions (0.22) may help reduce the income share of the top quintiles by around 0.01, increase marginally the income share of the middle quintile by 0.003, and increase the income share of the bottom quintiles between 0.008 and 0.004. Active labor policies that raise public employment also work as an effective tool in raising the income share of the bottom quintiles of the population (although the economic impact is negligible). When we analyze the sample of developing countries, we find that the redistributive impact of $\mathbf{L}_1$ across income shares is mainly attributed to mandated benefits (as proxied by the social security contributions as a percentage of salaries). The redistributive effects of increased social security contributions are larger for the full sample of countries: a one-standard-deviation increase in social security contributions would reduce the shares of the top quintiles by 0.018 to 0.02, increase the middle quintile by 0.01, and raise the income share of the bottom quintiles by 0.004 to 0.011.

In addition, an increase in labor market regulations—approximated by a rise in the $\mathbf{L}_2$ index—would reduce the income shares of the top quintiles of the population and increase the income shares of the bottom quintiles. Its impact on the income share of the middle quintile is statistically negligible. The redistributive effects across income shares are basically attributed to mandated benefits (as proxied by the number of days of maternity leave). A one-standard-deviation increase in mandated benefits (that is, in maternity leave) would reduce the shares of the top quintiles by between 0.013 and 0.0171, increase the middle quintile by 0.004, and raise the income share of the bottom quintiles by between 0.005 and 0.01. The number of days of maternity leave (our proxy for mandated benefits) drives the redistributive effects of the $\mathbf{L}_2$ index in developing nations, which is consistent with our findings for the $\mathbf{L}_1$ index. The quantitative effects of increasing mandated benefits are similar to those found for the full sample of countries.

Finally, an increase in our measures of compliance (as proxied by a reduction in the gap between de jure and de facto regulations) significantly improves income inequality. This proposition holds for the full sample of countries when the gap is measured with the $\mathbf{L}_1$ index and for the sample of developing countries regardless of the measure of de facto regulations used. If the compliance in the extent of regulations in the labor markets improves (as proxied by a decrease in the gap between the $\mathbf{L}_0$ and $\mathbf{L}_1$ indices), the Gini coefficient would decrease by 0.03 (when using the full sample regressions) to 0.05 (when using the developing country regressions).

3.5 A Scorecard on the Redistributive Benefits of Labor Regulations

In similar fashion to Calderón and Chong (in this volume), we construct a scorecard to evaluate the redistributive benefits of labor market regulations for the full sample of countries and for the sample of developing countries. The scorecard assesses the relation between our indicators of labor regulations and inequality measures such as the Gini coefficient and the income shares of the top, middle and bottom 20 percent of the population. We summarize the information from our different panel estimations by inputting the value of –1 for a negative and significant coefficient estimate, +1 for a positive and significant coefficient estimate, and 0 for an insignificant coefficient. The proportion of these negative and positive coefficients is presented in table 11. Our discussion of the summary results centers on the full sample of countries.

Regarding the relation between labor regulations and the Gini coefficient, we find, first, that de jure regulations have a positive, but weak correlation with income inequality. Second, de facto regulations—measured by either the $\mathbf{L}_1$ or $\mathbf{L}_2$ aggregate index—have a negative association with income inequality. The robust relation between the $\mathbf{L}_1$ index and the Gini coefficient may be attributed to the redistributive effects of both trade union membership and public employment. Mandated benefits (as proxied by the number of days of maternity leave) seem to explain the robust relation between the $\mathbf{L}_2$ index and the Gini coefficient. Finally, our two measures of enforcement of labor regulations seem to have a negative and robust relation with the Gini coefficient.

The aggregate $\mathbf{L}_1$ index of de facto labor regulations is negatively associated with the income share of the top 20 percent of the population and positively associated with the income shares of bottom and middle quintiles. The negative relation between the $\mathbf{L}_1$ index and the income share of the top quintile may be explained by the robust negative relation with trade union membership and public employment. The positive correlation between the $\mathbf{L}_1$ index and the income share of the bottom quintile may be explained by social security contributions. Finally, the aggregate $\mathbf{L}_2$ index of de facto labor regulations

Table 11. A Scorecard of Labor Regulations and Income Inequality[a]

	Full sample of countries				Developing countries			
		Quintile				Quintile		
Labor rigidity indicator	*Gini*	*Top*	*Middle*	*Bottom*	*Gini*	*Top*	*Middle*	*Bottom*
De jure index $\mathbf{L}_0$	0.2	0.0	–0.6	–0.2	0.4	0.0	–0.6	–0.4
De facto index $\mathbf{L}_1$	–0.6	–1.0	0.6	0.6	–0.6	–0.6	0.6	0.4
Minimun wage[b]	0.0	–0.2	0.4	–0.2	–0.4	–0.4	0.2	0.0
Social security contribution	–0.2	–0.4	0.0	0.6	0.2	–0.6	–0.2	0.4
Trade union membership	–1.0	–0.8	0.4	0.4	–0.8	–0.8	0.2	0.6
General government employment	–0.8	–1.0	0.4	0.4	–0.8	–0.8	0.4	0.4
De facto index $\mathbf{L}_2$	–0.6	–0.6	0.4	0.2	–0.4	–0.2	0.2	0.0
Minimun wage[c]	0.4	0.4	0.0	0.0	0.4	0.4	–0.6	–0.2
Maternity leave (no. days)	–1.0	–1.0	0.6	1.0	–1.0	–1.0	0.8	0.8
Ratification of ILO convention 87	–0.2	–0.6	0.0	–0.2	–0.2	0.0	0.0	–0.2
Central government employment	–0.4	–0.2	0.0	0.0	–0.4	0.0	0.0	0.0
De jure versus de facto								
$\mathbf{L}_1$ relative to $\mathbf{L}_0$	–1.0	–0.6	0.6	0.8	–1.0	–1.0	1.0	0.8
$\mathbf{L}_2$ relative to $\mathbf{L}_0$	–0.6	0.0	0.4	0.2	–0.6	–0.6	0.6	0.2

Source: Authors' calculations, based on data from Rama and Artecona (2002).
a. Based on five different panel data, we input the value of –1 for a negative and significant coefficient estimate, +1 for a positive and significant coefficient estimate, and 0 for insignificant coefficients; the table reports the proportion of significant negative and positive coefficients.
b. Minimum wages are normalized with the average labor cost in the manufacturing sectors.
c. Minimum wages are normalized with real income per capita. All labor indicators are normalized as specified in the text.

has a robust negative relation with the income share of the top quintile of the population and a positive, but weak association with the income share of both the middle and bottom quintiles. The negative robust association with the income share of the top quintile may be attributed to mandated benefits (proxied by maternity leave rights) and trade union membership.

4. Conclusions

We have analyzed the relationship between labor regulations and income inequality. Finding robust results is not a straightforward process, however, because there are alternative ways of measuring regulations and alternative estimation techniques for addressing (albeit imperfectly) simultaneity and probable measurement errors. We thus used alternative econometric approaches and considered two data sets and two alternative samples. A number of results appear to be fairly robust.

The main results in our paper can be grouped in three types. First, we find that de jure regulations do not improve income distribution. The Rama-Artecona indicator (the $\mathbf{L}_0$ index) does not display any consistent pattern, and the Djankov-La Porta indicators either have no effect or worsen income distribution. Second, relative compliance with existing regulations, particularly the ratio between the $\mathbf{L}_1$ and $\mathbf{L}_0$ indices of the Rama-Artecona data set, seems to improve income distribution, although we cannot rule out the possibility that this measure is proxying for other factors such as institutional development. Third, de facto regulations are weakly associated, overall, with improving income inequality. This result is partly due to the fact that different regulations have quite distinct effects. In particular, we find that a higher minimum wage tends to worsen income distribution, whereas the extent of trade union membership, the importance of government employment and maternity leave improve it. Finally, some of these positive results do not carry through to the bottom quintile of the population.

Appendix A
List of Countries

— Industrial countries (twenty-two countries): Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Luxembourg, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, United Kingdom, and United States.

— Latin America and the Caribbean (twenty-one countries): Argentina, Bahamas, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Guatemala, Honduras, Jamaica, Mexico, Nicaragua, Panama, Paraguay, Peru, Trinidad and Tobago, Uruguay, and Venezuela.

— East Asia and the Pacific (twelve countries): China, Hong Kong, Indonesia, Korea, Malaysia, Mongolia, Papua New Guinea, Philippines, Singapore, Taiwan, Thailand, and Vietnam.

— Eastern Europe and Central Asia (seventeen countries): Belarus, Bulgaria, Croatia, Czech Republic, Estonia, Hungary, Kazakhstan, Kyrgyz Republic, Latvia, Lithuania, Poland, Romania, Russian Federation, Slovak Republic, Slovenia, Ukraine, and Yugoslavia.

— Middle East and North Africa (twenty-one countries): Algeria, Bahrain, Cyprus, Egypt, Iran, Iraq, Israel, Jordan, Kuwait, Lebanon, Libya, Malta, Morocco, Oman, Qatar, Saudi Arabia, Syria, Tunisia, Turkey, United Arab Emirates, and Yemen.

— South Asia (five countries): Bangladesh, India, Nepal, Pakistan, and Sri Lanka.

— Sub-Saharan Africa (twenty-three countries): Botswana, Burkina Faso, Côte d'Ivoire, Ethiopia, Ghana, Guinea, Guinea-Bissau, Kenya, Lesotho, Madagascar, Mali, Mauritania, Mauritius, Niger, Nigeria, Rwanda, Senegal, Sierra Leone, South Africa, Tanzania, Uganda, Zambia, and Zimbabwe.

Appendix B
Supplemental Tables

In addition to the exercises reported in this appendix, we performed cross-country regression analysis between income inequality and labor market regulations using income shares as a proxy for our dependent variable, for both the Rama-Artecona and Djankov-La Porta databases. We also carried out sensitivity analyses on panel regressions for different measures of labor regulations for the full sample of countries and the sample of developing countries, using OLS and IV with pooled and time-effects estimators. These results are available on request.

Table B1. Cross-section Regression Analysis for Labor Market Regulations and Income Inequality: Ordinary Least Squares[a] (Dependent variable: Gini coefficient)

	Full sample			*Developing countries*			*Full sample*			*Developing countries*		
Explanatory variable	L_0	L_1	L_2	L_0	L_1	L_2	EL_0	IR_0	SS_0	EL_0	IR_0	SS_0
Constant	0.362	–0.166	0.354	0.763	–0.375	0.700	0.235	0.460	0.188	0.534	0.866	0.373
	(0.88)	(0.85)	(0.94)	(1.08)	(1.05)	(1.16)	(1.04)	(1.05)	(0.86)	(1.23)	(1.25)	(1.02)
Output per capita (in logs)	0.125**	0.118**	0.119**	0.143**	0.175**	0.148**	0.057**	0.056**	0.058**	0.110**	0.170**	0.163**
	(0.06)	(0.06)	(0.05)	(0.07)	(0.07)	(0.06)	(0.03)	(0.03)	(0.02)	(0.03)	(0.03)	(0.03)
Output per capita squared	–0.008**	–0.007**	–0.007**	–0.009**	–0.010**	–0.009**	–0.003**	–0.003**	–0.003**	–0.008**	–0.011**	–0.010**
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Economic growth	–0.958*	–0.911*	–1.016*	–0.766*	–0.771*	–0.779*	–1.692**	–1.694**	–1.741**	–0.911**	–0.804**	–1.482**
	(0.62)	(0.60)	(0.62)	(0.47)	(0.48)	(0.48)	(0.79)	(0.83)	(0.71)	(0.45)	(0.39)	(0.71)
Secondary schooling	–0.020*	–0.020*	–0.021*	–0.028*	–0.034*	–0.027*	–0.016*	–0.020*	–0.019*	–0.018	–0.035	–0.047
	(0.01)	(0.01)	(0.01)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)	(0.04)	(0.04)	(0.04)
Liquid liabilities	–0.015	–0.023	–0.019	–0.007	–0.013	–0.010	–0.002	–0.013	–0.001	0.033	0.001	0.010
	(0.03)	(0.03)	(0.02)	(0.04)	(0.04)	(0.04)	(0.03)	(0.03)	(0.02)	(0.04)	(0.05)	(0.04)
Inflation rate	0.079**	0.076**	0.080**	0.069*	0.072*	0.085**	0.055*	0.064*	0.078*	0.049*	0.058*	0.088**
	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.03)	(0.04)	(0.03)	(0.03)	(0.04)
Modern sector	0.294*	0.274*	0.295*	0.285*	0.279*	0.289*	0.265*	0.299*	0.262*	0.261	0.312*	0.216
	(0.16)	(0.15)	(0.16)	(0.16)	(0.16)	(0.16)	(0.17)	(0.19)	(0.17)	(0.18)	(0.19)	(0.17)
Physicians per 1,000 people	–6.117**	–4.222**	–5.461**	–6.550**	–5.486**	–5.887**	–6.722**	–6.569**	–7.964**	–7.704**	–6.712**	–9.537**
	(2.17)	(1.55)	(2.00)	(2.68)	(2.43)	(2.53)	(1.99)	(1.91)	(2.06)	(2.40)	(2.49)	(2.50)
Labor regulation (L_0, L_1, L_2)	0.040	–0.123*	0.026	0.084	–0.215*	0.047	0.054**	0.022	0.043**	0.084**	0.031	0.058**
	(0.07)	(0.07)	(0.08)	(0.10)	(0.11)	(0.09)	(0.02)	(0.02)	(0.02)	(0.03)	(0.03)	(0.03)

Table B1. (continued)

	Full sample			*Developing countries*			*Full sample*			*Developing countries*		
Explanatory variable	$\mathbf{L}_0$	$\mathbf{L}_1$	$\mathbf{L}_2$	$\mathbf{L}_0$	$\mathbf{L}_1$	$\mathbf{L}_2$	$\mathbf{EL}_0$	$\mathbf{IR}_0$	$\mathbf{SS}_0$	$\mathbf{EL}_0$	$\mathbf{IR}_0$	$\mathbf{SS}_0$
Summary statistic												
No. observations	68	67	68	53	52	53	53	53	53	38	38	38
R^2	0.41	0.41	0.40	0.22	0.24	0.22	0.48	0.45	0.48	0.34	0.26	0.34
Turning point	8.0	8.6	8.1	8.4	8.6	8.5	8.5	8.3	8.6	7.2	7.8	8.0

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002) and Djankov and others (2003).
a. The dependent variable is the Gini coefficient (0–1). The $\mathbf{L}_0$, $\mathbf{L}_1$, and $\mathbf{L}_2$ indices are the Rama-Artecona aggregate de jure and de facto labor rigidity indices. The $\mathbf{EL}_0$, $\mathbf{IR}_0$, and $\mathbf{SS}_0$ indices are the Djankov-La Porta aggregate indices of employment laws, industrial relations laws, and social security laws, respectively. Standard errors are in parentheses.

Table B2. Cross-section Regression Analysis for Labor Regulation and Income Inequality: Instrumental Variables[a] (Dependent variable: Gini coefficient)

	Full sample			Developing countries			Full sample			Developing countries		
Explanatory variable	L_0	L_1	L_2	L_0	L_1	L_2	EL_0	IR_0	SS_0	EL_0	IR_0	SS_0
Constant	0.409	–0.039	0.495	0.609	0.062	0.863	0.372	0.402	0.354	1.170	0.890	0.863
	(0.87)	(0.86)	(0.92)	(1.08)	(1.05)	(1.18)	(1.08)	(1.08)	(1.06)	(1.26)	(1.22)	(1.25)
Output per capita (in logs)	0.077**	0.079**	0.059**	0.049*	0.052*	0.063*	0.068**	0.052**	0.060**	0.317**	0.295**	0.304**
	(0.02)	(0.02)	(0.02)	(0.03)	(0.03)	(0.03)	(0.03)	(0.02)	(0.03)	(0.03)	(0.03)	(0.03)
Output per capita squared	–0.004**	–0.004**	–0.003**	–0.003**	–0.002**	–0.003**	–0.004*	–0.003*	–0.004*	–0.018*	–0.017*	–0.018*
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)
Economic growth	–1.001**	–1.048**	–0.966**	–0.779**	–0.905**	–0.719**	–1.159**	–1.188**	–1.096**	–1.154**	–1.552**	–1.060**
	(0.17)	(0.17)	(0.18)	(0.09)	(0.08)	(0.09)	(0.09)	(0.09)	(0.09)	(0.21)	(0.27)	(0.22)
Secondary schooling	–0.021*	–0.023*	–0.022*	–0.028*	–0.027*	–0.024*	–0.012*	–0.017*	–0.018*	–0.027*	–0.037*	–0.029*
	(0.01)	(0.01)	(0.01)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.02)
Liquid liabilities	–0.033	–0.035	–0.028	–0.030	–0.018	–0.030	–0.002	–0.013	–0.010	–0.072	–0.103*	–0.052
	(0.04)	(0.04)	(0.03)	(0.06)	(0.06)	(0.06)	(0.03)	(0.03)	(0.03)	(0.06)	(0.06)	(0.06)
Inflation rate	0.078**	0.074*	0.079**	0.077**	0.077**	0.078**	0.060	0.045	0.058	0.065*	0.051	0.061
	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)
Modern sector	0.300*	0.251*	0.298*	0.304*	0.302*	0.305*	0.275*	0.302*	0.276*	0.278*	0.302*	0.255*
	(0.16)	(0.16)	(0.16)	(0.16)	(0.17)	(0.16)	(0.18)	(0.18)	(0.18)	(0.16)	(0.17)	(0.17)
Physicians per 1,000 people	–5.332**	–4.734**	–5.768**	–5.675**	–4.788**	–6.428**	–7.813**	–7.463**	–6.840**	–9.743**	–9.468**	–8.566**
	(1.93)	(1.76)	(2.12)	(2.56)	(2.41)	(2.76)	(2.14)	(2.11)	(2.22)	(2.44)	(2.36)	(2.80)
Labor regulation (L_0, L_1, L_2)	–0.008	–0.125	0.128	0.047	–0.055	0.258	0.092*	0.058*	0.062	0.151**	0.096**	0.107
	(0.18)	(0.23)	(0.28)	(0.21)	(0.30)	(0.36)	(0.05)	(0.03)	(0.06)	(0.06)	(0.04)	(0.08)

Table B2. (continued)

	Full sample			*Developing countries*			*Full sample*			*Developing countries*		
Explanatory variable	L_0	L_1	L_2	L_0	L_1	L_2	EL_0	IR_0	SS_0	EL_0	IR_0	SS_0
Summary statistic												
No. observations	66	65	66	51	50	51	51	51	51	36	36	36
R^2	0.407	0.409	0.409	0.216	0.210	0.225	0.482	0.479	0.456	0.359	0.342	0.285
Turning point	9.6	9.4	9.6	9.7	10.8	10.6	8.6	8.3	8.4	8.8	8.5	8.6

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002) and Djankov, and others (2003).
a. The dependent variable is the Gini coefficient (0–1). The $\mathbf{L}_0$, $\mathbf{L}_1$, and $\mathbf{L}_2$ indices are the Rama-Artecona aggregate de jure and de facto labor rigidity indices. The $\mathbf{EL}_0$, $\mathbf{IR}_0$, and $\mathbf{SS}_0$ indices are the Djankov-La Porta aggregate indices of employment laws, industrial relations laws, and social security laws, respectively. Our set of instruments for the labor indicators consists of the level of development, trade openness adjusted by geographic variables, leftist political orientation of the government, British legal origin, German legal origin, and institutionalized autocracy; the set of instruments was chosen from the existing literature, following Djankov and others (2003). Standard errors are in parentheses.

Table B3. The Impact of De Jure Regulations: Panel Data Regression Analysis with Least Squares[a] (Dependent variable: Gini coefficient)

	Full sample			Developing countries		
Explanatory variable	*Pooled*	*Time fixed effects*	*Country fixed effects*	*Pooled*	*Time fixed effects*	*Country fixed effects*
Output per capita (in logs)	0.180**	0.181**	0.068	0.200**	0.255**	0.138
	(0.08)	(0.09)	(0.12)	(0.10)	(0.11)	(0.15)
Output per capita squared	–0.011**	–0.010**	–0.004	–0.012*	–0.014**	–0.008
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Economic growth	–0.229	–0.164	0.135	–0.143	–0.060	0.172
	(0.16)	(0.15)	(0.10)	(0.17)	(0.17)	(0.13)
Secondary schooling	–0.021**	–0.027**	–0.018**	–0.027**	–0.039**	–0.031**
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Liquid liabilities	–0.040**	–0.050**	0.026	–0.047**	–0.048**	0.026
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.03)
Physicians per 1,000 people	–3.773**	–4.521**	1.260*	–5.565**	–6.157**	2.331*
	(0.84)	(0.90)	(0.76)	(1.02)	(1.12)	(1.29)
Inflation	0.022	0.026*	–0.011	0.022	0.034*	–0.013
	(0.02)	(0.02)	(0.01)	(0.02)	(0.02)	(0.01)
Size of the modern sector	0.294**	0.257**	–0.088	0.294**	0.263**	–0.075
	(0.06)	(0.07)	(0.08)	(0.06)	(0.07)	(0.09)
Labor rigidity indicator (L_0)	0.022	0.024	–0.110*	0.067*	0.055	–0.154*
	(0.03)	(0.03)	(0.06)	(0.04)	(0.04)	(0.09)
Summary statistic						
No. observations	327	327	327	263	263	263
R^2	0.378	0.410	0.908	0.267	0.303	0.892
Adjusted R^2	0.361	0.383	0.847	0.241	0.263	0.787
GDP turning point	7.97	8.68	8.08	8.52	8.87	8.46

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002).
a. We report the regression coefficient for the indicator of labor rigidity according to equations (2) and (3) in the text. The dependent variable is the Gini coefficient (0–1). Our control variables are output per capita (in logs), output per capita squared, secondary schooling, liquid liabilities, inflation, size of the modern sector, physicians per 1,000 people, and the labor regulation indicator. The pooled regressions include a constant. Our set of instruments for the labor indicators consists of the level of development, trade openness adjusted by geographic variables, leftist political orientation of the government, British legal origin, German legal origin, and institutionalized autocracy; this set of instruments was chosen based on the existing literature, following Djankov and others (2003). Standard errors are in parentheses and are robust to autocorrelation and heteroskedasticity (White, 1980). Full regression results and standard errors of the coefficients of the labor regulation variables are not reported for reasons of space, but they are available from the authors on request.

Table B4. The Impact of *De Jure* Regulations: Panel Data Regression Analysis with Instrumental Variables[a]

Explanatory variable	*Full sample* Pooled	*Full sample* Time fixed effects	*Full sample* Country fixed effects	*Developing countries* Pooled	*Developing countries* Time fixed effects	*Developing countries* Country fixed effects
Output per capita (in logs)	0.221**	0.214**	0.112	0.372**	0.455**	0.232
	(0.08)	(0.09)	(0.14)	(0.12)	(0.14)	(0.19)
Output per capita squared	–0.014**	–0.013**	–0.007	–0.024**	–0.028**	–0.014
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Economic growth	–0.338**	–0.282*	0.116	–0.247	–0.158	0.143
	(0.17)	(0.16)	(0.11)	(0.17)	(0.18)	(0.14)
Secondary schooling	–0.019**	–0.025**	–0.020**	–0.025**	–0.039**	–0.033**
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Liquid liabilities	–0.048**	–0.056**	0.025	–0.067**	–0.074**	0.026
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.03)
Physicians per 1,000 people	–3.117**	–3.785**	0.741	–4.359**	–4.832**	1.145
	(0.82)	(0.85)	(0.72)	(1.04)	(1.07)	(1.23)
Inflation	0.018	0.022	–0.010	0.022	0.033*	–0.011
	(0.02)	(0.02)	(0.01)	(0.02)	(0.02)	(0.01)
Size of the modern sector	0.303**	0.268**	–0.090	0.316**	0.278**	–0.065
	(0.06)	(0.07)	(0.08)	(0.06)	(0.07)	(0.10)
Labor rigidity indicator (L_0)	0.033	0.015	–0.030	0.102	0.055	0.007
	(0.07)	(0.07)	(0.15)	(0.09)	(0.08)	(0.19)
Summary statistic						
No. observations	312	312	312	248	248	248
R^2	0.396	0.425	0.906	0.296	0.332	0.889
Adjusted R^2	0.378	0.398	0.840	0.269	0.292	0.769
GDP turning point	7.78	8.40	8.23	7.84	8.20	8.25

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002).
a. We report the regression coefficient for the indicator of labor rigidity according to equations (2) and (3) in the text. The dependent variable is the Gini coefficient (0–1). Our control variables are output per capita (in logs), output per capita squared, secondary schooling, liquid liabilities, inflation, size of the modern sector, physicians per 1,000 people, and the labor regulation indicator. The pooled regressions include a constant. Our set of instruments for the labor indicators consists of the level of development, trade openness adjusted by geographic variables, leftist political orientation of the government, British legal origin, German legal origin, and institutionalized autocracy; this set of instruments was chosen based on the existing literature, following Djankov and others (2003). Standard errors are in parentheses and are robust to autocorrelation and heteroskedasticity (White, 1980). Full regression results and standard errors of the coefficients of the labor regulation variables are not reported for reasons of space, but they are available from the authors on request.

Table B5. The Impact of the L_1 Index of De Facto Regulations: Panel Data Regression Analysis with Least Squares[a] (Dependent variable: Gini coefficient)

	Full sample			*Developing countries*		
Explanatory variable	*Pooled*	*Time fixed effects*	*Country fixed effects*	*Pooled*	*Time fixed effects*	*Country fixed effects*
Output per capita (in logs)	0.217**	0.213**	0.049	0.279**	0.311**	0.071
	(0.08)	(0.08)	(0.11)	(0.09)	(0.11)	(0.15)
Output per capita squared	−0.013**	−0.012**	−0.003	−0.016**	−0.017**	−0.004
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Economic growth	−0.159**	−0.116	0.129	−0.072	−0.034	0.170
	(0.16)	(0.14)	(0.10)	(0.16)	(0.16)	(0.12)
Secondary schooling	−0.025**	−0.029**	−0.023**	−0.034**	−0.043**	−0.035**
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Liquid liabilities	−0.058**	−0.065**	0.035*	−0.064**	−0.062**	0.031
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.03)
Physicians per 1,000 people	−1.883**	−2.590**	0.780	−1.852**	−2.582**	1.473
	(0.55)	(0.80)	(0.66)	(0.83)	(1.11)	(1.11)
Inflation	0.018	0.022	−0.015	0.019	0.029*	−0.015
	(0.02)	(0.02)	(0.01)	(0.02)	(0.02)	(0.01)
Size of the modern sector	0.219**	0.191**	−0.156**	0.205**	0.187**	−0.144*
	(0.06)	(0.06)	(0.07)	(0.06)	(0.07)	(0.08)
Labor rigidity indicator (L_1)	−0.174**	−0.159**	0.162**	−0.248**	−0.231**	0.160*
	(0.03)	(0.04)	(0.06)	(0.05)	(0.06)	(0.09)
Summary statistic						
No. observations	341	341	341	269	269	269
R^2	0.41	0.43	0.91	0.30	0.32	0.89
GDP turning point	8.54	9.07	8.61	8.85	9.03	9.66

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002).
a. We report the regression coefficient for the indicator of labor rigidity according to equations (2) and (3) in the text. The dependent variable is the Gini coefficient (0–1). Our control variables are output per capita (in logs), output per capita squared, secondary schooling, liquid liabilities, inflation, size of the modern sector, physicians per 1,000 people, and the labor regulation indicator. The pooled regressions include a constant. Our set of instruments for the labor indicators consists of the level of development, trade openness adjusted by geographic variables, leftist political orientation of the government, British legal origin, German legal origin, and institutionalized autocracy; this set of instruments was chosen based on the existing literature, following Djankov and others (2003). Standard errors are in parentheses and are robust to autocorrelation and heteroskedasticity (White, 1980). Full regression results and standard errors of the coefficients of the labor regulation variables are not reported for reasons of space, but they are available from the authors on request.

Table B6. The Impact of the L_1 Index of De Facto Regulations: Panel Data Regression Analysis with Instrumental Variables[a] (Dependent variable: Gini coefficient)

	Full sample			*Developing countries*		
Explanatory variable	*Pooled*	*Time fixed effects*	*Country fixed effects*	*Pooled*	*Time fixed effects*	*Country fixed effects*
Output per capita (in logs)	0.277**	0.262**	0.161	0.423**	0.474**	0.252
	(0.08)	(0.09)	(0.12)	(0.11)	(0.14)	(0.17)
Output per capita squared	−0.017**	−0.015**	−0.008	−0.026**	−0.028**	−0.013
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Economic growth	−0.284**	−0.234*	0.068	−0.165	−0.111	0.077
	(0.16)	(0.16)	(0.10)	(0.17)	(0.18)	(0.13)
Secondary schooling	−0.024**	−0.030**	−0.017**	−0.030**	−0.042**	−0.031**
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Liquid liabilities	−0.060**	−0.069**	0.025	−0.079**	−0.081**	0.026
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.03)
Physicians per 1,000 people	−2.503**	−3.147**	0.780	−3.765**	−4.253**	1.297
	(0.73)	(0.81)	(0.68)	(0.96)	(1.103)	(1.15)
Inflation	0.018	0.022	−0.013	0.022	0.033*	−0.012
	(0.02)	(0.02)	(0.01)	(0.02)	(0.02)	(0.01)
Size of the modern sector	0.253**	0.222**	−0.174**	0.253**	0.231**	−0.155*
	(0.06)	(0.07)	(0.08)	(0.06)	(0.07)	(0.09)
Labor rigidity indicator (L_1)	−0.103	−0.130	−0.360**	0.023	−0.028	−0.498*
	(0.08)	(0.10)	(0.14)	(0.12)	(0.13)	(0.20)
Summary statistic						
No. observations	326	326	326	254	254	254
R^2	0.39	0.42	0.91	0.28	0.31	0.90
GDP turning point	8.36	8.94	9.93	8.21	8.44	9.95

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002).
a. We report the regression coefficient for the indicator of labor rigidity according to equations (2) and (3) in the text. The dependent variable is the Gini coefficient (0–1). Our control variables are output per capita (in logs), output per capita squared, secondary schooling, liquid liabilities, inflation, size of the modern sector, physicians per 1,000 people, and the labor regulation indicator. The pooled regressions include a constant. Our set of instruments for the labor indicators consists of the level of development, trade openness adjusted by geographic variables, leftist political orientation of the government, British legal origin, German legal origin, and institutionalized autocracy; this set of instruments was chosen based on the existing literature, following Djankov and others (2003). Standard errors are in parentheses and are robust to autocorrelation and heteroskedasticity (White, 1980). Full regression results and standard errors of the coefficients of the labor regulation variables are not reported for reasons of space, but they are available from the authors on request.

Table B7. The Impact of the L_2 Index of De Facto Regulations: Panel Data Regression Analysis with Least Squares[a] (Dependent variable: Gini coefficient)

	Full sample			*Developing countries*		
Explanatory variable	*Pooled*	*Time fixed effects*	*Country fixed effects*	*Pooled*	*Time fixed effects*	*Country fixed effects*
Output per capita (in logs)	0.185**	0.182**	0.101	0.217**	0.261**	0.154
	(0.08)	(0.09)	(0.12)	(0.10)	(0.12)	(0.15)
Output per capita squared	−0.012**	−0.011	−0.005	−0.013**	−0.015**	−0.008
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Economic growth	−0.281*	−0.238*	0.125	−0.203	−0.162**	0.172
	(0.16)	(0.15)	(0.10)	(0.17)	(0.17)	(0.12)
Secondary schooling	−0.024**	−0.029**	−0.022	−0.032**	−0.043**	−0.034**
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Liquid liabilities	−0.051**	−0.058**	0.025	−0.064**	−0.062**	0.022
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.03)
Physicians per 1,000 people	−2.753**	−3.398**	0.806	−3.963**	−4.574**	1.295
	(0.72)	(0.80)	(0.70)	(0.97)	(1.01)	(1.16)
Inflation	0.013	0.018	−0.018	0.017	0.028*	−0.017
	(0.02)	(0.02)	(0.01)	(0.02)	(0.02)	(0.01)
Size of the modern sector	0.265**	0.230**	−0.175	0.261**	0.232**	−0.168**
	(0.06)	(0.06)	(0.07)	(0.06)	(0.07)	(0.08)
Labor rigidity indicator (L_2)	−0.065**	−0.061**	0.126	−0.053	−0.054*	0.159**
	(0.03)	(0.03)	(0.05)	(0.04)	(0.03)	(0.06)
Summary statistic						
No. observations	344	344	344	272	272	272
R^2	0.38	0.41	0.90	0.26	0.29	0.89
GDP turning point	7.98	8.60	9.34	8.53	8.82	9.49

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002).
a. We report the regression coefficient for the indicator of labor rigidity according to equations (2) and (3) in the text. The dependent variable is the Gini coefficient (0–1). Our control variables are output per capita (in logs), output per capita squared, secondary schooling, liquid liabilities, inflation, size of the modern sector, physicians per 1,000 people, and the labor regulation indicator. The pooled regressions include a constant. Our set of instruments for the labor indicators consists of the level of development, trade openness adjusted by geographic variables, leftist political orientation of the government, British legal origin, German legal origin, and institutionalized autocracy; this set of instruments was chosen based on the existing literature, following Djankov and others (2003). Standard errors are in parentheses and are robust to autocorrelation and heteroskedasticity (White, 1980). Full regression results and standard errors of the coefficients of the labor regulation variables are not reported for reasons of space, but they are available from the authors on request.

Table B8. The Impact of the L_2 Index of De Facto Regulations: Panel Data Regression Analysis with Instrumental Variables[a] (Dependent variable: Gini coefficient)

	Full sample			Developing countries		
Explanatory variable	*Pooled*	*Time fixed effects*	*Country fixed effects*	*Pooled*	*Time fixed effects*	*Country fixed effects*
Output per capita (in logs)	0.215**	0.210**	0.158	0.330**	0.403**	0.269
	(0.09)	(0.09)	(0.13)	(0.12)	(0.15)	(0.18)
Output per capita squared	−0.014**	−0.012**	−0.010	−0.020**	−0.024**	−0.016
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Economic growth	−0.302*	−0.262*	0.073	−0.201	−0.158	0.104
	(0.17)	(0.16)	(0.11)	(0.17)	(0.18)	(0.14)
Secondary schooling	−0.021**	−0.027**	−0.017	−0.032**	−0.043**	−0.033**
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Liquid liabilities	−0.048**	−0.055**	0.027	−0.065**	−0.069**	0.022
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.03)
Physicians per 1,000 people	−3.121**	−3.664**	0.994	−4.463**	−4.760**	1.521
	(0.85)	(0.86)	(0.73)	(1.06)	(1.08)	(1.26)
Inflation	0.016	0.020	−0.016	0.021	0.031*	−0.017
	(0.02)	(0.02)	(0.01)	(0.02)	(0.02)	(0.01)
Size of the modern sector	0.257**	0.224**	−0.130*	0.265**	0.234**	−0.122
	(0.07)	(0.07)	(0.08)	(0.07)	(0.07)	(0.10)
Labor rigidity indicator (L_2)	0.091	0.059	−0.364	0.181	0.109	−0.144
	(0.10)	(0.10)	(0.31)	(0.13)	(0.13)	(0.39)
Summary statistic						
No. observations	330	330	330	258	258	258
R^2	0.40	0.41	0.90	0.29	0.31	0.88
GDP turning point	7.90	8.44	8.30	8.15	8.37	8.54

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
Source: Authors' calculations, based on data from Rama and Artecona (2002).
a. We report the regression coefficient for the indicator of labor rigidity according to equations (2) and (3) in the text. The dependent variable is the Gini coefficient (0–1). Our control variables are output per capita (in logs), output per capita squared, secondary schooling, liquid liabilities, inflation, size of the modern sector, physicians per 1,000 people, and the labor regulation indicator. The pooled regressions include a constant. Our set of instruments for the labor indicators consists of the level of development, trade openness adjusted by geographic variables, leftist political orientation of the government, British legal origin, German legal origin, and institutionalized autocracy; this set of instruments was chosen based on the existing literature, following Djankov and others (2003). Standard errors are in parentheses and are robust to autocorrelation and heteroskedasticity (White, 1980). Full regression results and standard errors of the coefficients of the labor regulation variables are not reported for reasons of space, but they are available from the authors on request.

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DID EUROPEAN LABOR MARKETS BECOME MORE COMPETITIVE IN THE 1990S? EVIDENCE FROM ESTIMATED WORKER RENTS

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Unemployment has been very high in a number of European countries for almost three decades. Many economists ascribe the problem to lack of competition in labor markets plagued by institutional rigidities, such as employment protection, generous unemployment benefits, compression in relative wages as a result of collective bargaining, and so on. Few countries have removed these rigidities, however. Instead, governments have developed a lot of (often very costly) policies with dubious effects, including permanent budget deficits, relief jobs in the public sector that do little to enhance the job prospects of the long-term unemployed, and voodoo economics such as working-time reduction. Some marginal reforms may have had an effect, as in the case of the liberalization of temporary contracts in Spain and other countries in the 1980s and 1990s or France's recent reform of its unemployment benefit system to monitor job search efforts. A detailed look at the history of labor market reforms in several European countries reveals the following characteristics. First, reforms are numerous and amount to an accumulation of small changes. Second, some reforms tend to increase labor market flexibility, while others tend to reduce it. Third, it is quite difficult to assess the magnitude of the impact of individual reforms and, in particular, whether they have made European labor markets more competitive.

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The degree of labor market competition may also be affected by developments such as increases in product market competition as a result of deregulation or increased openness to international trade. Such changes may help reduce European unemployment and thus spare painful reforms of the labor market—although groups that benefit from labor rigidities have an interest in blocking these changes.[1] Consequently, labor market competition might increase even in the absence of labor market reforms.

Rather than looking directly at policy measures, this paper looks at the evolution of some quantitative measures of labor market competition. Specifically, I look at the evolution of two very different measures of labor market competition in a number of European countries between 1994 and 2000. The first measure captures interindustry differences in wages, while the second is a proxy for the welfare difference, in present discounted value terms, between the employed and the unemployed.

1. Rents and Their Meaning

I define the rent of an employed worker as the present discounted value of his or her expected flow of future incomes, minus the present discounted value of the income flow of an unemployed worker with similar characteristics. This measure indicates how uncompetitive the labor market is. In a perfectly competitive labor market, the unemployed would be able to underbid the employed to the point that people would be indifferent between being employed or unemployed. That may mean full employment, in which case an unemployed person would immediately find a job, such that his or her situation would in effect be no different from that of an employed person. Alternatively, it may mean that the wage has fallen to the level of unemployment benefits (adjusted for the disutility of effort), in which case there is voluntary unemployment in the sense that the unemployed are indifferent about getting a job.

The rent also identifies how much workers lose when they lose their job. In a no-rent society, the risk of job loss is not a risk at all. People are insured against it by the perfectly competitive labor market, which makes them indifferent between working and not working.

1. See Blanchard and Philippon (2003) for an analysis.

All the implications of job loss being painful derive from the fact that employed workers have rents.

Rents may originate from microeconomic frictions, which prevent the labor market from being competitive. The theoretical literature identifies a number of channels. The efficiency wage theory, for example, states that it is costly for firms to monitor their workers' effort level.[2] They therefore prefer to pay above market-clearing wages so as to deter shirking. This theory implies that the rent will rise in line with the severity of the informational problems associated with observing effort. The insider-outsider theory, in turn, states that firms have sunk specific investments in locating and training workers, which generates a hold-up problem.[3] Once the investment is paid for, the worker can expropriate part of it by asking above market-clearing wages. This theory predicts that rents will be higher for jobs requiring larger specific investments ex ante.[4] It also predicts that larger rents will be accompanied by greater workers' bargaining power —that is, the share of the total surplus that they are able to appropriate, although there is no straightforward empirical equivalent of that parameter. The search-and-matching theory extends the insider-outsider theory to a general equilibrium framework in which there is a per-unit-of-time cost of maintaining a vacancy and in which the rate at which vacancies are filled depends on the ratio between the stock of unemployment and the stock of vacancies.[5] The tighter the labor market, the longer it takes to fill a vacancy—which raises both the sunk costs of hiring and rents. The theory thus predicts a positive relation between rents and labor market tightness. It also predicts that higher rents raise the cost of vacancies and lower the efficiency of the process of matching workers and firms. Finally, union-wage-setting models directly generate rents, as unions act as monopolies in the labor market.

All these models further predict that a number of labor market regulations affect the rent. Firing costs increase the rent under all of the models. In the efficiency wage model, it makes it more costly to dismiss workers when they have been caught shirking, which raises the rent that must be paid to deter the behavior. In the insider-outsider model,

2. See Shapiro and Stiglitz (1984); Solow (1979); Schlicht (1978).

3. See Lindbeck and Snower (1988); Blanchard and Summers (1986); Layard, Nickell, and Jackman (1990).

4. The macroeconomic consequences of the degree of specificity in investments are explored by Caballero and Hammour (1998).

5. See, for example, Mortensen and Pissarides (1994).

it acts as a sunk cost, as it must be paid to get rid of any workers in order to replace them with others. Minimum wages directly increase the rent for those employed workers for whom they are binding. Work rules may also increase rents to the extent that they impose specific investments on firms and more generally reduce competition among workers.

Product market regulation also affects rents. By increasing monopoly power, regulation increases a firm's total revenue per worker; the rent is increased as long as the workers have some ability to seize part of that revenue. Under union-wage-setting models, workers' rents are linked to product market competition via a simple law of derived demand. A more regulated product market implies a lower price-elasticity of demand for each firm, which in turn implies a lower wage-elasticity of labor demand and thus a higher wage.

I now briefly discuss the political consequences of rents.[6] The observation made above, that rents show how much workers lose when they lose their job, implies that in an economy with rents, workers have a general aversion to job loss—the more so, the greater the rent. Incumbent employees tend to oppose policies that threaten their jobs and to promote policies that protect them. That incentive would be absent in an economy without rents. If rents differ among workers, then workers will support different policies, with workers with greater rents in favor of increasing protection.

This implies that greater rents heighten support for employment protection legislation. Since employment protection itself also tends to increase rents, the system involves a mutual feedback mechanism. Beyond that, any shift that tends to increase rents should enhance support for employment protection. Thus, political support for employment protection should be higher after a hike in the minimum wage, after a period of tight labor markets, or after any technological or organizational change that reduces a firm's ability to monitor workers or raises its required specific investment in a job.

Rents also easily generate political-economic complementarities between different labor market institutions. By a political-economic complementarity between institution A and institution B, I mean that the political support for institution A is greater if institution B is in place, and vice-versa. As I just argued, institutions that create (or increase) rents increase the political support for employment protection. Employment protection itself, however, increases the political

6. See Saint-Paul (1997, 2000, 2002) for an analytical treatment.

support from employed workers for institutions that create rents, because it reduces their exposure to unemployment and thus their prospects of losing the rent. Political-economic complementarities imply that a comprehensive labor market reform will have more support than a piecemeal approach.

While rents increase the support for institutions that directly increase employment protection, they also have a pervasive effect on the way people view most policy changes. When the rent is high, incumbent employees have a vested interest in opposing policies that threaten their jobs. This means that any policy change that implies some labor reallocation will face greater political opposition in economies with high rents.[7] This applies to trade liberalization, changes in the level and structure of government spending, and so on. In other words, rents tend to generate a bias in favor of the status quo in virtually any policy area.

The story of labor market flexibility in Europe in the 1990s is very much that of a half-full, half-empty glass: measures that increased labor flexibility alternated with measures that reduced it. As a result, rents will not necessarily have fallen, but their evolution in a given country may illustrate which reforms had the strongest effects. At the same time, increased trade integration and deregulation in product markets is a clear trend that should push rents downward.

2. Measuring Competition in European Labor Markets

There are various ways to assess whether European labor markets are becoming more competitive. One possibility is to construct indices of labor market regulation and look at their evolution over time in different countries. This approach has mostly been pioneered by the Organization for Economic Cooperation and Development (OECD).[8] The reliability of these indices depends on how quantitative the underlying variables are, together with the accuracy of the researcher's assessment of the importance of a given change in regulation. In some cases, it is easy to construct an index because the regulation being measured has a clear quantitative definition. This is the case, for example, for unemployment benefits, for which fairly

7. See Saint-Paul (1996b).

8. Typical examples include Grubb and Wells (1993) and the OECD's job study (1994).

reliable indices of replacement ratios have been constructed. Even in such a case, however, the index is not fully accurate because it fails to capture the diversity of individual situations and the way the unemployment benefit system is actually administered. Constructing indices of more qualitative regulations such as employment protection is obviously even more complicated. These indices do well in cross-sectional comparisons, but they are more problematic for assessing evolutions over time.[9] For example, in the 1990s many countries moved back and forth in the liberalization of temporary contracts, and this was sometimes accompanied by moves in the opposite direction concerning the degree of protection for permanent contracts. It is not easy to determine whether employment protection goes up or down if a reform makes it harder to use temporary contracts but at the same time eases the conditions under which a permanent worker may be dismissed.

It is thus useful to pursue a different approach, namely, to look at direct quantitative indicators of worker's rents. The drawback is that this approach does not specify which reforms have been implemented; workers' rents may fall under a number of labor market reforms, product market reforms, or the sheer pressure of international competition. It does, however, provide an idea of the evolution of the true degree of competition in labor markets, and it can help one avoid misclassifying a policy change or taking one seriously when it actually has only second-order effects on labor market flexibility or when is not enforced. To measure rents, I use two different approaches: the interindustry approach and the transition approach, which are described in the next two sections.

3. The Interindustry Approach

The first approach exploits variation of wages across industries. This empirical regularity was much studied in the 1980s and 1990s, following Krueger and Summers (1988). The literature shows that these differentials are not associated with compensating differentials for working conditions or nonwage benefits or with unobservable worker heterogeneity. On the other hand, they are correlated with a number of industry characteristics—such as union density, capital

9. Indeed, indices such as the one in Bertola (1990) are typically used for cross-sectional studies.

intensity, and product market competition—that are likely to be associated with the rent that can be extracted by workers and their power to do so. In other words, there is a strong presumption that differences in wages between industries represent differences in rents rather than anything else. Therefore, I analyze the evolution of labor market rents over time in a number of European countries by looking at trends in the estimated coefficients of a wage equation, in an individual data set, with industry dummies. If rents are falling over time, then I expect the dispersion in these coefficients across sectors to be falling, too: in a rent-free economy, they would all be equal to zero. Assuming that the least-paying sector is more or less perfectly competitive, I define an average rent by looking at the employment-weighted average of the difference between a sector's coefficient and that of the least-paying sector. That alternative measure captures changes in the rent that are due to labor reallocation from high-rent to low-rent sectors, whereas the dispersion measure gives an indication of the evolution of the rent in a given sector.

The data is the European Household Panel Survey. The advantage of this data set is that it includes data on wages, individual characteristics, and labor market status, which are consistent across countries and available for all European Union (EU) members. Its panel dimension allows me to control for unobserved heterogeneity among individuals by making use of fixed-effect estimators. The drawbacks are that it has fewer observations than a typical national labor-force survey and that data for Germany and the United Kingdom are not available after 1996.

I estimate wage equations for each of the countries, in which each observation is an individual at a given date. The specification is

$$\ln w_{it} = b_0 ED3_{it} + b_1 ED2_{it} + b_2 AGE_{it} + b_3 AGE_{it}^2 + b_4 MARRIED + b_5$$

$$+ \sum_{s=2}^{T} \sum_{k=1}^{N} c_{ks} \left(ID_{it}^{k} TD_{it}^{s} \right) + \sum_{k=2}^{N} c_{k1} \left(ID_{it}^{k} TD_{it}^{1} \right) + c_0 \qquad (1)$$

where TD^s is a time dummy for date s ($TD_{it}{}^{s} = 1$ if $t = s$ and 0 otherwise); ID^k is an industry dummy for industry k ($ID_{it}{}^{k} = 1$ if individual i works in industry k at date t and 0 otherwise); T is the number of periods; and the other variables are self-explanatory.

The above equation can be estimated with and without individual fixed effects. The fixed effects eliminate potential sources of bias like unobserved heterogeneity among workers. If workers with greater unobserved ability are more likely to work in certain industries, then part of the industry dummy reflects the return to unobserved ability rather than a rent. The earlier literature finds that interindustry wage differentials are typically robust to the introduction of individual fixed effects, although the coefficients are somewhat smaller than when the specifications are run without fixed effects.[10]

Next, I construct synthetic indicators of labor market rents by first defining the spread indicator for any date, s, as

$$SPREAD_S = \max_k c_{ks} - \min_k c_{ks} \ .$$

This equation captures the difference in wages for similar workers in the best-paying and the worst-paying sectors. If the worst-paying sector is interpreted as perfectly competitive, then the spread indicator is a measure of the highest rent paid to workers in that economy, irrespective of the number of workers who earn the rent.[11] It would fail to capture a reduction in rents stemming from a fall in the employment share of the best-paying sectors. I therefore also compute an average rent indicator (*ARENT*) for date s as follows:

$$ARENT_s = \frac{\sum_{k=1}^{N} n_{ks}\left(c_{ks} - \min_j c_{js}\right)}{\sum_{k=1}^{N} n_{ks}} \ ,$$

where n_{ks} is the number of people employed in industry k at date s and where $c_{jkN} = 0$ by extension.[12]

ARENT measures the average rent earned by a worker in that economy, as compared with the least-paying sector. If that sector is competitive, *ARENT* also provides an idea of the welfare difference, in annuity terms, between an employed person and an unemployed person.

10. See Saint-Paul (1996a) for a survey.
11. For date $s = 1$, the formula is slightly different: $SPREAD_1 = \max(\max_k c_{k1}, 0) - \min(\min_k c_{ks}, 0)$.
12. For $s = 1$ the formula is again slightly different:

$$ARENT_1 = \frac{\sum_{k=1}^{N} n_{k1}\left(c_{k1} - \min(\min_j c_{j1}, 0)\right)}{\sum_{k=1}^{N} n_{k1}} \ .$$

Once these indicators are constructed, I look at their evolution over time in each country. One shortcoming with the data used is that they are only available for seven consecutive years (three for Germany and the United Kingdom), which may cause problems if there are long lags between reforms and their effects on labor market competition. I also perform another exercise, namely, looking at wage differentials across size categories of firms rather than industries, using the same methodology.

4. The Transition Approach

The second approach, in the spirit of Cohen (1999), estimates a dynamic process for individual transitions between employment and unemployment and uses the estimated coefficients to compute the present discounted value of being employed and the present discounted value of being unemployed for any given category of worker. The difference between the two represents the total rent of the employed.

Assume that for a given category of workers, individuals move between two states, employed and unemployed. The transition rate from employment to unemployment is s; the transition rate from unemployment to employment is h. The income in unemployment is b, and the income in employment is w. The real interest rate is r. Workers are risk neutral.

Then, the evolution equation for the value of being employed, V_e, defined as the expected present discounted value of income flows when employed, is the following:

$$rV_e = w + s(V_u - V_e) + \dot{V}_e \ .$$

Similarly, the evolution equation for the value of being unemployed, V_u, is

$$rV_u = b + h(V_e - V_u) + \dot{V}_u \ .$$

In the steady state, the total rent—defined as the difference between the utility of the employed and that of the unemployed (that is, by $Q = (V_e - V_u)$—is

$$Q = \frac{w - b}{r + s + h} \cdot$$

Another concept of interest is the cost per unit of time to the employer of having to pay the rent, Q, in addition to the worker's alternative wage. It is given by the annuity equivalent of the rent, Q, that is, $q = (r + s)Q$:

$$q = \frac{(r+s)w-b}{r+s+h} \cdot$$

While the total rent, Q, is measured in terms of workers' welfare, the annuity rent q expresses the same concept from the point of view of the firm's labor cost. The rent, q, tells us how much firms have to pay workers per unit of time in addition to their alternative wage, rV_u: $q = w - rV_u$. The two differ because welfare can be transferred to workers not only in the form of wages, but in the form of job security. The rent, q, goes up with s, because a higher job loss rate reduces the welfare of unemployed workers, since their prospective jobs do not last as long. It goes down with h for the opposite reason. In contrast, Q falls with s, because the employed workers are worse off when their jobs are insecure, all else equal. Nevertheless, the gap between their wage and their alternative wage widens.

In principle, if I can estimate transition rates between employment and unemployment, as well as the income of the employed and the unemployed, then I can compute Q and q.

The most important shortcoming with that approach is that if w, b, s, and h have different cyclical elasticities, variations in q and Q over a period of a few years are as likely to result from the influence of business cycles as from underlying changes in the degree of labor market competition. To control for that possibility, I pool all the countries together and impose a common response of these variables to country-specific business cycle conditions. This leads to the following specification:

$$\begin{aligned} Y_{it} = {} & \sum_{j=1}^{P} CD_{it}^{j}\left(a_{j0} + a_{j1} SB_{it}\right) + \\ & \sum_{j=1}^{P}\left(b_0^j ED3_{it} + b_1^j ED2_{it} + b_2^j AGE_{it} + b_3^j AGE_{it}^2 + b_4^j MARRIED + b_5^j SEX_{it}\right. \\ & +\left(c_0 U_{it} + c_1 U_{it-1} + c_2 \ln GDP_{it}\right) \end{aligned} \tag{2}$$

where Y_{it} is one of the four variables of interest, w, s, b, and h (defined below); and P is the number of countries. There are three blocks. The first block captures the country-specific evolution of Y over time, where CD_{it}^{j} is a country dummy. The second block captures the effect of individual characteristics, assuming country-specific responses. The third block captures the effect of the business cycle: U_{it} is the unemployment rate in the country where the individual observation is located, and GDP_{it} is the country's real GDP. The coefficients are assumed common across countries, which allows identification. The structural break dummies, SB_{it}, are defined by

$$\begin{aligned} SB_{it} &= 0 \quad \text{if} \;\; t \le t_0 \\ SB_{it} &= 1 \quad \text{if} \;\; t > t_0 \;. \end{aligned} \tag{3}$$

These allow me to compute the country-specific change in w, s, b, and h between the two subperiods defined by equation (3).

A second shortcoming with the approach is that it is difficult to get reliable estimates of b, the unemployment benefit payments, from the data. The problem is that the database is silent about the flow of unemployment benefits payments. Rather, unemployment benefits payments are reported for the whole year, and there appears to be a lag between unemployment spells and the actual payment of corresponding benefits. My attempts to solve this issue using econometric methods failed in that they yielded estimates for Δb that are not plausible for many countries and that do not match the evolution of unemployment benefits replacement ratios over time as estimated by the OECD.

I therefore use equation (2) only for estimating Δw, Δs, and Δh. The three variables of interest are defined as follows: $\ln w_{it}$ is the log of individual earnings for an employed person, in which case the regression is estimated using only observations such that the individual is employed at t (regression 1); ED_{it} is a dummy equal to 1 if the individual is employed at t, in which case the regression uses only observations such that the individual was unemployed at $t - 1$ (regression 2); and UD_{it} is a dummy equal to 1 if the individual is unemployed at t, in which case the regression uses only observations such that the individual was employed at $t - 1$ (regression 3). The coefficient a_{j1} represents the change in the relevant variable between the two subperiods. As for Δb, I use estimates of the benefit replacement ratio, $\rho = b/w$, in the first subperiod as reported by Nickell (2003).[13]

13. One problem with that study is that its estimate of the replacement ratio for Italy in the second subperiod is unreliable. A discussion by the author with Pietro Ichino suggests progressive moving toward a replacement ratio of 0.4 in the second subperiod, starting in 1997, and a value of 0.26 in the first one, while estimating a version of equation (1) yields an increase in $\Delta \ln \rho$ of just 0.02 between the two subperiods. As a reasonable compromise, I take $\rho = (0.26 + 0.40)/2 = 0.33$ in the second subperiod.

For any country, this allows me to compute the average change in the total rent:

$$\frac{\Delta Q}{Q} \approx \frac{w}{w-b}\Delta \ln w - \frac{b}{w-b}\Delta \ln b - \frac{\Delta h}{r+s+h} - \frac{\Delta s}{r+s+h} \ ,$$

or, equivalently,

$$\frac{\Delta Q}{Q} \approx \Delta \ln w - \frac{\rho \Delta \ln \rho}{1-\rho} - \frac{\Delta h}{r+s+h} - \frac{\Delta s}{r+s+h} \ .$$

This number is computed using the average unconditional values of w, b, h, and s in the first subsample (t = 1,... S) and r = 0.03. Similarly, I can compute the change in the rent in annuity terms:

$$\frac{\Delta q}{q} \approx \Delta \ln w - \frac{\rho \Delta \ln \rho}{1-\rho} - \frac{\Delta h}{r+s+h} + \frac{h \Delta s}{(r+s+h)(r+s)} \ .$$

5. Results I: The Interindustry Approach

The estimated industry coefficients are highly significant and typically range up to 50–60 percent. In some cases the number of observations is too low in a given time × country × industry cell, and the coefficient cannot be used. I have therefore dropped Luxembourg, Greece, and years 1999 and 2000 for Belgium. Also, the Panel stops in 1996 for Germany and the United Kingdom, and it starts in 1995 for Austria and in 1996 for Finland. Finally, the Netherlands includes what is probably an aberrant observation in 1998, owing to a sharp drop in the estimated industry dummy coefficient for textiles.

The tables in this section report the results for the two rent indicators, *SPREAD* and *ARENT*. In table 1, the *SPREAD* measure fluctuates in all countries, but it does not seem to follow any clear trend. In other words, the rent of the best-paid workers relative to their characteristics does not seem to vanish. The exceptions are Austria, where rents seem to go down, and Finland and the Netherlands, where they go up. Overall, the results confirm the findings of Krueger and Summers (1988) that interindustry wage differentials are quite persistent over time.

Table 2 reports the results for the *ARENT* measure. Most countries display no clear upward or downward trend for the estimated average rent. In the cases of Spain and Italy, the measure is remarkably stable. Again, the rent seems to have gone down in Austria, and to have gone up in Finland.[14]

Table 1. Evolution of *SPREAD*

	Year						
Country	*1994*	*1995*	*1996*	*1997*	*1998*	*1999*	*2000*
Austria	–	0.59	0.55	0.40	0.46	0.42	0.37
Belgium	0.24	0.20	0.26	0.23	0.23	–	–
Denmark	0.27	0.31	0.26	0.17	0.33	0.31	0.26
Finland	–	–	0.24	0.23	0.23	0.25	0.35
France	0.45	0.40	0.41	0.37	0.42	0.38	0.45
Germany	0.53	0.43	0.43	–	–	–	–
Ireland	0.67	0.61	0.57	0.59	0.79	0.56	0.70
Italy	0.47	0.35	0.44	0.36	0.40	0.42	0.41
Netherlands	0.33	0.32	0.33	0.30	0.63	0.46	0.42
Portugal	0.49	0.54	0.51	0.58	0.50	0.53	0.53
Spain	0.55	0.54	0.54	0.56	0.55	0.56	0.60
United Kingdom	0.66	0.57	0.62	–	–	–	–

– Data do not cover this year.

Table 2. Evolution of *ARENT*

	Year						
Country	*1994*	*1995*	*1996*	*1997*	*1998*	*1999*	*2000*
Austria	–	0.47	0.36	0.30	0.35	0.34	0.26
Belgium	0.12	0.08	0.17	0.13	0.13	–	–
Denmark	0.17	0.14	0.18	0.09	0.20	0.21	0.15
Finland	–	–	0.16	0.16	0.14	0.18	0.27
France	0.23	0.18	0.20	0.16	0.22	0.18	0.21
Germany	0.32	0.29	0.26	–	–	–	–
Ireland	0.47	0.40	0.37	0.38	0.45	0.41	0.50
Italy	0.20	0.18	0.20	0.19	0.18	0.20	0.17
Netherlands	0.15	0.21	0.19	0.13	0.45	0.22	0.17
Portugal	0.14	0.15	0.15	0.18	0.14	0.15	0.16
Spain	0.24	0.20	0.19	0.20	0.22	0.21	0.23
United Kingdom	0.40	0.31	0.32	–	–	–	–

– Data do not cover this year.

14. One shortcoming is that the results are substantially driven by the difference between the agricultural sector and all other sectors, as the former pays considerably lower wages than the rest. This would not be a problem if, for example, the agricultural sector pays no rent at all—such that people are indifferent between working in that sector and being unemployed—while all other sectors pay rents that are similar. However, no clear pattern emerged when I dropped the agricultural sector from my computations of the rent indicators.

These results may be driven by sectors with too few observations, which would imply a potentially volatile associated coefficient. To check for that, I constructed alternative estimates of *ARENT* and *SPREAD*, which use only sectors with more than a hundred observations in the first wave. This means that the variables are defined using a different set of industries in different countries, but that is unimportant here because I am not comparing the average level of the rent across countries. The results for *ARENT* (reported in table 3) are slightly different from those of table 2. Rents now seem to go down in Ireland and perhaps France and Italy, and to go up again in Finland, with no clear pattern elsewhere. In particular, they no longer seem to be falling in Austria.

The usual problem of unobserved heterogeneity among workers also applies. I therefore also computed the fixed effect estimator. One problem, though, is that if people do not move much between industries, then such a panel, with relatively few periods and observations, is likely to present fixed effects that are highly collinear with the vectors of industry dummies. The following results should thus be taken with caution. As shown in table 4, the estimated spread is quite volatile, although there is still evidence of a downward trend in rents in Austria. Also, in many countries, rents computed using the fixed effect estimators are smaller than under random effects, as expected.

When fixed effects are applied to the average rent, a few strange phenomena arise, like the quasi-disappearance of the average rent in

Table 3. Evolution of *ARENT*, Robust Definition

				Year			
Country	*1994*	*1995*	*1996*	*1997*	*1998*	*1999*	*2000*
Austria	–	0.06	0.04	0.04	0.04	0.07	0.07
Belgium	0.06	0.08	0.08	0.09	0.07	0.12	0.05
Denmark	0.06	0.06	0.07	0.05	0.06	0.06	0.06
Finland	–	–	0.04	0.03	0.04	0.04	0.09
France	0.24	0.16	0.16	0.15	0.19	0.19	0.18
Germany	0.13	0.11	0.11	–	–	–	–
Ireland	0.20	0.23	0.14	0.15	0.13	0.11	0.11
Italy	0.20	0.17	0.20	0.18	0.18	0.19	0.17
Netherlands	0.08	0.06	0.06	0.06	0.12	0.05	0.08
Portugal	0.12	0.13	0.12	0.15	0.12	0.13	0.14
Spain	0.23	0.19	0.18	0.19	0.21	0.20	0.22
United Kingdom	0.22	0.25	0.20	–	–	–	–

– Data do not cover this year.

France, Spain, and Italy (see table 5). Again, the measure seems highly volatile, but there is still a downward trend in Austria.

To conclude, no country demonstrates a clear trend. There is mild evidence of falling rents in Austria and Ireland, but it is not robust across estimators. If I had to choose a preferred estimation, however, I would opt for that of table 3, which is based on the least volatile estimates of the interindustry dummies. That table suggests a sharp drop of rents in Ireland but not elsewhere.

Table 4. Evolution of *SPREAD*, Fixed Effects

	Year						
Country	*1994*	*1995*	*1996*	*1997*	*1998*	*1999*	*2000*
Austria	–	0.39	0.28	0.20	0.27	0.26	0.11
Belgium	0.15	0.11	0.24	0.13	0.19	–	–
Denmark	0.30	0.41	0.28	0.32	0.41	0.31	0.26
Finland	–	–	0.28	0.22	0.26	0.20	0.17
France	0.29	0.25	0.32	0.24	0.32	0.27	0.28
Germany	0.31	0.35	0.38	–	–	–	–
Ireland	0.49	0.40	0.35	0.36	0.66	0.56	0.60
Italy	0.30	0.19	0.28	0.18	0.26	0.27	0.26
Netherlands	0.37	0.23	0.43	0.31	0.73	0.54	0.55
Portugal	0.13	0.10	0.17	0.13	0.12	0.16	0.18
Spain	0.40	0.16	0.12	0.24	0.18	0.30	0.32
United Kingdom	0.65	0.41	0.38	–	–	–	–

– Data do not cover this year.

Table 5. Evolution of *ARENT*, Fixed Effects

	Year						
Country	*1994*	*1995*	*1996*	*1997*	*1998*	*1999*	*2000*
Austria	–	0.31	0.21	0.15	0.23	0.19	0.08
Belgium	0.10	0.05	0.13	0.08	0.10	–	–
Denmark	0.16	0.17	0.15	0.22	0.21	0.18	0.21
Finland	–	–	0.14	0.08	0.13	0.06	0.09
France	0.06	0.05	0.09	0.07	0.09	0.07	0.08
Germany	0.18	0.21	0.23	–	–	–	–
Ireland	0.28	0.22	0.19	0.22	0.30	0.47	0.37
Italy	0.08	0.07	0.09	0.07	0.06	0.07	0.06
Netherlands	0.12	0.17	0.31	0.06	0.41	0.18	0.20
Portugal	0.07	0.06	0.10	0.06	0.07	0.10	0.11
Spain	0.06	0.05	0.04	0.07	0.08	0.08	0.10
United Kingdom	0.32	0.19	0.18	–	–	–	–

– Data do not cover this year.

5.1 Size Effects

While interindustry wage differentials are the most widely documented and discussed phenomenon, one may also want to look at wage differentials in other dimensions. Hence, I also present results from partitioning by firm size instead of by industry (see tables 6 and 7). I replaced the industry dummies with two size dummies for the regression, corresponding to three size categories: fewer than 100 employees, 100–500 employees, and more than 500 employees. The results look somewhat more plausible and of better quality than those obtained based on interindustry differences, but unfortunately they do not confirm the early results. Rather, they suggest that rents are declining in Belgium, France, Ireland, Italy (mildly), Portugal, and the United Kingdom, while they seem to be rising in the Netherlands and Spain—two countries where unemployment actually fell over the period! Rents are stable in other countries. Thus, the only country for which these estimates support those of the previous section is Ireland.

To conclude, the interindustry approach does not suggest a systematic pattern of falling rents in Europe. When it does for a given country, the decline does not seem to be related to any fall in unemployment in the corresponding country. Finally, breaking down industries by sector of activity or firm size generates different results. The only inference, if any, that one can confidently make from the exercise is that rents have fallen in Ireland.

One potential problem with the approach is that labor market liberalization may have conflicting effects on the estimated rents. On the one hand, it eliminates pure rents that are not the return to productive ability. On the other, it removes wage compression induced by regulation and collective bargaining, which may widen wage differentials by increasing the return to unobserved ability, match-specific human capital, and so on. If these latter factors are more present in some industries than others, then measured interindustry differences may well widen. A fixed effects estimator does not solve that problem: a given individual will earn different returns in different years if these years are associated with a different regulatory environment.

My provisional conclusion, however, is that there is no firm ground for believing that European labor markets have generally become more competitive in the 1990s on the basis of these estimates.

Table 6. Evolution of *SPREAD*, Size Differentials

	Year						
Country	*1994*	*1995*	*1996*	*1997*	*1998*	*1999*	*2000*
Austria	–	0.10	0.08	0.06	0.08	0.10	0.09
Belgium	0.11	0.12	0.12	0.12	0.12	0.09	0.06
Denmark	0.07	0.08	0.09	0.08	0.09	0.06	0.08
Finland	–	–	0.08	0.10	0.09	0.09	0.12
France	0.20	0.22	0.22	0.25	0.24	0.18	0.16
Germany	0.24	0.25	0.26	–	–	–	–
Ireland	0.24	0.20	0.21	0.16	0.15	0.10	0.13
Italy	0.09	0.09	0.08	0.08	0.08	0.08	0.06
Netherlands	0.07	0.07	0.08	0.10	0.09	0.12	0.10
Portugal	0.12	0.10	0.14	0.11	0.09	0.08	0.07
Spain	0.11	0.15	0.15	0.14	0.16	0.14	0.17
United Kingdom	0.16	0.13	0.12	–	–	–	–

– Data do not cover this year.

Table 7. Evolution of *ARENT*, Size Differentials

	Year						
Country	*1994*	*1995*	*1996*	*1997*	*1998*	*1999*	*2000*
Austria	–	0.03	0.02	0.02	0.02	0.02	0.02
Belgium	0.05	0.05	0.05	0.04	0.05	0.03	0.01
Denmark	0.03	0.02	0.02	0.02	0.02	0.02	0.02
Finland	–	–	0.02	0.02	0.02	0.02	0.03
France	0.06	0.07	0.07	0.04	0.05	0.04	0.03
Germany	0.08	0.08	0.08	–	–	–	–
Ireland	0.07	0.06	0.06	0.03	0.03	0.02	0.03
Italy	0.02	0.02	0.02	0.01	0.01	0.01	0.01
Netherlands	0.03	0.03	0.03	0.04	0.04	0.05	0.04
Portugal	0.01	0.01	0.02	0.01	0.01	0.01	0.01
Spain	0.03	0.01	0.03	0.03	0.03	0.03	0.03
United Kingdom	0.10	0.07	0.06	–	–	–	–

– Data do not cover this year.

6. Results II: The Transition Approach

This section presents the results of the transition approach to measuring rents. Unfortunately, they are not much more conclusive than the previous exercise, in part because of data problems.

Table 8 reports the evolution of rents under four alternative measures. The two unadjusted rents are defined in section 4. The two adjusted rents deflate the unadjusted ones to allow for growth. (They subtract the expected difference in GDP between the two subperiods on the basis of average GDP growth between 1980 and 2000.) If the rents grow slower than the

Table 8. The Transision Approach

	Unadjusted		*Adjusted*	
Country	$\Delta Q/Q$	$\Delta q/q$	$\Delta Q/Q$	$\Delta q/q$
Austria	0.0139	–0.0200	–0.0580	–0.0900
Belgium	–0.0450	0.1490	–0.1130	0.0800
Denmark	0.0450	–0.0450	–0.0230	–0.1100
Finland	–0.0100	–0.1890	–0.0870	–0.2670
France	0.0480	–0.0580	–0.0110	–0.1100
Ireland	0.0956	0.1940	–0.0840	0.0144
Italy	0.0500	0.0000	–0.0100	–0.0600
Portugal	0.1330	0.0810	0.0258	–0.0256
Spain	0.1120	0.1460	0.0226	0.0563

economy, in the long run they account for a negligible fraction of labor costs and the economy converges toward a competitive labor market. Thus, the first two columns express rents in real consumption units, while the last two columns express them relative to GDP.

A first aspect of the results is that they are not very robust: rents are quite sensitive to whether the specification allows for growth, and the evolution of the total rent, Q, often diverges from that of the rent per unit of time. This suggests that changes in the separation rate, s, play a quantitatively important role in the results (in that an increase in s reduces Q but increases q, all else equal). The evolution of rents again is not closely related to that of unemployment over the period. However, the adjusted $\Delta Q/Q$ does a better job than the other measures; its correlation with the change in unemployment is 0.3. This measure would thus appear to be the preferred one. On that basis, these estimates suggest a more optimistic conclusion than the interindustry approach: rents fall significantly in four countries (Austria, Belgium, Finland, and Ireland), and they fall moderately in three other countries (Denmark, France, and Italy).

7. Conclusion

This paper has provided some quantitative evidence on the evolution of labor market competition in Europe in the 1990s, based on various estimates of labor market rents. The results are rather inconclusive, probably as a result of the quality of the data. A general conclusion is that there is no strong evidence that labor markets became either more or less competitive in any European country over that period. One exception seems to be Ireland, though, for which a number of estimated rents fell significantly over the period.

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Assessing the Flexibility of the Labor Market in Chile: An International Perspective

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The unemployment rate in Chile averaged slightly over 6.5 percent throughout a ten-year period of high economic growth that ended in 1997. Unemployment then rose significantly at the outset of the Asian crisis, reaching levels near 11 percent. This broadly coincided with the implementation of a set of legal initiatives that increased protection standards in labor regulation. After the end of the military government in 1990, labor codes began to revert to their previous trend of high levels of regulation, with reforms approved in congress in 1990, 1993 and 2001. These introduced higher costs of dismissal, sanctions against firms that fire without just cause, extended provisions for and broader coverage of union bargaining, and a significant increase in the minimum wage (which was implemented in three stages beginning in 1998). These regulations are closer to the European-style labor market protection than to the Anglo-American tradition, which relies more on market forces. A quick look at the relative performance of unemployment rates under these two approaches can help clarify the controversy generated among economists and policymakers during this period.

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One side of the argument holds that the recent trend in unemployment is basically cyclical in nature, responding to the GDP slowdown triggered by the Asian crisis and reinforced by the quick succession of negative external shocks that included a large fall in the terms of trade, September 11, and turbulence in Argentina and Brazil. The persistent behavior of unemployment is thus merely a reflection of persistent shocks and not the result of an intrinsically rigid labor market limited by the current legal framework. This framework, the argument continues, is still far from European standards, especially when de jure regulations are placed in the context of their de facto repercussions in a country with weak legal enforcement. The difference between de jure and de facto regulation is considered in Rama and Artecona (2002) and Calderón and Chong (in this volume).

The other side argues that the recent rise and persistence of unemployment is mainly explained by the growing trend in labor regulations since the early 1990s, which have increased rigidity in the labor market. This has been particularly important since the last reform implemented in 2001, which has, according to this view, significantly discouraged employment creation and caused factor substitution toward capital utilization.[1]

Although the first line of argumentation sustains that significant increases in regulation can coexist with low levels of unemployment conditional on the cyclical position (as they did before 1998), some selected facts make a strong case for the second view. First, it is hard to reconcile the cyclical hypothesis with the evidence of some Southeast Asian economies, which managed to recover rather quickly from the crisis despite being severely affected. Second, an argument that draws on factor substitution trends raise the alert on repeating the European experience of growth with unemployment. Indeed, the recent evolution of relative factor costs suggests that the trend in regulation may indeed have strengthened union bargaining power. Despite sharp levels of unemployment, the path of real earnings remained almost unaltered, showing a steady increase throughout the 1990s, while the cost of capital decreased over the period.[2] Something similar happens with minimum wages: the ratio of the minimum wage to average wages was stable in the mid-1990s and then increased sharply

1. Bergoeing and Morandé (2003) argue that the discussion of the reform itself generated an anticipated fall of labor demand.

2. An alternative explanation could be posed based on the efficiency wages theory.

after 1999, whereas the share of young workers in total employment declined steadily across the same period. This combination of events supports the view that the minimum wage is way above equilibrium, which is formally sustained by recent evidence in Cowan and others (2003).

Finally, the concern with the level and persistence of unemployment seems particularly well-founded when two additional aspects are considered. First, two opposite forces usually affect the labor supply in the face of rising unemployment: the discouraging effect on job search among workers with worsened employment chances and the marginal worker effect, in which additional family members must join the labor force to augment domestic earnings. The Chilean experience clearly points to a dominant role for the former, as demonstrated by the decrease in labor-force participation and by empirical estimations of the procyclicality of the labor supply (García, 1995). Second, a growing portion of the recent job creation figures corresponds to self-employment in the informal sector, which increased markedly in the late 1990s in response to the Asian crisis. These jobs are typically low in productivity and socially unprotected. The joint effect of the drop in labor participation and the rise in informal jobs implies that reported unemployment figures actually underestimate the extent of the problem.

However suggestive, the above analysis does not assess the degree of rigidity present in the Chilean labor market, as it fails to identify whether observed unemployment reflects a current shock or persistence stemming from lack of market flexibility. In this context, our goal is to measure the relative flexibility of the labor market by using a performance-based indicator that can account for this distinction, in order to rank Chile within a group of countries that includes both members of the Organization for Economic Cooperation and Development (OECD) and emerging economies.

Our indicator is defined as the half-life of unemployment after the economy is hit by a shock, which is compatible with the cyclical rigidity we examine in this paper. When unemployment quickly converges to its natural rate after a shock, the country's labor market is ranked as being highly flexible, no matter what that natural rate of unemployment might be. We do not address the kind of rigidity that would explain differences in the natural rate of unemployment among countries, although the two issues might be related.

The model that guides our empirical approach is very much in the spirit of Dolado and Jimeno (1997) and Balmaseda, Dolado, and

Lopez-Salido (2000), who associate labor market rigidity with the persistence of unemployment in the presence of macroeconomic shocks.[3] Indeed, their international evidence reveals a clear relation between institutional measures of rigidity and the macroeconomic dynamics. Since our main goal is to rank Chile in terms of labor market rigidity, we naturally consider a model for an emerging open economy that is frequently affected by large movements in the terms of trade, in addition to other supply and demand forces.

The model assumes that wages are set in a bargaining framework in which insiders and outsiders interact, following Blanchard and Summers (1986) and Blanchard (1991). This setting is used to introduce rigidity in the labor market, which prevents nominal wages from adjusting rapidly to equilibrium and leads to partial hysteresis of the unemployment rate. Over the very long run, however, unemployment should be zero (after normalizing for the country-specific natural unemployment rate), which is compatible with a vertical aggregate supply and no trend in the natural unemployment rate (Blanchard and Katz, 1997). All shocks could have an impact on unemployment in the short run, though.

The labor market indicator should depend exclusively on labor market rigidity, since it needs to be comparable across countries. However, some of the rigidity indexes found in the related literature also depend on the elasticity of labor supply to real wages. An open economy version of those indexes would further depend on the share of tradable goods consumed in the economy. Our labor market flexibility index depends exclusively on the coefficient of the model associated with labor market rigidity in the wage equation.

The empirical strategy allows us to compute the direct measure of persistence with which to assess the actual performance of labor markets, by simulating responses to properly identified, isolated shocks. We use a structural vector autoregression (SVAR), with the long-run restriction identification strategy developed by Blanchard and Quah (1989). We use the SVAR to study the dynamics of the real wage, the real exchange rate, output, and unemployment in a sample of both OECD countries and emerging markets. The model helps us to impose the long-run restrictions and interpret the shocks. With the purpose of analyzing unemployment persistence, we focus on the impulse response functions of unemployment after the economy is hit by the structural shocks.

3. Other examples include Fabiani and Rodriguez-Palenzuela (2001) and Viñals and Jimeno (1996).

Our main conclusion is that Chile's labor market reactions to structural shocks are among the most flexible economies, ranking third after Korea and Hong Kong and followed by the United States and Mexico. At the other end of the ranking, Germany, Sweden, Spain, and Colombia have the most rigid labor markets.

The rest of the paper is structured as follows. The first section presents the model. The second describes the empirical strategy and main results. The third section assesses the labor market index, and the final section of the paper concludes.

1. The Model

This section reproduces the basic insights of the model developed in Albagli, García, and Restrepo (2004), starting with aggregate supply and demand. The economy is characterized by the supply of a domestic tradable good by firms, which hire labor as the only factor of production. The technology is assumed to be characterized by constant returns. Aggregate supply is given by

$$y_t = n_t + x_t \ , \qquad (1)$$

where x is the productivity of labor and n is aggregate employment (all variables are in natural logs throughout the description of the model). Consumption is divided into the domestic good and an imported good. To obtain aggregate demand, we use IS-LM analysis for an open economy. The saving-investment equilibrium is given by

$$y_t = -aE\left[r_t\right] + q_t + \eta_z z_t + \eta_x x_t + \tau_t \ , \qquad (2)$$

where $E\left[r_t\right]$ is the expected real interest rate, q is the real exchange rate, z is the relative price of domestic to foreign goods (the terms of trade), x is labor productivity, and τ is a labor-force shock,[4] while a, η_z, η_x are parameters.

4. These different shocks are included separately because they can conceivably affect aggregate demand through different channels. For example, productivity and the terms of trade affect permanent income, while the real exchange rate affects aggregate demand through expenditure-switching effects and balance-sheet effects. Although in the long run the real exchange rate is determined by productivity and the terms of trade, in the short run it behaves according to the nominal price rigidities embedded in the present framework, and should therefore be considered separately. The labor supply shock is included as a scaling factor.

Money market equilibrium is described by

$$m_t - p_t = -bi_t + y_t \ , \tag{3}$$

where m is money suply, p is the price level and b is the semielasticity of real money demand with respect to i, the nominal interest rate. Given perfect capital mobility, nominal interest rates depend on the parity condition, which—together with the Fischer equation—leads to aggregate demand:

$$y_t = \frac{a}{a+b}\left(m_t - p_t\right) + \frac{ab}{a+b} E\left[p_{t+1} - p_t\right] + \frac{b}{a+b}\left(q_t + \eta_z z_t + \eta_x x_t + \tau_t \right. \tag{4}$$

Domestic producer prices depend on nominal wages and productivity, through

$$p_t^p = w_t - x_t \ . \tag{5}$$

The aggregate price level is then given by a weighted average of domestic and foreign prices:

$$p_t = \gamma p_t^p + \left(1-\gamma\right) s_t \ , \tag{6}$$

where s is the nominal exchange rate and γ is the imported fraction of aggregate consumption. The real exchange rate, which is given by $q_t = s_t - p_t$, can be combined with (5) in (6) to obtain consumers' real wages:

$$w_t - p_t = x_t - \frac{\left(1-\gamma\right)}{\gamma} q_t \ . \tag{7}$$

We follow the precedent of papers such as Blanchard and Summers (1986) in establishing nominal wage bargaining as a function of union power. In our particular framework, unions negotiate nominal wages at the beginning of the period (before shocks arrive) to keep real wages equal to the previous period's level, as opposed to the real market-clearing or long-run wage *(w – p)**. This is represented by the following wage setting condition:

$$E\ \left[w_t - p_t\right] = \lambda\left(w_{t-1} - p_{t-1}\right) + \left(1-\lambda\right)\left(w-p\right)^* \ . \tag{8}$$

where E is the expectations operator and λ represents unions' bargaining power.[5]

Labor supply is modeled as a function of real wages and a labor-force shock,

$$l_t = c\left(w_t - p_t\right) + \tau_t \ , \tag{9}$$

where c is the elasticity of labor supply to real wages. Unemployment is then given by

$$u_t = l_t - n_t \ . \tag{10}$$

This basic framework thus defines a long-run equilibrium level of real and nominal variables that depends on four exogenous shocks: namely, shocks to productivity, the terms of trade, the labor supply, and the quantity of money. Each variable is assumed to follow a random walk process:

$$\Delta x_t = \varepsilon_t^x \ , \tag{11}$$

$$\Delta z_t = \varepsilon_t^z \ ,$$

$$\Delta \tau_t = \varepsilon_t^\tau \ , \text{ and}$$

$$\Delta m_t = \varepsilon_t^m$$

where $\varepsilon_t^x, \varepsilon_t^z, \varepsilon_t^\tau$, and ε_t^m are all uncorrelated independent and identically distributed (i.i.d.) shocks.

The economy starts from a position of equilibrium and is then subject to one or more exogenous shocks. The purpose of the model is to highlight how labor market rigidities affect the system's convergence to a new steady state. Once the economy is hit by any of the exogenous shocks, price rigidities stemming from wage bargaining will cause temporary misalignment of the real exchange rate, which directly affects aggregate demand and unemployment.

5. This wage setting framework implies that after a shock, prices are affected by the contemporaneous movement of the nominal exchange rate but not by movements of the nominal wage, which is fixed for the current period.

1.1 Dynamics

In the long run, real variables such as real wages, output, the real exchange rate and employment depend on real determinants only—namely, productivity, terms-of-trade and labor-supply shocks—through the values of x, z, and τ. From equation (7), the long-run workers' (consumers') real wage is

$$(w-p)^* = x - \frac{(1-\gamma)}{\gamma} q^* , \tag{12}$$

where an asterisk represents a long-run value. Equating supply and demand renders

$$q^* = \frac{(1+c-\eta_x)x - \eta_z z}{1+c(1-\gamma)/\gamma} . \tag{13}$$

Using equations (12) and (13) in equation (9) and setting $l = n$ in the long run, we get the steady-state value of output:

$$y^* = \left\{1 + c\left[1 - \frac{(1-\gamma)(1+c-\eta_x)}{\gamma + c(1-\gamma)}\right]\right\} x + \left[\frac{c(1-\gamma)\eta_z}{\gamma + c(1-\gamma)}\right] z + \tau . \tag{14}$$

Nominal variables, therefore, adjust to equations (12), (13), and (14), given the monetary stance, such that the price level in the long run is given by

$$p^* = m - y^* . \tag{15}$$

Nominal wages and the nominal exchange rate are finally obtained by adding equation (15) to the respective real values.

The specification of the shocks allows us to derive long-run identifying restrictions, by reducing the model to a system of four equations. These are equations (5), (13), (14), and (10), which relate real producer wages (labor cost), the real exchange rate, output, and unemployment, respectively, to the four exogenous shocks given by equation (11). In the long run, labor costs depend on productivity shocks only; the real exchange rate is affected by both productivity and terms-of-trade shocks; output reflects productivity, terms-of-trade, and labor-supply shocks; and unemployment responds only temporarily to all shocks, being zero in steady state. The nominal rigidity introduced by wage bargaining, therefore, has no role in the long-run equilibrium.

Turning now to the short run, we analyze the dynamics that are triggered when the system is hit by any of the four shocks, calling for an adjustment in nominal variables in order to reach the new steady state. Although the rigidity introduced by wage bargaining produces symmetric responses of output and employment below and above their long-run levels, we focus only on situations that cause temporal unemployment—that is, shocks that call for a downward nominal wage adjustment. For simplicity, we normalize each variable to zero in the initial state. The timing of the model is the following. First, unions and firms negotiate contracts (nominal wages) at the beginning of the period. Second, the economy is hit by a structural shock, which requires a downward adjustment of the nominal wage to reach the new long-run equilibrium. Since wages are fixed for the present period, prices adjust only partially (through the response of the nominal exchange rate), holding back aggregate demand and causing positive unemployment. Finally, wages are partially adjusted at the beginning of the next period, reflecting the previous real wage level and the new lower steady-state value.

Given the stickiness originating from the wage-bargaining process, prices follow a gradual adjustment path to their full-employment level. The asset channel reacts without this delay, however, so the real exchange rate departs from its long-run level. Thus, output and employment are determined by demand in the short run. If, as the result of a shock, the real exchange rate is below its long-run level, aggregate demand and employment will also be below their long-run levels. In contrast, the labor supply will temporarily rise with respect to its new equilibrium level given the higher real wages (which depend negatively on q); this results in a sharp increase in unemployment.

Panel A in figure 1 presents the comparative statics analysis in the face of a monetary contraction. When m_t falls, aggregate demand moves as a result of two driving forces: the direct impact of the money supply, which lowers demand, and the fall in prices caused by the instantaneous nominal appreciation, which compensates the first effect. The compensation is only partial, however, because nominal wages are fixed at the time of the shock; this ensures a negative dominant effect on labor demand for a given real wage. At the same time, the real wage increases as prices fall, which causes an increase in the labor supply. The result is an expansion of unemployment. This unemployment is gradually reduced as market pressure pushes down real wages during subsequent nominal wage negotiations. Based on equations (4), (8), (9), and (10), we derive that

$$u_{t+s} = \lambda^s u_t \ , \tag{16}$$

which means that the persistence of unemployment is actually given by the unions' bargaining power, λ. The dynamics of the monetary contraction are simulated in panel B of figure 1, using the Anderson-Moore algorithm for solving dynamic systems. Structural parameters were approximated with Chilean data, varying only the hypothetical level of rigidity.

1.2 Building an Index of Labor Market Rigidity

To create a measure that captures the cyclical persistence of the labor market, we need to build an index that satisfies two necessary conditions. First, it must be related to λ. Second, it must be related only to λ. This stems from the fact that two economies with the same degree of labor rigidity may respond differently in output, wages, and unemployment for a given shock. Such differences arise from the other structural parameters introduced in the model, such as c (the response of the labor supply to the real wage) and γ (the relative importance of the tradable sector).

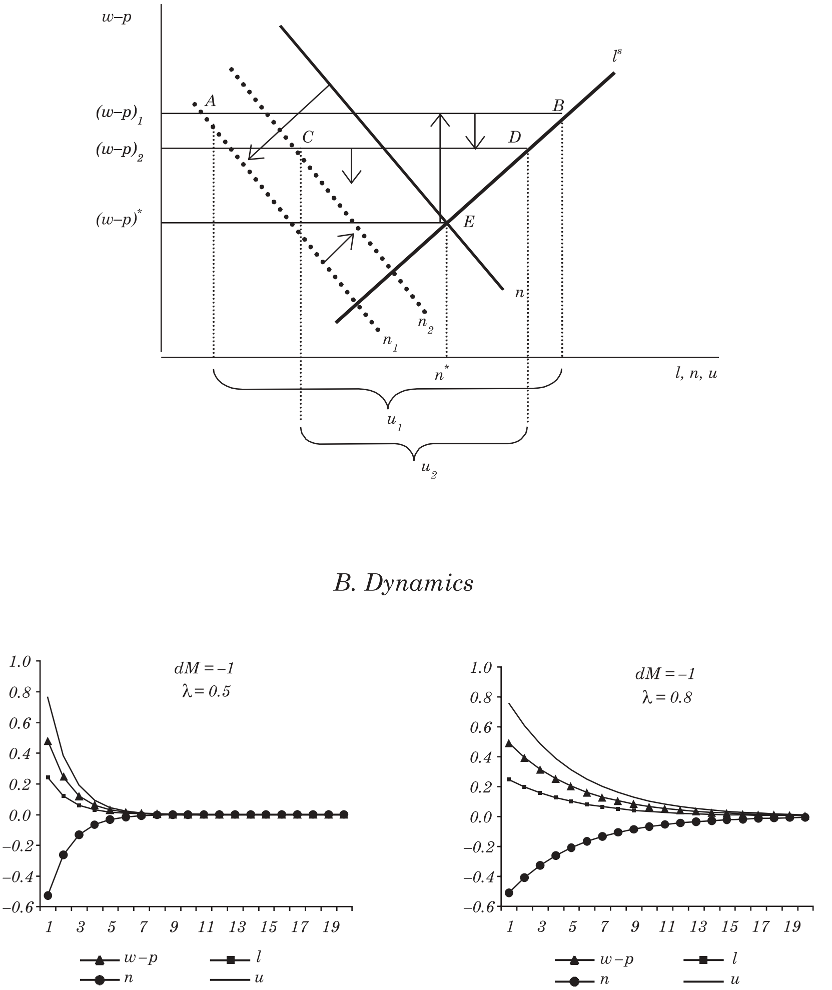
A standard measure of wage rigidity is the index developed in Layard, Nickell, and Jackman (1991) and Balmaseda, Dolado, and López-Salido (2000), which computes the ratio of the accumulated response of unemployment to the change in the real wage after the shock. This type of measure is not appropriate in our current framework, however, because it depends on c and γ as well.

While the assumption of a constant value of c over a rather homogeneous sample of OECD countries seems acceptable, it is not satisfactory when the sample is extended to include less-developed countries, as we do. Assuming similar degrees of openness further deteriorates the power of the measure. We therefore use an alternative measure that depends only on λ: namely, the half-life of unemployment after a shock, or the number of periods required for unemployment to decrease to one-half of its maximum value. From equation (16), unemployment becomes one-half of its initial level at period s^*, where

$$s^* = \frac{\ln(1/2)}{\ln \lambda} \ , \tag{17}$$

Figure 1. A Monetary Contraction

A. Comparative statics



Source: Authors' calculations.

which depends positively and solely on the value of our measure of labor market rigidity.

2. The Empirical Strategy

Since our purpose is to measure unemployment persistence in the presence of shocks, we use the SVAR methodology. In particular, we follow Balmaseda, Dolado, and López-Salido (2000) in identifying the VAR with long-run restrictions, as in Blanchard and Quah (1989) and Clarida and Galí (1994). These authors assume that some shocks have permanent effects on some variables and transitory effects on others and that some shocks may not have any permanent effect on any variable. This procedure fits the intuition of a growing economy in which unemployment goes back to its natural rate, although wages and employment may change because of structural factors, and the supply curve is vertical in the long run.

2.1 Structural Identification

The structural VAR identification, derived directly from the model, provides a clear interpretation of the structural shocks. For clarity, we rewrite several equations taken from the long-run equilibrium of the model:

$$\Delta\left(w - p^{p}\right)^{*} = \varepsilon^{x} \ . \tag{18}$$

Only productivity shocks affect the real producer wage in the long run.

$$\Delta q^{*} = \frac{\left(1 + c - \eta_{x}\right)\varepsilon^{x}}{1 + c\left(1-\gamma\right)/\gamma} - \frac{\eta_{z}\varepsilon^{z}}{1 + c\left(1-\gamma\right)/\gamma} \ . \tag{19}$$

In the long run, the real exchange rate depends only on productivity and the terms of trade.

$$\Delta y^{*} = \left\{1 + c\left[1 - \frac{\left(1-\gamma\right)\left(1 + c - \eta_{x}\right)}{\gamma + c\left(1-\gamma\right)}\right]\right\}\varepsilon^{x} + \left[\frac{c\left(1-\gamma\right)\eta_{z}}{\gamma + c\left(1-\gamma\right)}\right]\varepsilon^{z} + \varepsilon^{\tau} \ . \tag{20}$$

Output is affected in the long run by productivity, the terms of trade, and the evolution of the labor force.

$$u^* = 0 \,. \tag{21}$$

Finally, although all shocks affect unemployment in the short run, the impact is not permanent since unemployment is stationary in a partial hysteresis setting.

The identification is based on the assumption that the matrix of structural long-run multipliers, $\mathbf{C}(1)$, is lower triangular. To find $\mathbf{C}(1)$, it is necessary to first build the matrix $\mathbf{f}(1)\mathbf{\Sigma f}(1)'$ from the reduced form estimation, where $\mathbf{f}(1)$ is the sum of the coefficients, and $\mathbf{\Sigma}$ is the variance-covariance matrix obtained. It is possible to show that $\mathbf{C}(1)$ is the Choleski factor of $\mathbf{f}(1)\mathbf{\Sigma f}(1)'$. Once $\mathbf{C}(1)$ is found, it is easy to compute all the structural coefficients, C, which are used to build the impulse-responses, because $\mathbf{C}_0 = \mathbf{f}(1)^{-1}\mathbf{C}(1)$, and with $\mathbf{C}_0$ all $\mathbf{C}_s$ can be computed given $\mathbf{C}(\mathrm{L}) = \mathbf{f}(\mathrm{L})\mathbf{C}_0$.[6]

2.2 Data

We use quarterly data from 1980:1 to 2002:4 for real producer wages (computed with the GDP deflator), the real exchange rate, output, and unemployment. Most countries' data sets come from the OECD database. For non-OECD countries, data were found in the respective central bank's web site and, in some cases, in the International Monetary Fund's *International Financial Statistics* data set. Table 1 reports the source of the time series for each country.

The model assumes that real wages, real exchange rates, and output are integrated processes, while unemployment is stationary.[7] We ran Dickey-Fuller tests for all variables, but in several countries the null hypothesis of a unit root for the unemployment rate could not be rejected (see table 2). However, we follow Balmaseda, Dolado, and López-Salido (2000) in assuming a partial hysteresis setup, because it seems unreasonable to consider the consequences of any

6. For a detailed explanation, see Clarida and Galí (1994) and Enders (1995).

7. The Dickey-Fuller test rejected the unit root hypothesis for the real exchange rate of Denmark and the Netherlands. We therefore ran the stationary VAR $[\Delta(w - p^p), q, \Delta y, u]$ for these two countries.

Table 1. Quarterly Series Sources[a]

Country	*Data span*	*Unemployment*	*GDP*	*GDP deflator*	*Nominal wage*	*Real exchange rate*
Australia	1984:1–2002:4	OECD	OECD	OECD	OECD[b]	OECD
Austria	1980:1–2002:4	OECD	IMF	IMF	OECD[c]	OECD
Belgium	1980:1–2002:4	OECD	OECD	OECD	OECD[d]	OECD
Canada	1980:1–2002:4	OECD	OECD	OECD	OECD[d]	OECD
Chile	1986:1–2002:4	NSO	Central Bank	Central Bank	NSO[b]	Central Bank
Colombia	1984:1–2002:4	Central Bank	Central Bank	Central Bank	Central Bank[c]	Central Bank
Denmark	1988:1–2002:4	OECD	OECD	OECD	OECD[d]	OECD
France	1980:1–2002:4	OECD	OECD	OECD	OECD[b]	OECD
Germany	1980:1–2002:4	OECD	IMF	IMF	OECD[d]	OECD
Hong Kong	1986:1–2002:4	HKMA	HKMA	HKMA	HKMA[b]	IMF[e]
Italy	1980:1–2002:4	OECD	OECD	OECD	OECD[c]	OECD
Korea	1983:1–2002:4	NSO	Bank of Korea	Bank of Korea	NSO[b]	IMF
Mexico	1981:1–2002:4	OECD	OECD	OECD	OECD[d]	OECD
Netherlands	1980:1–2002:4	OECD	OECD	OECD	OECD[d]	OECD
Spain	1980:1–2002:4	OECD	OECD	OECD	OECD[b]	OECD
Sweden	1980:1–2002:4	OECD	OECD	OECD	OECD[d]	OECD
United Kingdom	1980:1–2002:4	OECD	OECD	OECD	OECD[b]	OECD
United States	1980:1–2002:4	OECD	OECD	OECD	OECD[b]	OECD

a. HKMA: Hong Kong Monetary Authority. IMF: from *International Financial Statistics*. OECD: from *Main Economic Indicators* and *Quarterly Labour Force Statistics* (various issues); unemployment corresponds in all cases to the standardized rate. NSO: national statistics office.
b. All sectors.
c. Industry.
d. Manufacturing
e. Constructed based on trade participation.

Table 2. Dickey–Fuller Unit Root Tests[a]

Country	Real wage ($w - p^p$)		Real exchange rate (q)		Output (y)		Unemployment (u)	
Australia	–1.75	(–3.46)	–2.81	(–2.89)	–1.68	(–3.46)	–2.81	(–2.90)
Austria	–2.02	(–3.46)	–1.93	(–2.89)	–2.56	(–3.47)	–3.08	(–2.89)
Belgium	–1.12	(–3.46)	–3.19	(–3.46)	–2.89	(–3.46)	–4.20	(–3.47)
Canada	–2.43	(–3.46)	–2.32	(–3.46)	–2.40	(–2.90)	–3.86	(–3.47)
Chile	–1.14	(–2.90)	–1.37	(–2.91)	–2.40	(–2.91)	–2.82	(–2.90)
Colombia	–2.47	(–3.47)	–1.91	(–2.90)	–1.86	(–2.90)	–2.75	(–2.90)
Denmark	–2.28	(–3.49)	–4.78	(–2.91)	–2.89	(–3.49)	–2.74	(–3.49)
France	–2.61	(–3.46)	–2.52	(–2.89)	–2.57	(–3.46)	–2.10	(–2.90)
Germany	–2.94	(–3.46)	–1.94	(–2.89)	–0.66	(–2.89)	–2.63	(–2.89)
Hong Kong	–1.84	(–2.90)	–1.46	(–2.90)	–1.39	(–2.89)	–3.78	(–3.48)
Italy	–2.72	(–2.90)	–1.96	(–2.90)	–1.94	(–3.46)	–2.54	(–2.89)
Korea	–2.46	(–3.47)	–2.77	(–2.90)	–1.61	(–2.89)	–3.44	(–2.89)
Mexico	–1.47	(–2.91)	–2.52	(–2.91)	–2.85	(–3.48)	–2.44	(–2.89)
Netherlands	–1.18	(–2.89)	–4.73	(–3.46)	–0.65	(–2.90)	–4.21	(–3.46)
Spain	–2.86	(–3.47)	–1.85	(–2.89)	–2.69	(–3.46)	–2.21	(–2.89)
Sweden	–1.48	(–3.46)	–2.87	(–3.46)	–1.93	(–3.46)	–2.17	(–2.89)
United Kingdom	–2.14	(–3.46)	–2.23	(–2.89)	–2.40	(–3.46)	–4.14	(–3.46)
United States	–0.25	(–3.46)	–1.55	(–2.89)	–3.01	(–3.46)	–3.04	(–3.46)

Source: Authors' calculations.
a. The values in parentheses are 5 percent critical values.

shock on unemployment as permanent, even in the most rigid economies.[8] We therefore estimated the following stationary VAR, imposing the long-run restrictions above described: $[\Delta(w - p^p), \Delta q, \Delta y, u]'$. For the purpose of comparison, we also run a three-variable (closed economy) VAR equivalent to the one found in Balmaseda, Dolado, and López-Salido (2000): $[\Delta(w - p^p), \Delta y, u]'$. Most VARs were estimated using two lags based on the LM multivariate residual test for autocorrelation and the other regular criteria.[9]

2.3 Estimation Results

Given the large number of economies in our sample, we decided to report the impulse responses of a small subgroup of countries with varying degrees of labor market flexibility (see figure 2). The confidence intervals were obtained with a bootstrap procedure using 500 replications.[10] In general, a positive productivity shock causes real wages to increase in both the short and long terms. When a terms-of-trade shock hits the economy, real (producer) wages increase only in the short run.[11] In the case of positive labor-force shocks, the response of real wages tends to be negative in the short run, but it is insignificant in several cases. Real wages also fall in response to a monetary shock (expansion), most notably in the cases of Chile, Colombia, and the United States. In Korea and Sweden wages are procyclical, and they do not move in the United Kingdom.

The real exchange rate tends to appreciate when a positive productivity shock strikes the economy. After a positive terms-of-trade shock, the real exchange rate appreciates in both the short and long

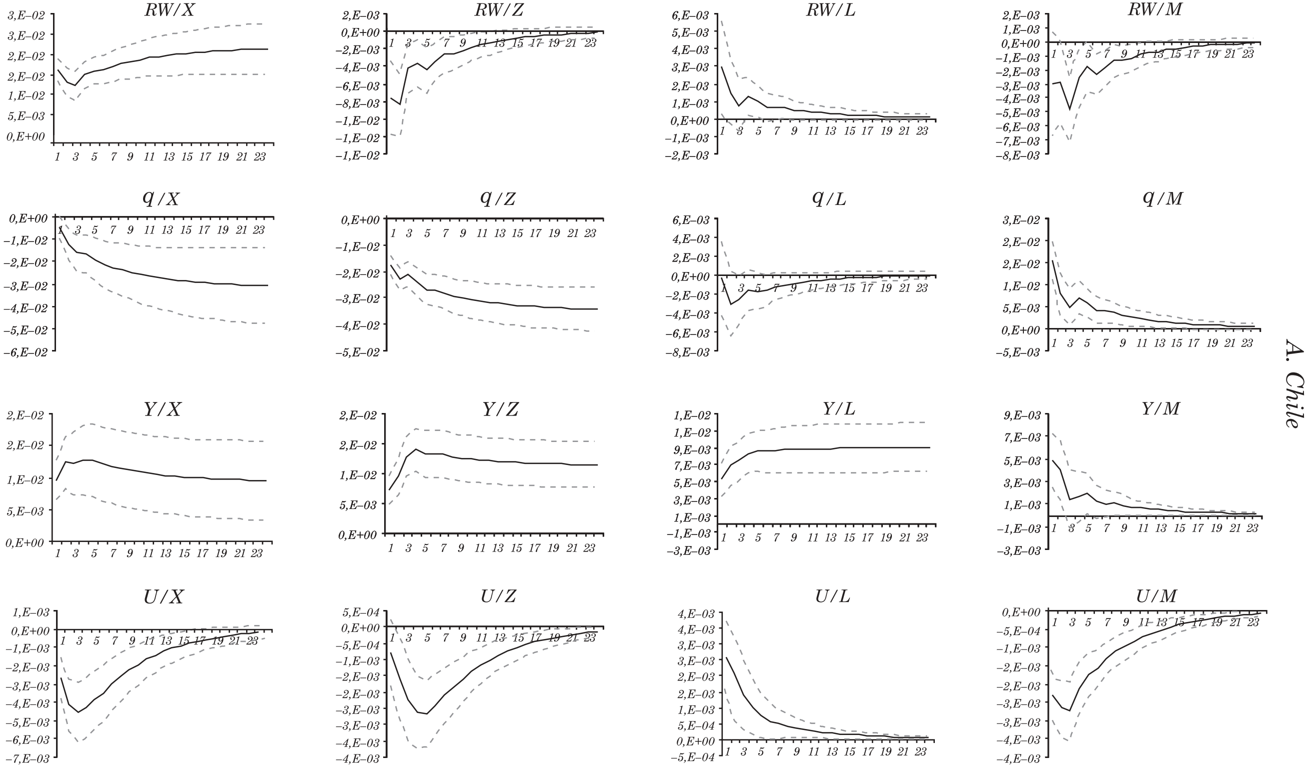
8. We also performed cointegration tests for all countries (as in Balmaseda, Dolado, and López-Salido, 2000); the null hyphotesis of no cointegration among the integrated variables $[w - p, q, y]$ was not rejected. With respect to the empirical rejection of the absence of integration in unemployment series, we agree with several authors mentioned in the paper who treat unemployment as an I(0) process, regardless of its severe persistence in the short run. From a theoretical perspective, which we consider the most relevant, the unemployment rate can hardly be considered a variable with a forecast of infinite variance.

9. We refer to the Akaike, Schwartz, and Hannan-Quinn information criteria. Kilian and Ivanov (2000) analyze which criterion performs better for VARs with different sample sizes.

10. See Benkwitz, Lütkepohl, and Wolters (2001) for an analysis of alternative bootstrap procedures.

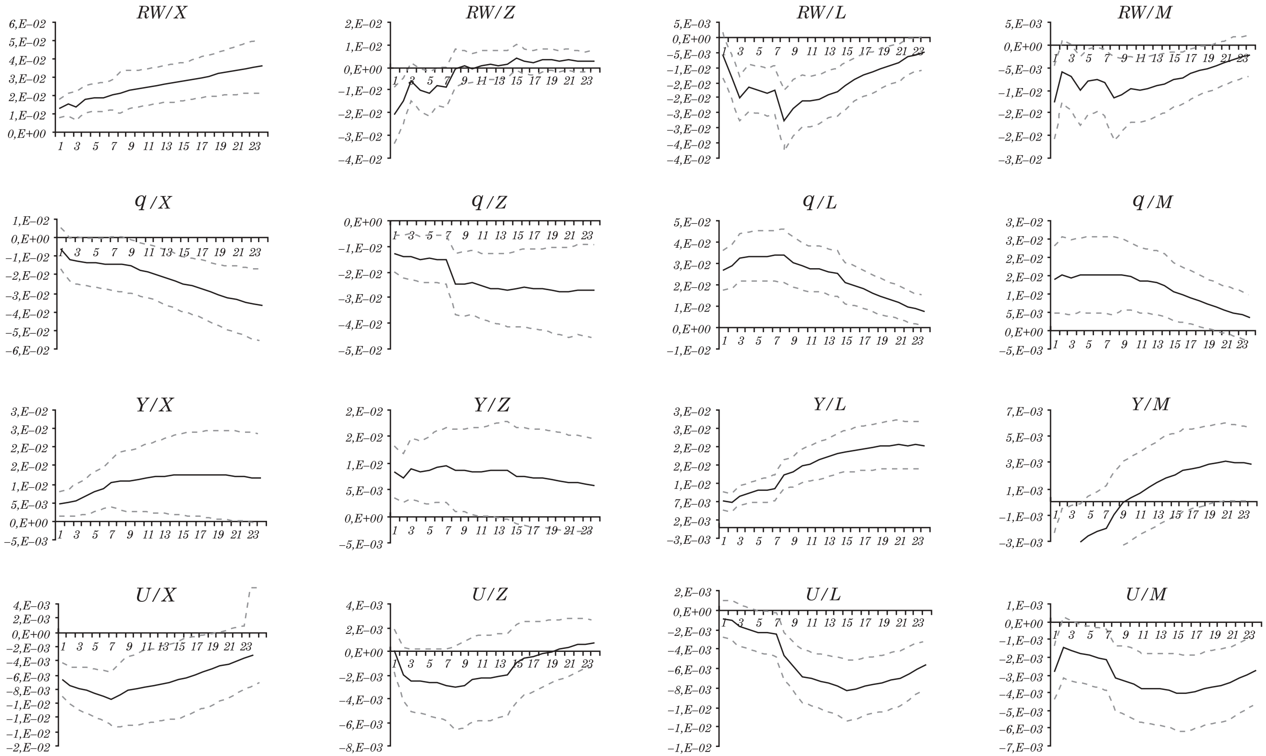
11. In Chile and Colombia, real wages go the wrong way, falling in the short run after a positive terms-of-trade shock.

Figure 2. Impulse-Response Functions to Structural Shocks, Selected Countries



Source: Authors' calculations.

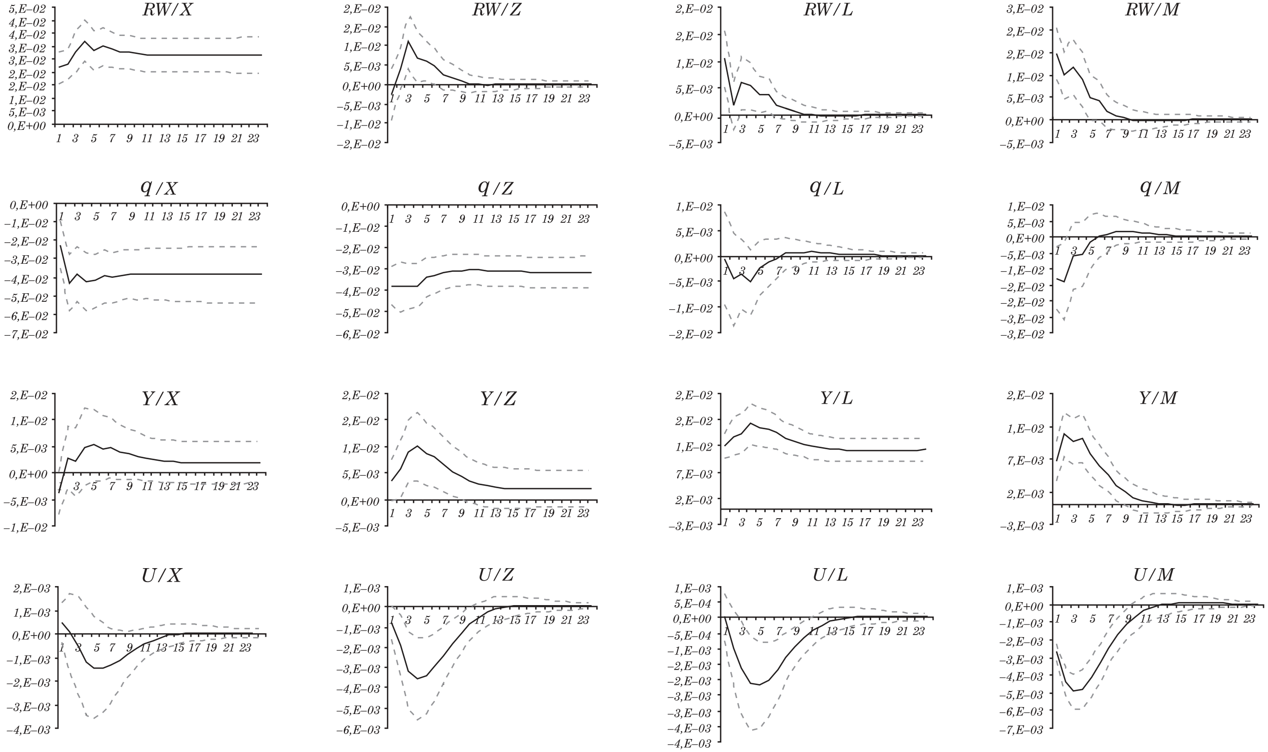
Figure 2. (continued)



B. Colombia

Source: Authors' calculations.

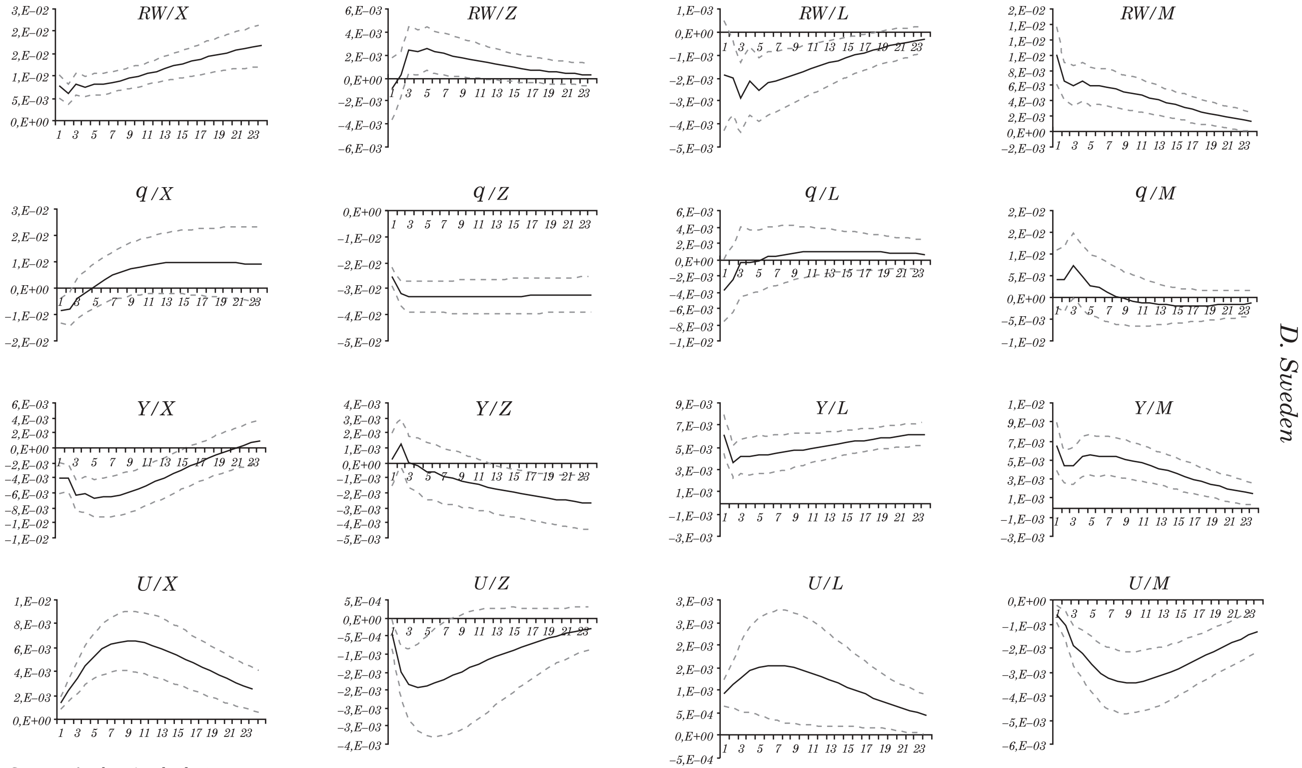
Figure 2. (continued)



C. Korea

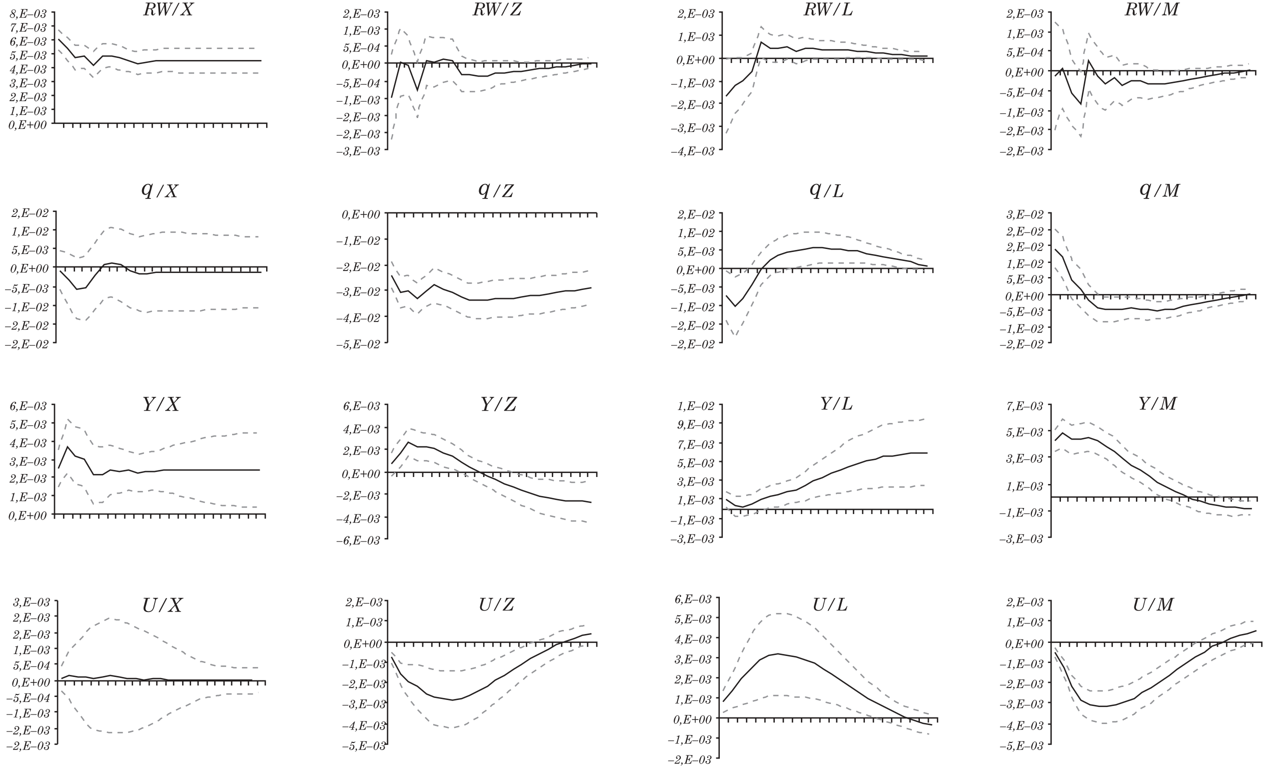
Source: Authors' calculations.

Figure 2. (continued)



Source: Authors' calculations.

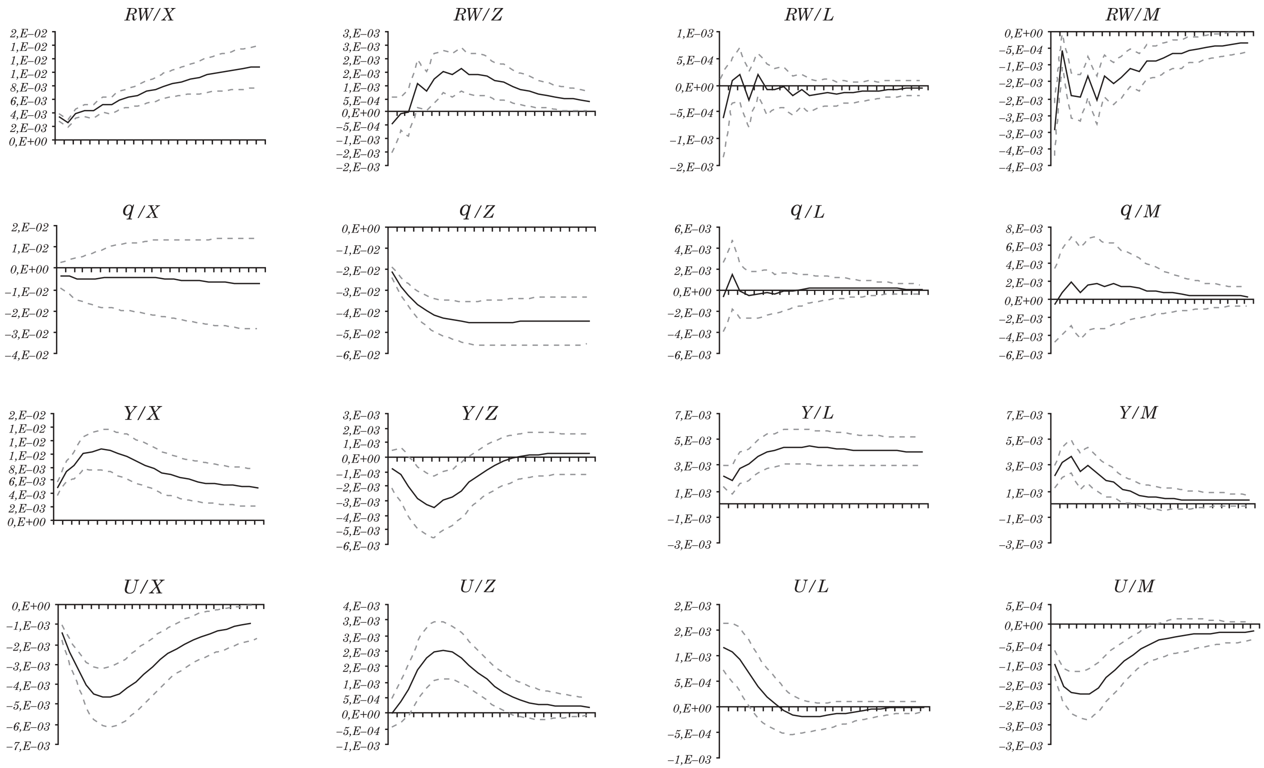
Figure 2. (continued)



E. United Kingdom

Source: Authors' calculations.

Figure 2. (continued)



F. United States

Source: Authors' calculations.

Table 3. Labor Market Rigidity Indexes

Country	*Average unemployment*[a]	*Closed economy (three variables)*	*Open economy (four variables)*
Hong Kong	3.11	4.25	4.25
Korea	3.37	3.08	4.33
Chile	8.27	5.67	4.63
Mexico	3.39	5.33	5.50
United States	6.30	4.42	5.75
Canada	9.17	5.75	6.75
Netherlands	5.87	6.13	7.33
United Kingdom	7.37	6.92	7.38
Australia	7.96	7.58	7.50
France	9.90	7.63	9.00
Belgium	11.63	9.25	9.38
Italy	10.61	9.58	9.58
Denmark	6.32	9.17	9.67
Austria	5.51	8.58	10.50
Germany	6.85	9.33	10.75
Spain	17.89	11.42	11.38
Sweden	4.94	12.83	12.13
Colombia	12.90	24.50	22.75
Correlation with unemployment		0.54	0.53

Source: Authors' calculations.
a. Corresponds to the OECD standardized rate of unemployment.

terms. The response of the real exchange rate after a labor-force shock is seldom significant. The real exchange rate tends to increase in the short run as a result of a monetary expansion shock. Finally, the response of unemployment after a positive productivity shock is not clear-cut: it increases in many countries, but it does the opposite in several others. Unemployment tends to increase after a labor-supply shock and to decrease with either a terms-of-trade shock or a positive monetary disturbance, but there are exceptions to these trends.

Labor Market Rigidity Index

Table 3 shows the rankings we built by computing the average half-life of the unemployment responses for all shocks. The table reports rankings estimated for both closed- and open-economy specifications (based on three and four variables, respectively). Korea and Hong Kong are the most flexible countries, followed by Chile, Mexico, and the United States. This is consistent with recent evidence for the Korean economy, where unemployment peaked after the Asian crisis but quickly returned to its previous level. On the other hand, Chile is still relatively flexible in an international context despite the two labor reforms that may have introduced some rigidity into the labor market after 1990.

At the other end of the spectrum, Germany, Sweden, Spain, and Colombia are ranked as the most rigid labor markets. In Colombia, unemployment increased sharply during the 1999 crisis, reaching 20 percent, but it decreased very slowly in 2003. In the middle range of rigidity are Australia, Austria, Belgium, Denmark, and the Netherlands among others. The ranking has a positive and significant correlation with average unemployment, as can be expected.

3. Assessment of the Chilean Labor Market Rigidity Ranking

Can we infer from the rankings presented in table 3 that Chile is a flexible economy in comparison with most of the countries considered? The evidence presented suggests precisely that. Nevertheless, we must mention a number of caveats before precipitating conclusions.

First, the ranking is based on unemployment persistence. Unemployment, however, is a net measure between labor supply and employment, so the rigidity implied by its persistence hides the true origin of market frictions. A rigid labor market can be dominated by an inefficient process of job reallocation, in which case the persistence is best attributed to rigid employment creation. Alternatively, the source of rigidity could be labor market institutions (that is, social security benefits) that foster persistent job search even when hiring prospects are low. In this context, the different correlations between unemployment and labor market participation across countries could help to disentangle the dominant source of rigidity. The fall in labor participation during high unemployment periods would suggest that at least part of this supposed flexibility comes from people exiting unemployment toward inactivity, not employment.

Second, as mentioned in the introduction, a growing part of employment creation is in the form of low-skilled, informal positions of self-employment. Any assessment of flexibility must therefore take into account the extent to which the adjustment actually occurs outside the more formal market. Unfortunately, lack of comparable data on informality among countries over long periods impedes further insights on this issue.

Finally, although the 1990s as a whole were characterized by increasing labor market regulation, reform critics point out that the structural change in adjustment capacity was only triggered by the recent regulations of 1999–2001. Whether this view is sustainable cannot be determined by the methodology presented. The empirical approach adopted here is limited to the small number of data points after the reform. When we ran the same VAR for Chile from 1986 to 1998 and from 1990 to 2002, we found no significant difference relative to the responses of unemployment using the whole sample.

Given these caveats, our performance-based ranking suggests that, when examined using a methodology that properly distinguishes shock responses from the persistent behavior of unemployment, the Chilean labor market appears to be relatively flexible.

4. Conclusions

In this paper, we ranked labor market rigidity for a sample of eighteen countries, with the purpose of characterizing the relative rigidity of the labor market in Chile. We analyzed the dynamic responses of unemployment in the presence of macroeconomic shocks identified with a structural VAR following Blanchard and Quah (1989). The setting of the empirical approach and the interpretation of the shocks are based on a model with rigidity in the labor market through the insider-outsider bargaining setup, which was popularized by Blanchard and Summers (1986) and is extended for an open economy in Albagli, García, and Restrepo (2004).

The restrictions derived from the model imply that in the long run, real producer wages grow only with productivity; the real exchange rate depends only on productivity and the terms of trade; output is affected by productivity, terms-of-trade, and labor-force shocks; and unemployment converges to its natural rate despite short-run disturbances. The model allows us to build an indicator of rigidity based on unemployment persistence and depending only on the rigidity coefficient of the wage-setting equation: namely, the half-life of unemployment after the shocks. We used this indicator to build rankings that served as the basis of comparison within our sample.

We found that Korea and Hong Kong have the most flexible labor markets, followed by Chile, Mexico and the United States. At the other end of the ranking, Germany, Sweden, Spain, and Colombia are the most rigid countries.

These findings support the view that despite the strict labor regulation introduced in the 1990s, labor market outcomes in Chile are still far from the European experience of high, persistent unemployment. Whether this results from weak institutional enforcement and whether the more recent reforms account for a structural change in labor market dynamics are interesting challenges for future research.

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Microeconomic Flexibility in Latin America

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Latin American economies have begun to leave behind some of the most primitive sources of macroeconomic fluctuations. Policy concern is gradually shifting toward increasing microeconomic flexibility. This is a welcome trend since microeconomic flexibility, which facilitates the ongoing process of creative-destruction, is at the core of economic growth in modern market economies.

But how poorly are these economies doing along this flexibility dimension? Answering this question requires measuring the important, but elusive concept of microeconomic flexibility. One way of doing this is to look directly at regulation, which is perhaps the main institutional factor hindering or facilitating microeconomic flexibility. Extensive studies examine labor market regulation, in particular. Heckman and Pagés (2000), for example, document that "even after a decade of substantial deregulation, Latin American countries remain at the top of the Job Security list, with levels of regulation similar to or higher than those existing in the highly regulated south of Europe." This is important work. However, in practice microeconomic flexibility depends not only on labor market regulation, but also on a wide variety of factors, including technological options, the nature of the production process, the political environment, the efficiency and

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biases of labor courts, cultural variables, and accepted practices. While studies of rules and regulation are useful for eventual policy formulation, they are unlikely to provide us with the big picture of a country's flexibility any time soon: understanding the complex interactions of different regulations and environments is a valuable, but very slow process.

At the other extreme, one can look directly at outcomes: how much factor reallocation occurs in different countries and episodes? This is also a useful exercise, but it is equally incomplete, since there is no reason to expect the same degree of aggregate flows in countries facing different idiosyncratic and aggregate shocks. It is always difficult to determine whether the observed reallocation is abnormally high or low, because the counterfactual is not part of the statistic.

A third approach, which remedies some of the main weaknesses of the previous two, is to measure microeconomic flexibility by the speed at which establishments reduce the gap between their labor productivity and the marginal cost of such labor. An economy is said to be inflexible at the microeconomic level if these gaps persist over time. Conversely, a very flexible economy, firm, or establishment is one in which gaps disappear quickly as a result of prompt adjustment. This is the approach we follow here, while extending a methodology developed in Caballero and others (2004). The main advantage of this methodology over conventional partial adjustment estimates is its ability to use limited information efficiently by correcting standard biases that are often present when estimating such models. Our methodology also allows for nonlinearities and state-dependent responses of employment to productivity gaps, as in Caballero and Engel (1993).[1]

We use establishment-level observations for all the Latin American economies for which we had access to fairly reliable data: Chile, Mexico, and, to a lesser extent, Brazil, Colombia, and Venezuela. All in all, this provides us with about 140,000 observations.

In the first part of the paper, we document the main features of adjustment for these economies. Our findings include the following:

1. Our definition of microeconomic flexibility refers to the speed at which establishments react to changing conditions, not to whether the labor market is flexible in responding to aggregate shocks. Thus, a labor market regulation that makes the real wage rigid will increase the unemployment response to aggregate shocks—that is, it will exhibit macroeconomic inflexibility—yet this will not be part of our measure of microeconomic inflexibility.

— While more inflexible than the United States, the economies of Brazil, Chile, and Colombia exhibit a relatively high degree of microeconomic flexibility, with over 70 percent of labor adjustment taking place within a year, on average (over time). Mexico ranks lower with about 60 percent of adjustment within a year, and Venezuela is the most inflexible of these economies, with slightly over 50 percent of adjustment within a year.

— In all the economies in our sample except Venezuela, small establishments (with fewer than the median number of employees) are substantially less flexible than large establishments (above the seventy-fifth percentile of employees). In Brazil, the former establishments close about 67 percent of their gap within a year, while the latter close about 81 percent. In Chile, the figures are 69 percent and 78 percent, respectively; in Colombia, 68 percent and 79 percent; in Mexico, 56 percent and 61 percent; and in Venezuela, 53 percent for both.

— It follows from the previous finding that the behavior of large establishments is primarily behind the substantial differences in flexibility across some of the economies we study. It may well be the case that large companies in Venezuela and Mexico are more insulated from competitive pressures than are their counterparts in Chile, Colombia, and Brazil.

— All these economies present evidence of an increasing hazard. That is, establishments are substantially more flexible with respect to large gaps than to small ones. This points to the presence of significant fixed costs of adjustment, which may have a technological or institutional origin.

— Increasing hazard is particularly pronounced in large establishments in the relatively more flexible economies. In fact, most of the additional flexibility experienced by large establishments in these economies is due to their rapid adjustment when gaps become large. For example, when gaps are below 25 percent in Chile, small establishments have an adjustment coefficient of 0.50, while large establishments have a coefficient of 0.51. For deviations above 25 percent, the coefficient is 0.79 for small establishments and 0.93 for large establishments. The patterns are similar in Brazil and Colombia, yet less pronounced in Mexico and Venezuela.

The second part of the paper focuses on Chile, which has the only long panel in our sample. We are thus able to explore the evolution of the economy's microeconomic flexibility over time. Our main findings are threefold:

- Microeconomic flexibility in Chile experienced a significant decline toward the end of our sample (1997–99). The coefficient fell from an average adjustment coefficient of 0.77 for the three years prior to the Asian/Russian crisis episode to 0.69 in the aftermath of the crisis.
- When the adjustment hazard is assumed to be constant, the decline in flexibility appears to subside toward the end of the sample. This finding is lost, however, when the hazard is allowed to be increasing, with no evidence of recovery. The reason for the misleading conclusion with a constant hazard is that toward the end of the sample, there is a sharp rise in the share of establishments with large (negative) gaps, to which establishments naturally react more strongly under increasing hazards.
- While it is too early to tell whether the decline we uncover is purely cyclical or whether it reflects a structural change, we can make a few interesting observations. First, much of the decline in flexibility is due to a decline in the flexibility of large establishments. Second, the speed of response to negative gaps remained fairly constant, while the speed at which establishments adjust to labor shortages slowed dramatically. This "reluctance to hire" may reflect pessimism regarding future conditions not captured in the contemporaneous gap. This is unlikely to be the only factor, however, because we do not observe a rise in the speed of firing (for a given hazard).[2] Finally, the sharpest decline in flexibility came from establishments in sectors that normally experience less restructuring, either because they experience smaller shocks than other sectors or because they are characterized by higher technological and institutional inflexibility. If either form of inflexibility is responsible for reduced restructuring, then the cost of the decline in flexibility can potentially be very large, since inflexible establishments spend a significant amount of time away from their frictionless optimum.

The last part of the paper explores a different metric for the degree of inflexibility and its economic impact. By impairing worker movements from less to more productive units, microeconomic inflexibility reduces

2. While we did see an increase in the speed of firing, this is accounted for by the interaction of a prolonged contraction with an increasing hazard.

aggregate output and slows economic growth. We develop a simple framework for quantifying this effect. Our findings suggest that the aggregate consequences of microeconomic inflexibilities in Latin America are significant. In particular, the impact of the decline in microeconomic flexibility in Chile following the Asian crisis is, in itself, large enough to account for a substantial fraction of the decline in total factor productivity (TFP) growth in Chile since 1997, which fell from an annual average of 3.1 percent for the preceding decade to about 0.3 percent after that date. Moreover, if the decline were to persist, it could permanently shave off almost half a percent from Chile's structural growth rate.

Section 1 presents the methodology, while section 2 describes the data. Section 3 characterizes average microeconomic flexibility in the Latin American economies in our data. Section 4 explores the case of Chile in more detail and describes the evolution of its index of flexibility. Section 5 presents a simple model with which to map microeconomic inflexibility into growth outcomes. Section 6 concludes and is followed by two appendixes.

1. Methodology and Data

The starting point for our methodology is a simple adjustment hazard model, in which the change in the number of filled jobs in establishment i in sector j between time $t-1$ and t is a probabilistic function (at least to the econometrician) of the gap between desired and actual employment (before the adjustment):

$$\Delta e_{ijt} = \psi_{ijt}(e^*_{ijt} - e_{ij,t-1}) \tag{1}$$

where e_{ijt} and e^*_{ijt} denote the logarithm of employment and desired employment, respectively. The random variable ψ_{ijt} is assumed to be independent and identically distributed (i.i.d.) both across establishments and over time; it takes values in the interval [0, 1] and has mean λ and variance $\alpha\lambda(1-\lambda)$, with $0 \leq \alpha \leq 1$. The case $\alpha = 0$ corresponds to the standard quadratic adjustment model; the case $\alpha = 1$ represents the Calvo (1983) model. The parameter λ captures microeconomic flexibility. As λ goes to one, all gaps are closed quickly and microeconomic flexibility is maximum. As λ decreases, microeconomic flexibility declines.

Equation (1) hints at two important components of our methodology: we need to find a measure of the employment gap, $e^*_{ijt} - e_{ijt-1}$, and an estimation strategy for the mean of the random variable ψ_{ijt}, λ. We describe both ingredients in detail below. In a nutshell, we construct estimates of e^*_{ijt}, the only unobserved element of the gap, by solving the optimization problem of the firm as a function of observables such as labor productivity and a suitable proxy for the average market wage. We estimate λ from equation (1), based on the large cross-sectional size of our sample and the well-documented fact that there are significant idiosyncratic components in the realizations of the gap and ψ_{ijt}.

An important aspect of our methodology is to find an efficient method of removing fixed effects, while at the same time avoiding the standard biases present in dynamic panel estimation.[3] The model we develop also leads to a standard dynamic panel formulation, namely,[4]

$$Gap_{ijt} = (1-\lambda)\Delta e^*_{ijt} + (1-\lambda)Gap_{ij,t-1} + \varepsilon_{ijt} \ . \tag{2}$$

We report results for this specification after presenting our main results. As we will show, the results are consistent with the estimates we obtain based on equation (1) and therefore provide a useful robustness check, although they are considerably less precise. Our methodology may thus be viewed as an alternative, for the particular problem at hand, that uses data more efficiently than standard dynamic panel estimation techniques. Finally, our methodology can be adapted to the case in which flexibility evolves over time (see section 4), which is not the case with standard panel techniques.[5]

1.1 Details

Output and demand faced by an establishment are given by

$$y = \mathrm{a} + \alpha e + \beta \ \mathrm{h} \tag{3}$$

3. As documented, for example, in Arellano and Bond (1991).
4. The gap below could be either before or after adjustments take place.
5. Panel techniques along the lines of Arellano and Bond (1991) cannot be extended to the case in which the economy wide average λ varies over time, since the instruments used in these procedures are no longer valid.

and

$$p = d - \frac{1}{\eta} y \tag{4}$$

where *y*, *p*, *e*, *a*, *h*, and *d* denote firm output, price, employment, productivity, hours worked, and demand shocks, and η is the price elasticity of demand. We let $\gamma \equiv (\eta - 1) / \eta$.[6] All variables are in logs.

Firms are competitive in the labor market but pay wages that are increasing in the average number of hours worked, according to[7]

$$w = w^0 + \mu(h - \bar{h}) \tag{5}$$

where $\bar{h}$ is constant over time and interpreted below.[8]

A key assumption is that firms only face adjustment costs when they change employment levels, not when they change the number of hours worked.[9] It follows that the firm's choice of hours in every period can be expressed in terms of its current level of employment, by solving the corresponding first-order condition for hours.

In a frictionless labor market, the firm's employment level also satisfies a first-order condition for employment. Our functional forms then imply that the optimal choice of hours does not depend on the employment level.[10] We denote the corresponding employment level by $\hat{e}$ and refer to it as the static employment target.[11] This leads to the following relation between the employment gap and the hours gap:

$$\hat{e} - e = \frac{\mu - \beta\gamma}{1 - \alpha\gamma}(h - \bar{h}) \; . \tag{6}$$

6. To ensure interior solutions, we assume $\eta > 1$ and $\alpha\gamma < 1$.

7. The expression below should be interpreted as a convenient approximation for $w = k^0 + \log(H^{\mu} + \Omega)$, with w^o and μ determined by k^o and Ω.

8. To ensure interior solutions, we assume $\alpha\mu > \beta$ and $\mu > \beta\gamma$.

9. See Sargent (1978) and Shapiro (1986).

10. A patient calculation shows that

$$\bar{h} = \frac{1}{\mu} \log\left(\frac{\beta\Omega}{\alpha\mu - \beta} \right) .$$

11. We have

$$\hat{e} = C + \frac{1}{1 - \alpha\gamma}\left[d + a - w^0 \right]$$

where ***C*** is a constant that depends on μ, α, β, and γ.

This is the expression used by Caballero and Engel (1993). It is not useful in our case, since we do not have information on hours worked. However, the argument used to derive equation (6) can also be used to express the employment gap in terms of the marginal labor productivity gap:

$$\hat{e} - e = \frac{\phi}{1 - \alpha\gamma}(v - w^0)$$

where v denotes marginal productivity; $\phi \equiv \mu / (\mu - \beta\gamma)$ is decreasing in the elasticity of the marginal wage schedule with respect to average hours worked, $\mu - 1$; and w^o was defined in equation (5). This result is intuitive: the employment response to a given deviation of wages from marginal product will be larger if the marginal cost of the alternative adjustment strategy—changing hours—is higher. Also note that $\hat{e} - e$ is the difference between the static target, $\hat{e}$, and realized employment, not the dynamic employment gap, $e^*_{ijt} - e_{ijt}$, related to the term on the right-hand side of equation (1). We assume, however, that demand, productivity, and wage shocks follow a random walk.[12] Consequently, e^*_{ijt} is equal to $\hat{e}_{ijt}$ plus a constant, δ_τ.[13] It follows that

$$e^*_{ijt} - e_{ij,t-1} = \frac{\phi}{1 - \alpha\gamma_j}(v_{ijt} - w^0_{ijt}) + \Delta e_{ijt} + \delta_t \tag{7}$$

where we have allowed for sector-specific differences in γ.

We estimate the marginal productivity of labor, v_{ijt}, using output per worker multiplied by an industry-level labor share, assumed constant over time.

Two natural candidates to proxy for w^0_{ijt} are the average (across each industry, at a given point in time) of either observed wages or observed marginal productivities. The former is consistent with our assumption of a competitive labor market; the latter may be expected to be more robust in settings with long-term contracts and multiple forms of worker compensation, where the wage may not represent the actual marginal cost of labor.[14] We performed our estimations

12. Given the preceding footnote, it suffices that $d + \gamma\alpha - w^o$ follows a random walk.

13. To allow for variations in future expected growth rates of a and d, the constant δ is allowed to vary exogenously over time.

14. While we have assumed a simple competitive market for the base wage within each firm (i.e., the wage for regular hours), our procedure can easily accommodate rent sharing as part of the wage-setting mechanisms (with a suitable reinterpretation of some parameters, but not λ).

using both alternatives, and we found no discernible differences. This suggests that statistical power comes mainly from the cross-section dimension, that is, from the well-documented large magnitude of the idiosyncratic shocks faced by establishments. We report the more robust alternative and approximate w^o by the average marginal productivity, which leads to

$$e^*_{ijt} - e_{ij,t-1} = \frac{\phi}{1-\alpha\gamma_j}(v_{ijt} - v_{.jt}) + \Delta e_{ijt} + \delta_t \equiv Gap_{ijt} + \delta_t \ . \tag{8}$$

The above expression ignores systematic variations in labor productivity that may occur across establishments, which would tend to bias estimates of the speed of adjustment downward. In appendix A we provide evidence in favor of incorporating this possibility by subtracting from $(v_{ijt} - v_{.jt})$ in equation (8) a moving average of lagged relative productivity by establishment, $\hat{\theta}_{ijt}$.[15] The resulting expression for the estimated employment gap is[16]

$$e^*_{ijt} - e_{ij,t-1} = \frac{\phi}{1-\alpha\gamma_j}(v_{ijt} - \hat{\theta}_{ijt} - v_{.jt}) + \Delta e_{ijt} + \delta_t \equiv Gap_{ijt} + \delta_t \ . \tag{9}$$

Finally, we estimate ϕ (related to the substitutability between hours worked and employment) using

$$\Delta e_{ijt} = -\frac{\phi}{1-\alpha\gamma_j}(\Delta v_{ijt} - \Delta v_{.jt}) + \kappa_t + \upsilon_{it} + \Delta e^*_{ijt} \equiv -\phi z_{ijt} + \kappa_t + \varepsilon_{ijt} \tag{10}$$

where κ is a year dummy, $\Delta e_{ijt}*$ is the change in the desired level of employment, and

$$z_{ijt} = (\Delta v_{ijt} - \Delta v_{.jt}) / \left(1 - \alpha\gamma_j\right) \ .$$

15. Where

$$\hat{\theta}_{ijt} \equiv \frac{1}{2}\left[\left(v_{ijt-1} - v_{.jt-1}\right) + \left(v_{ijt-2} - v_{.jt-2}\right)\right].$$

The alternative specification, with relative wages instead of relative marginal productivities, leads to almost identical results.

16. Where $\alpha\gamma_j$ is constructed using the sample median of the labor share for sector j across years and countries (Brazil, Chile, Colombia, Mexico, and Venezuela).

By assumption, Δe^{*}_{ijt} is i.i.d. and independent of lagged variables. To avoid endogeneity and measurement error bias, we estimate equation (10) using $(\Delta w_{ij,t-1} - \Delta w_{.jt-1})$ as an instrument for $(\Delta v_{ijt} - \Delta v_{.jt})$.[17] Table 1 reports the estimation results of equation (10) across the countries in our sample.[18] We report estimates both with and without the one percent of extreme values for the independent variable. Based on the estimates reported in table 1, we chose a common value of ϕ equal to 0.40, to facilitate comparison across countries.

Table 1. Estimating ϕ[a]

	Country			
Parameter	*Chile*	*Colombia*	*Mexico*	*Venezuela*
ϕ (with extreme values)	0.460	0.414	0.372	0.336
	(0.028)	(0.035)	(0.033)	(0.108)
ϕ (without extreme values)	0.495	0.394	0.365	0.317
	(0.037)	(0.035)	(0.037)	(0.118)
No. observations	21,149/20,938	20,268/20,065	27,752/27,475	2,906/2,877

a. The parameter ϕ is estimated using equation 10 in the text. The regression is run both with and without the one percent of extreme values for the independent variable. Brazil is excluded from the analysis because wage data are not available. Robust standard errors in parentheses.

1.2 Summary

Our methodology has four advantages over previous specifications used to estimate cross-country differences in speed of adjustment. First, it only requires data on nominal output and the employment level, two standard and well-measured variables in most industrial surveys. Most previous studies on adjustment costs require measures of real output or an exogenous measure of sector demand.[19] Second,

17. We lag the dependent variable because it is correlated with the error term, and we use lagged wages to instrument lagged labor productivity to avoid measurement errors.

18. We do not have wage data for Brazil, so we cannot estimate the parameter for this country.

19. Abraham and Houseman (1994), Hamermesh (1993), and Nickel and Nunziata (2000) evaluate the differential response of employment to observed real output. A second option is to construct exogenous demand shocks. Although this approach overcomes the concerns associated with real output, it requires constructing an adequate sectoral demand shock for every country. Burgess and Knetter (1998) and Burgess, Knetter, and Michelacci (2000), for example, use the real exchange rate as their demand shock. The estimated effects of the real exchange on employment are usually marginally significant and often of the opposite sign than expected.

it summarizes in a single variable all shocks faced by a firm. This feature allows us to increase precision and, therefore, the power of hypothesis testing, as well as to study the determinants of the speed of adjustment using interaction terms. Third, in contrast with standard panel techniques, the methodology we develop can be used when the average microeconomic flexibility is evolving over time (see section 4). Finally, our approach can be extended to incorporate nonlinearities in the adjustment function—that is, the possibility that ψ in equation (1) depends on the gap before adjustments take place. This feature also turns out to be useful.

Summing up, in our basic setup we estimate the microeconomic flexibility parameter, λ, from

$$\Delta e_{ijt} = \left(Gap_{ijt} + \delta_t\right) + \varepsilon_{ijt} \tag{11}$$

where Gap_{ijt} is proportional to the gap between marginal labor productivity and the market wage. To correct for labor heterogeneity across establishments, a fixed effect is also included in the gap measure. This fixed effect is estimated by the average labor productivity in the two preceding periods. As shown in appendix A, the resulting estimator is unbiased, on average. It forces us to discard only two time periods, and it can adapt to slow time variations in heterogeneity.

2. Data and Basic Facts

This section describes the data and sources used in the empirical analysis. These data are from manufacturing censuses and surveys conducted by national government statistical agencies in five Latin American countries: Brazil, Chile, Colombia, Mexico, and Venezuela. The variables used in our analysis are nominal output, employment, total compensation, and industry classification within the manufacturing sector (from the International Standard Industrial Classification, or ISIC, at three digits). For the case of Chile, we also use capital stock and a measure of cash flow defined as sales minus total input costs. In all countries, we include only plants that existed during the full sample period (continuous plants).

For Brazil, the data are from the annual manufacturing survey *(Pesquisa Industrial Anual)* conducted by the *Instituto Brasileiro de Geografia e Estatística*. This survey started in 1967 but underwent a severe methodological change in 1996, so we only use observations from 1996 to 2000. In the case of Chile, the data are from the Chilean manufacturing census *(Encuesta Nacional Industrial Anual)* conducted by the country's *Instituto Nacional de Estadísticas*. In principle, the survey covers all manufacturing plants in Chile with more than ten employees during the period 1979–97. However, the years before 1985 are characterized by large macroeconomic shocks and structural adjustments that introduce too much noise for our methodology to handle properly. We thus use only continuous plants from the period 1985–97. For Colombia, we use the Colombian manufacturing census *(Encuesta Anual Manufacturera y Registro Industrial)* conducted by the *Departamento Administrativo Nacional de Estadística*. The survey covers all manufacturing plants with more than twenty employees during the period 1982–99. For plants with less than twenty employees, only a random sample is covered. Again, we limit the sample to continuous plants in the period 1992–99 because of a methodological change in the survey in 1992.

For Mexico, we use the annual manufacturing survey *(Encuesta Industrial Anual)* conducted by the *Instituto Nacional de Estadística, Geografía e Informática*. The survey covers a random sample of establishments in the manufacturing sector during the period 1993–2000. Finally, the data for Venezuela are from the manufacturing survey *(Encuesta Industria Manufacturera)* conducted by the country's *Instituto Nacional de Estadística*. The survey covers all plants with more than fifty employees, and it has a yearly random sample for plants with less than fifty employees. As a result of changes in the methodology, we are only able to follow firms during the 1995–1999 period.

Table 2 presents the number of observations per size bracket (measured by the number of employees) for each of the five countries, for the relevant sample period. The coverage of plants by size differs across countries. Chile and Colombia have the largest coverage of small plants (fewer than fifty employees), whereas Venezuela's survey mainly covers large establishments.

In table 3, we compute the average job creation and job destruction for each country. We also report the simple average over time of

Table 2. Descriptive Statistics: Breakdown of Establishments by Size[a]

	Country				
Statistic	*Brazil*	*Chile*	*Colombia*	*Mexico*	*Venezuela*
No. observations	42,525	24,450	27,440	37,384	4,950
No. establishments	8,505	1,630	3,430	4,673	990
Observations by plant size (percent)					
Under 50 employees	15.9	56.7	45.1	21.0	9.9
50–99 employees	28.5	17.9	22.8	21.4	31.5
100–249 employees	28.9	15.4	19.5	29.4	33.7
250 employees and above	26.6	9.9	12.7	28.2	24.9
Period	1996–2000	1985–99	1992–99	1993–2000	1995–99

a. Only continuous plants are considered.

Table 3. Descriptive Statistics: Job Creation and Destruction[a]

	Country				
Statistic	*Brazil*	*Chile*	*Colombia*	*Mexico*	*Venezuela*
Total jobs in sample	2,555,035	169,813	461,441	1,214,776	233,746
Net change	–0.024	0.021	–0.013	0.018	–0.023
Job creation	0.074	0.080	0.072	0.071	0.069
Job destruction	0.098	0.059	0.086	0.053	0.091
Reallocation	0.173	0.139	0.158	0.123	0.160
Excess reallocation	0.135	0.099	0.124	0.086	0.125
Period	1997–2000	1986–99	1993–99	1994–2000	1996–99

a. Quantities reported are yearly averages over the sample period. Definition of all variables follows Davis, Haltiwanger, and Schuh (1996).

the net change in employment and the excess turnover (that is, the sum of job flows net of the change in employment stemming from cyclical factors). All statistics are defined following Davis, Haltiwanger, and Schuh (1996). These numbers suggest that microeconomic flexibility in these countries is limited: they are of the same order of magnitude as those of developed economies—which presumably need less restructuring than emerging economies that are still catching up—and substantially below economies such as Taiwan.[20]

20. See, for example, Caballero and Hammour (2000) and references therein.

3. Microeconomic Flexibility

In this section, we report our average (over time) flexibility findings. The basic results are reported in table 4. All of our regressions include year dummies, d_t. That is, for each country, we estimate the following equation (we now drop the sector j subscript):

$$\Delta e_{it} = d_t + \lambda Gap_{it} + \varepsilon_{it} \quad . \tag{12}$$

The first apparent result is that microeconomic flexibility is more limited in our economies than in the very flexible United States. In the latter, estimates of λ using annual data are much closer to one.[21]

Table 4. Average Flexibility Estimates[a]

	Country				
Explanatory variable	*Brazil*	*Chile*	*Colombia*	*Mexico*	*Venezuela*
Gap	0.701	0.724	0.722	0.581	0.539
	(0.004)	(0.005)	(0.005)	(0.004)	(0.014)
Summary statistic					
R^2	0.50	0.50	0.53	0.47	0.37
No. observations	25,260	20,979	20,375	27,757	2,941
Period	1998–2000	1988–99	1995–99	1995–2000	1997–99

a. All regressions include year dummies. All estimates are based on one regression per country, using all available observations, and are significant at the 1 percent level. Observations corresponding to extreme values of regressors (0.5 percent in right tail and 0.5 percent in left tail) are excluded. Robust standard errors are in parentheses.

Although comparisons must be interpreted with caution since the samples differ in number of observations, time periods, establishments' demographics, and so forth, we can identify a discernible pattern. Within the region, Brazil, Chile, and Colombia exhibit a relatively high degree of microeconomic flexibility with over 70 percent of labor adjustment taking place within a year. Mexico ranks lower with about 60 percent of adjustment within a year, and Venezuela is the least

21. For example, Caballero, Engel, and Haltiwanger (1997) find a quarterly λ exceeding 0.4 for U.S. manufacturing, which implies an annual λ of approximately 0.90.

flexible of these economies, with slightly more than 50 percent of adjustment within a year.

Our ranking is essentially uncorrelated with the ranking obtained by Heckman and Pagés (2000) and Djankov and others (2003) based on measuring labor market regulations (see table 5). For example, the Djankov and others (2003) index of job security places Venezuela at a level of flexibility similar to that of Brazil and Chile, while Colombia is significantly more flexible than all of the above.[22] These contrasting results lend support to our earlier motivation for adopting our approach in constructing a broad measure of microeconomic inflexibility.

Table 5. Comparing Flexibility Measures[a]

	Country				
Measure	*Brazil*	*Chile*	*Colombia*	*Mexico*	*Venezuela*
Job security index (Heckman and Pagés, 2000)	3.04	3.38	3.79	3.16	4.54
Job security index (Djankov and others, 2003)	0.69	0.62	0.31	0.71	0.64
Excess reallocation (from table 3)	0.135	0.099	0.124	0.086	0.125
Microeconomic flexibility index (this paper)	0.701	0.724	0.722	0.581	0.539

a. Flexibility is decreasing in the index for the first two measures and increasing for the remaining two measures. Yearly values for 1990–99 are available for the Heckman-Pagés index only; the numbers reported for this index are the average over the sample period (years before 1990 are proxied by the 1990 value, and years after 1999 by the 1999 value).

Table 6 reports the results from repeating the estimation of regression 1, but conditioning on whether establishments are small or large. Small establishments are defined as those with a number of employees below the median in the preceding year, while large establishments are those above the seventy-fifth percentile in number of employees, also in the preceding year.

In all our economies but Venezuela, small firms are substantially less flexible than large establishments. In Brazil, the former close about 67 percent of their gap within a year, while the latter close

22. According to the Heckman and Pagés (2000) index, the most flexible countries in our sample are Brazil and Mexico—not Chile and Colombia, as suggested by our index.

Table 6. Average Flexibility Estimates by Plant Size[a]

Explanatory variable and plant size	*Country* Brazil	Chile	Colombia	Mexico	Venezuela
Gap					
Small plants	0.670	0.685	0.675	0.561	0.529
	(0.006)	(0.007)	(0.007)	(0.006)	(0.020)
Large plants	0.808	0.783	0.790	0.607	0.529
	(0.009)	(0.010)	(0.010)	(0.007)	(0.026)
Summary statistic					
R^2					
Small plants	0.47	0.49	0.52	0.44	0.35
Large plants	0.57	0.54	0.56	0.53	0.39
No. observations					
Small plants	12,560	10,404	10,087	13,784	1,469
Large plants	6,340	5,265	5,131	7,008	741
Period	1998–2000	1988–99	1995–99	1995–2000	1997–99

a. Small plants are below the fiftieth percentile of the lagged employment distribution; large plants are above the seventy-fifth percentile of the lagged employment distribution. All regressions include year dummies. All estimates in this table are significant at the 1 percent level. Observations corresponding to extreme values (0.5 percent in right tail and 0.5 percent in left tail) of regressor excluded. Robust standard errors are in parentheses.

about 81 percent. In Chile, the figures are 69 and 78, respectively; in Colombia, 68 and 79; in Mexico, 56 and 61; and in Venezuela, 53 percent for both.

The behavior of large establishments primarily explains the substantial differences in flexibility across some of these economies. Again, this need not come from differences in labor market regulation, in which case it would not be captured by indices based on this variable. It could also reflect, for example, barriers to entry or social objectives assigned to large firms by the government.

Table 7 further splits the observations by the magnitude of the employment gap. Small gaps are defined as gaps of less than 25 percent in absolute value, while large gaps are above 25 percent. That is, we reestimate equation (1) for each country-size/size-of-gap combination (*jsg*):

$$\Delta e_{ijsgt} = d_{jsgt} + \lambda_{jsg} Gap_{ijsgt} + \varepsilon_{ijsgt} \quad . \tag{13}$$

Table 7. Average Flexibility Estimates by Plant Size and Gap Size[a]

Explanatory variable, plant size, and gap size	*Country*				
	Brazil	*Chile*	*Colombia*	*Mexico*	*Venezuela*
Gap					
Small plants, small gap	0.473	0.499	0.440	0.330	0.275
	(0.010)	(0.009)	(0.010)	(0.009)	(0.033)
Small plants, large gap	0.722	0.790	0.752	0.626	0.570
	(0.013)	(0.016)	(0.012)	(0.010)	(0.031)
Large plants, small gap	0.541	0.513	0.551	0.418	0.222
	(0.011)	(0.013)	(0.014)	(0.010)	(0.044)
Large plants, large gap	0.870	0.927	0.890	0.682	0.540
	(0.018)	(0.023)	(0.020)	(0.015)	(0.040)
Summary statistic					
R^2					
Small plants, small gap	0.21	0.27	0.22	0.14	0.08
Small plants, large gap	0.56	0.65	0.65	0.57	0.41
Large plants, small gap	0.28	0.29	0.29	0.26	0.06
Large plants, large gap	0.64	0.68	0.65	0.68	0.40
No. observations					
Small plants, small gap	9,204	8,844	7,493	9,812	886
Small plants, large gap	3,356	1,560	2,594	3,972	583
Large plants, small gap	4,903	4,342	4,052	5,729	441
Large plants, large gap	1,437	923	1,079	1,279	300
Period	1998–2000	1988–99	1995–99	1995–2000	1997–99

a. Small plants are below the fiftieth percentile of the lagged employment distribution; large plants are above the seventy-fifth percentile of the lagged employment distribution. A small gap has an absolute value less than or equal to 25 percent; a large gap has an absolute value larger than 25 percent. All regressions include year dummies. All estimates in this table are significant at the 1 percent level. Observations corresponding to extreme values of regressors (0.5 percent in right tail and 0.5 percent in left tail) are excluded. Robust standard errors are in parentheses.

Several significant conclusions follow from this table. First, all the economies we study show evidence of an increasing hazard.[23] In other words, establishments are substantially more flexible with respect to large gaps than to small ones. This hints at the presence of significant fixed costs (increasing returns) in the adjustment technology. These fixed costs may have a technological origin, as when there are strong complementarities in production or fixed proportion with sunk capital, or an institutional origin, as when dismissals require approval by a government agency or are likely to be litigated in court.

23. See Caballero and Engel (1993) for a description of increasing hazard models and their aggregate implications.

Second, the increasing hazard is particularly pronounced in large establishments in the relatively more flexible economies. This does not mean that these firms face larger fixed costs than the same establishments in less flexible economies; quite the opposite is the case, since they adjust more frequently than their counterparts in inflexible economies. Rather, it means that the benefits of adjustment overcome fixed costs sooner in large establishments in flexible economies vis-à-vis inflexible economies and that the adjustment decisions of large establishments in inflexible economies are less predictable (that is, they are not correlated with the size of the gap).

Finally, most of the additional flexibility experienced by large establishments in the more flexible Latin American economies is due to their rapid adjustment when gaps get to be very large (over 25 percent). For example, both small and large establishments have an adjustment coefficient of approximately 0.50 for gaps below 25 percent in Chile. For large deviations, on the other hand, small establishments have a coefficient of 0.79, while large establishments have one of 0.93. The patterns are similar in Brazil and Colombia and less pronounced in Mexico and Venezuela.

In conclusion, the Latin American economies present evidence of microeconomic inflexibility, and in some cases, such as Mexico and Venezuela, the problem is quite severe. Studies based only on quantifying job flows are unable to detect either of these facts: gross job flows are comparable in magnitude to those in the United States and across all the economies we study, or they yield the wrong ranking (for example, Chile would be the second-most inflexible of these economies, according to the excess reallocation numbers presented in table 3); the same remark applies to studies based solely on studying labor market regulation.[24]

We also find that allowing for an increasing hazard is important: there is clear evidence of increasing hazards, especially for large establishments in the more flexible economies. To a substantial extent, more inflexible economies seem to be those in which large imbalances go uncorrected for sustained periods of time. Conversely, large establishments in the more flexible economies seldom tolerate (or can afford to tolerate) large microeconomic imbalances.

24. Our remarks refer only to the measurement of a broad concept of microeconomic flexibility, not to the general merit of such studies.

Table 8. Flexibility Estimates Based on Equation (2)[a]

	Country			
Explanatory variable	*Brazil*	*Chile*	*Mexico*	*Venezuela*
Gap	0.855	0.675	0.592	0.401
	(0.048)	(0.034)	(0.037)	(0.184)
No. observations	8,322	17,631	18,368	968
Period	1998–2000	1988–99	1995-2000	1997–99

a. The dependent variable is the change in the gap after adjustments. Second and third lags are used as instruments. All estimates in this table are significant at the 1 percent level, with the exception of Venezuela, which is significant at the 5 percent level. All estimates are based on one regression per country, using all available observations. Colombia was not included because we did not have access to the data. All regressions that consider more than one year (Chile and Mexico) use year dummies. Observations corresponding to extreme values of regressors (0.5 percent in right tail and 0.5 percent in left tail) are excluded. Robust standard errors are in parentheses.

Finally, as mentioned in section 1, our model can also be estimated with standard dynamic panel methods. Table 8 shows that our basic conclusions remain unchanged when we use this procedure, but the precision of the estimates falls significantly, as expected.

4. The Evolution of Flexibility

Has microeconomic flexibility improved over time? Unfortunately, we only have a long time dimension for the case of Chile. In this section, we focus our analysis on this case, and we conclude that the answer to this question is negative. Flexibility has declined significantly since the Asian crisis.

All our results in this section are obtained from running variants of the following regression:

$$\begin{aligned}\Delta e_{ijt} = {} & \left[\lambda_{0jt} + \lambda_{1j}\left\{\left|Gap_{ijt}\right| > 0.25\right\} + \lambda_{2j}\left\{Gap_{ijt} < -0.05\right\}\right] Gap_{ijt} + \\ & d_{1j}\left\{\left|Gap_{ijt}\right| > 0.25\right\} + d_{2j}\left\{Gap_{ijt} < -0.05\right\} + \varepsilon_{ijsgt}\end{aligned} \tag{14}$$

where we include, but do not report, constants, time dummies, and group dummies (for example, $\left|Gap_{ijt}\right| > 0.25$). The results of these variants are reported in table 9.

Table 9. Evolution of Flexibility by Plant Size: Chile, 1987–99[a]

	Constant hazard			*Increasing hazard*			*Increasing and asymmetric hazard*		
Explanatory variable	*All plants (1)*	*Small plants (2)*	*Large plants (3)*	*All plants (4)*	*Small plants (5)*	*Large plants (6)*	*All plants (7)*	*Small plants (8)*	*Large plants (9)*
Gap 1987	0.745	0.742	0.782	0.490	0.514	0.537	0.343	0.384	0.365
	(0.030)	(0.036)	(0.068)	(0.030)	(0.038)	(0.064)	(0.030)	(0.039)	(0.063)
Gap 1988	0.674	0.707	0.716	0.424	0.481	0.445	0.272	0.344	0.270
	(0.031)	(0.041)	(0.059)	(0.031)	(0.040)	(0.058)	(0.031)	(0.040)	(0.060)
Gap 1989	0.776	0.714	0.854	0.533	0.504	0.564	0.381	0.377	0.381
	(0.038)	(0.042)	(0.054)	(0.034)	(0.043)	(0.054)	(0.035)	(0.043)	(0.055)
Gap 1990	0.677	0.656	0.765	0.441	0.478	0.488	0.274	0.326	0.289
	(0.031)	(0.039)	(0.072)	(0.030)	(0.039)	(0.068)	(0.032)	(0.041)	(0.072)
Gap 1991	0.731	0.688	0.806	0.501	0.503	0.578	0.335	0.362	0.374
	(0.033)	(0.053)	(0.058)	(0.032)	(0.050)	(0.055)	(0.034)	(0.051)	(0.058)
Gap 1992	0.740	0.705	0.758	0.520	0.522	0.503	0.359	0.380	0.302
	(0.039)	(0.063)	(0.065)	(0.036)	(0.058)	(0.063)	(0.038)	(0.062)	(0.064)
Gap 1993	0.706	0.640	0.812	0.492	0.474	0.547	0.322	0.327	0.347
	(0.034)	(0.047)	(0.066)	0.032)	(0.046)	(0.060)	(0.033)	(0.047)	(0.065)
Gap 1994	0.730	0.656	0.913	0.515	0.487	0.639	0.345	0.339	0.443
	(0.036)	(0.050)	(0.071)	(0.035)	(0.049)	(0.066)	(0.036)	(0.050)	(0.070)
Gap 1995	0.775	0.743	0.907	0.547	0.569	0.641	0.370	0.415	0.434
	(0.034)	(0.048)	(0.072)	(0.032)	(0.044)	(0.065)	(0.033)	(0.046)	(0.069)
Gap 1996	0.808	0.706	0.856	0.577	0.531	0.582	0.402	0.378	0.386
	(0.035)	(0.055)	(0.059)	(0.034)	(0.054)	(0.056)	(0.035)	(0.055)	(0.059)
Gap 1997	0.686	0.648	0.667	0.469	0.495	0.395	0.301	0.346	0.206
	(0.033)	(0.043)	(0.073)	(0.032)	(0.042)	(0.072)	(0.034)	(0.046)	(0.074)
Gap 1998	0.669	0.614	0.667	0.425	0.446	0.377	0.242	0.285	0.168
	(0.040)	(0.051)	(0.095)	(0.038)	(0.051)	(0.091)	(0.040)	(0.052)	(0.092)
Gap 1999	0.705	0.655	0.712	0.418	0.455	0.367	0.250	0.309	0.172
	(0.034)	(0.045)	(0.076)	(0.035)	(0.048)	(0.075)	(0.038)	(0.050)	(0.080)

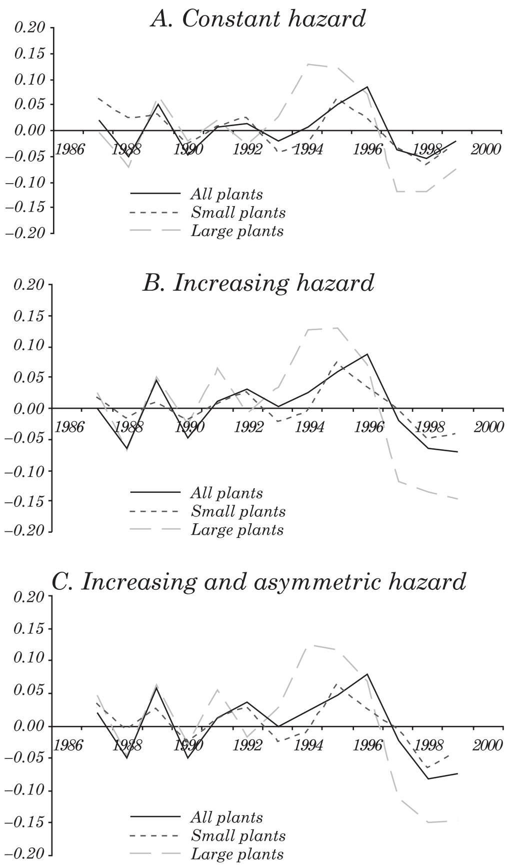
Table 9. (continued)

	Constant hazard			Increasing hazard			Increasing and asymmetric hazard		
Explanatory variable	*All plants (1)*	*Small plants (2)*	*Large plants (3)*	*All plants (4)*	*Small plants (5)*	*Large plants (6)*	*All plants (7)*	*Small plants (8)*	*Large plants (9)*
Gap (|Gap| > 0.25)				0.371	0.295	0.407	0.479	0.410	0.508
				(0.016)	(0.023)	(0.031)	(0.016)	(0.023)	(0.032)
Gap (Gap < –0.05)							–0.095	–0.172	–0.012
							(0.031)	(0.420)	(0.062)
|Gap| > 0.25				0.002	0.027	–0.023	0.004	0.019	–0.012
				(0.004)	(0.006)	(0.009)	(0.005)	(0.007)	(0.010)
Gap < –0.05							–0.093	–0.097	–0.087
							(0.003)	(0.004)	(0.007)
R^2	0.50	0.49	0.54	0.53	0.51	0.57	0.55	0.54	0.59

a. Small plants are below the fiftieth percentile of the lagged employment distribution; large plants are above the seventy-fifth percentile of the lagged employment distribution. All regressions include year and group dummies. Observations corresponding to extreme values of regressors (0.5 percent in right tail and 0.5 percent in left tail) are excluded. Robust standard errors are in parentheses.

Figure 1 plots the path of the estimated values of l_{0jt}, with their mean subtracted. A high value represents an upward shift in the adjustment hazard. We focus on the shift in the hazard itself as an index of flexibility, rather than on the average speed of adjustment, because in the realistic context of increasing hazard, the latter depends on the endogenous path of the cross section. When the hazard is constant, its shift also represents an equal shift in the speed of adjustment. When the hazard is increasing, on the other hand, the mapping from a vertical shift in the hazard to a change in the average speed of adjustment

Figure 1. Time-varying Adjustment Hazards



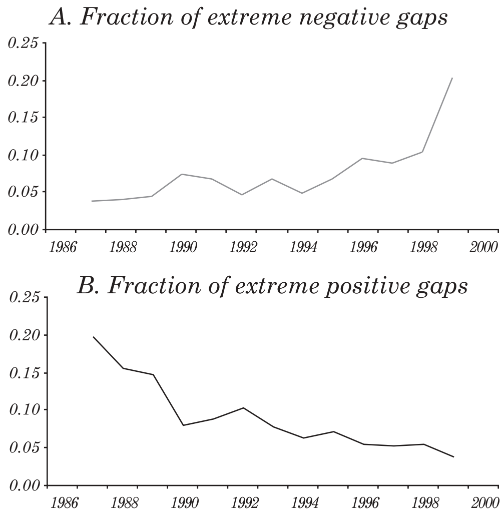
Source: Authors' calculations.

is not one for one, since the interactions with the cross-sectional distribution of gaps complicates the mapping.

The first column in table 9 and the continuous line in the upper panel of figure 1 show the results for the constant-hazard case. Under this assumption, the index of flexibility exhibits fluctuations in the second half of the 1980s and early 1990s, settles at a fairly high value in the mid-1990s, and then declines sharply during the 1997–99 period. This coefficient fell from an average adjustment coefficient of 0.77 for the three years prior to the Asian/Russian crisis episode to 0.69 in the aftermath of the crisis.

In this case, the decline in flexibility appears to be subsiding toward the end of the sample. However, columns 4 and 7 in table 9 (corresponding to all plants under increasing hazard) and the continuous lines in the middle and lower panels of figure 1 show that this finding is lost, with no evidence of recovery, once the hazard is allowed to be nonlinear. The reason for the misleading conclusion with a constant hazard is that the share of establishments with large negative gaps rises sharply toward the end of the sample (see figure 2), and establishments naturally react more strongly to this situation

Figure 2. Fraction of Extreme Gaps



Source: Authors' calculations.

under increasing hazards.[25] That is, the average speed of adjustment rises—even if the hazard does not change—when a large number of establishments accumulate substantial negative gaps.

While it is too early to tell whether this decline in microeconomic flexibility is purely cyclical or whether it reflects some structural change, we can make a few interesting observations. First, the remaining columns in table 9 and series in figure 1 show that much of the decline in flexibility is due to a decline in the flexibility of large establishments (as measured by their lagged employment).

Second, table 10 shows that while the speed of response to negative gaps remained fairly constant, the speed at which establishments adjust to shortages of labor slowed dramatically.[26] This reluctance to hire may reflect pessimism with respect to future conditions not captured in the current gap. This is unlikely to be the only factor, however, since we do not observe the rise in the speed of firing that should accompany it. The increasing nature of the adjustment hazard partly explains the asymmetry seen in the decline of the speed of

Table 10. Evolution of Flexibility and Asymmetric Hazards[a]

	Gap		*Gap < –0.05*		
Year	*Coefficient*	*Standard error*	*Coefficient*	*Standard error*	*No. observations*
1987	0.689	0.030	0.227	0.062	1,300
1988	0.720	0.030	–0.079	0.058	1,216
1989	0.729	0.033	0.155	0.061	1,248
1990	0.702	0.036	0.016	0.060	1,155
1991	0.815	0.036	–0.097	0.061	1,153
1992	0.752	0.035	0.061	0.067	1,151
1993	0.721	0.037	0.034	0.064	1,124
1994	0.831	0.039	–0.135	0.066	1,073
1995	0.891	0.036	–0.152	0.060	1,134
1996	0.859	0.039	–0.040	0.063	1,139
1997	0.710	0.039	0.028	0.062	1,146
1998	0.734	0.046	–0.078	0.069	1,144
1999	0.698	0.052	0.031	0.070	1,252
Simple average	0.758		–0.002		

a. Estimates for parameters in equation (14) in the main text. All regressions include year and gap size dummies. Observations corresponding to extreme values of regressors during the whole sample (0.5 percent in right tail and 0.5 percent in left tail) are excluded.

25. Large negative gaps are defined as being smaller than –0.25, and large positive gaps are gaps larger than 0.25.

26. Between 1994–96 and 1997–99, the latter fell from 0.86 to 0.71, while the former fell from 0.75 to 0.71.

adjustment with respect to positive and negative gaps. As we mentioned above, a substantial number of establishments developed large negative gaps (excess labor) during the slowdown; the increasing hazard implies that their adjustment did not slow as much as the decline in the average speed of adjustment.

Table 11 illustrates that the sharpest decline in flexibility came from establishments in sectors that normally experience a low degree of restructuring, either because they experience smaller shocks or because they are characterized by more technological and institutional inflexibility. Normal restructuring for sectors with a high or low degree of restructuring is measured by the excess reallocation above or below the median in Chile prior to 1997.[27] If inflexibility rather than shocks explains the ranking, then the cost of the increase in flexibility can potentially be very large, as inflexible establishments spend a significant amount of time away from their frictionless optimum.

Table 11. Evolution of Flexibility and Ex Ante Restructuring[a]

	Type of sector					
	High restructuring			*Low restructuring*		
Year	*Coefficient*	*Standard error*	*No. observations*	*Coefficient*	*Standard error*	*No. observations*
1987	0.745	0.024	902	0.749	0.030	709
1988	0.750	0.023	898	0.552	0.029	712
1989	0.824	0.023	904	0.698	0.031	705
1990	0.704	0.025	911	0.640	0.026	706
1991	0.722	0.023	902	0.748	0.030	710
1992	0.722	0.025	908	0.768	0.031	709
1993	0.786	0.024	909	0.575	0.027	713
1994	0.767	0.025	913	0.689	0.029	711
1995	0.765	0.023	904	0.788	0.030	717
1996	0.824	0.024	906	0.788	0.029	705
1997	0.722	0.026	912	0.634	0.027	702
1998	0.723	0.026	911	0.580	0.029	705
1999	0.733	0.027	895	0.664	0.029	700
Simple average	0.753			0.682		

a. High- and low-restructuring sectors are defined using the median sector excess reallocation in Chile prior to 1997. All regressions include year and gap size dummies. Observations corresponding to extreme values of regressors during the whole sample (0.5 percent in right tail and 0.5 percent in left tail) are excluded.

27. Similar results are obtained when sectors are classified according to the excess reallocation in the corresponding U.S. sectors (which serves as an instrumental variable for technological factors).

In conclusion, we clearly identified a decline in microeconomic flexibility toward the end of the 1990s, although we cannot pinpoint a specific reason for the decline. We also found that the increasing nature of the hazard is important for showing that the recovery in average flexibility around 1999 does not seem to correspond to a real increase in flexibility. Instead, it simply reflects the interaction between an increasing hazard and a depressed phase of the business cycle. Flexibility declined in 1997 and remained down until the end of our sample, particularly in the case of large establishments. We also found that the decline in flexibility is more pronounced in sectors that normally have a low degree of restructuring. If the latter is a consequence of large technological or institutional adjustment costs, then their relative slowdown is worrisome since the cost of further reducing their restructuring is particularly large.

5. Gauging the Costs of Microeconomic Inflexibility

By impairing worker movements from less productive to more productive units, microeconomic inflexibility reduces aggregate output and slows economic growth. In this section, we develop a simple framework to quantify this effect. Any such exercise requires strong assumptions, and our approach is no exception. Nonetheless, our findings suggest that the costs of microeconomic inflexibilities in Latin America are significant. The impact of the decline in microeconomic flexibility in Chile following the Asian crisis accounts for a substantial fraction of the large decline in TFP growth in Chile since 1997, which fell from an annual average of 3.1 percent for the preceding decade to about 0.3 percent after the crisis. If the decline were to persist, it could permanently shave about 0.4 percent off Chile's structural growth rate.

5.1 The Model

Consider a continuum of establishments, indexed by i, that adjust labor in response to productivity shocks, while their share of the economy's capital remains fixed over time. Their production functions exhibit constant returns to (aggregate) capital, K_t, and decreasing returns to labor:

$$Y_{it} = B_{it} K_t L_{it}^{\alpha} , \tag{15}$$

where B_{it} denotes plant-level productivity and $0 < \alpha < 1$. The values of B_{it} follow geometric random walks, which can be decomposed into the product of a common and an idiosyncratic component:

$$\Delta \log B_{it} \equiv b_{it} = v_t + v_{it}^I ,$$

where *the* v_t's are i.i.d. $N(\mu_A, \sigma_A^2)$ and the v_{it}'s are i.i.d. (across productive units, over time, and with respect to shocks) $N(0, \sigma_I^2)$. We set $\mu_A = 0$, since we are interested in the interaction between rigidities and idiosyncratic shocks, rather than in Jensen's inequality effects associated with aggregate shocks.

The price-elasticity of demand is $\eta > 0$. Aggregate labor is assumed constant and set equal to 1. We define aggregate productivity, A_t, as

$$A_t = \int B_{it} L_{it}^{\alpha} di , \tag{16}$$

so that aggregate output,

$$Y_i \equiv \int Y_{it} di ,$$

satisfies $Y_t = A_t K_t$.

Units adjust with probability λ in every period, independent of their history and of what other units do that period.[28] The parameter that captures microeconomic flexibility is λ. Higher values of λ are associated with a faster reallocation of workers in response to productivity shocks.

Standard calculations show that the growth rate of output, g_Y, satisfies: [29]

$$g_Y = sA - \delta , \tag{17}$$

where s denotes the savings rate (which we assume to be exogenous) and δ the capital depreciation rate.

When microeconomic flexibility decreases from λ_0 to λ_1, aggregate productivity decreases, reflecting slower reallocation of workers

28. More precisely, whether unit i adjusts at time t is determined by a Bernoulli random variable, ξ_{it}, with probability of success λ, where the values of ξ_{it} are independent across units and over time.

29. Here we use $g_A = 0$, since we assumed $\mu_A = 0$.

from less productive to more productive units. Equation (16) implies that the reduction in aggregate productivity is given by

$$\Delta A = \int B_{it} \Delta L_{it}^{\alpha} di \ ,$$

where ΔL_{it}^{α} denotes the difference between the value of L_{it}^{α} for the new value of λ and the value it would have had under the old λ. A tedious, but straightforward calculation relegated to appendix B shows that

$$\Delta A \cong \left[\frac{1}{\lambda_0} - \frac{1}{\lambda_1} \right] \theta A_0 \ ,$$

where

$$\theta = \frac{\alpha\gamma(2-\alpha\gamma)}{2(1-\alpha\lambda)^2} (\sigma_I^2 + \sigma_A^2) \ ,$$

and

$$\gamma = (\eta - 1)/\eta \ .$$

Using equation (17) to eliminate A_0 yields our main result:

$$\Delta g_Y \cong (g_{Y,0} + \delta) \left[\frac{1}{\lambda_0} - \frac{1}{\lambda_1} \right] \theta \ , \tag{18}$$

where $g_{Y,0}$ denotes the growth rate of output before the change in λ.

We choose parameters to apply equation (18) as follows. The markup is set at 20 percent. Parameters $g_{Y,0}$, σ_P, and σ_A are set at their average values for Chile over the 1987–96 period, namely 7.9 percent, 19 percent, and 4 percent, respectively. We also set δ equal to 6 percent. The microeconomic flexibility parameters are set at their average values for large establishments in 1994–96 and 1997–99, since these firms account for most production.[30] From this exercise we conclude that the reduction in flexibility has lowered structural output growth by 0.4 percent. This permanent cost is due to the

30. The values are 0.688 in 1994–96 and 0.892 in 1997–99; see table 9.

effect of reduced productivity on capital accumulation. One must add to this the initial direct effect of a decline in productivity on output growth, which amounts to 2.7 percent.[31] The sum of these two structural costs is very relevant. As mentioned earlier, it can account for a significant share of the decline in Chilean TFP growth from an annual average of 3.1 percent in the decade preceding the Asian crisis to 0.3 in the 1997-99 period.

Table 12 reports the potential gain in structural growth that each country in our sample could obtain from raising microeconomic flexibility to U.S. levels. Our estimates indicate that, on the low end, Chile and Colombia would have an initial gain in the range of 2 to 4 percent and a permanent increase in their structural growth rate of approximately 0.3 percent. On the high end, Venezuela would see an initial gain of 22.2 percent, although the impact on its growth rate is less pronounced because it had the lowest growth rate in our sample. Mexico could expect an initial gain of 7.4 percent and an impressive permanent increase in growth of 0.7 percent, while the corresponding percentages for Brazil are 5.0 and 0.43. These numbers are large. We are fully aware of the many caveats that such ceteris paribus comparisons carry, but the table provides an alternative metric of the potential significance of observed levels of inflexibility in the region.

Table 12. Gains from Acquiring U.S. Level of Flexibility[a]
Percent

	Country				
Indicator	*Brazil*	*Chile*	*Colombia*	*Mexico*	*Venezuela*
σ_1	27.6	19.3	25.8	24.1	38.1
g_{Y0}	2.7	6.6	2.7	3.5	2.0
Additional growth on impact	5.0	2.1	3.8	7.4	22.2
Increase in growth rate	0.43	0.27	0.33	0.70	0.18

a. The volatility of idiosyncratic shocks by country is computed using equation (10) in the text and $\phi = 0.4$. Observations corresponding to extreme values of gaps (0.5 percent in right tail and 0.5 percent in left tail) are excluded.

31. This is equal to $\frac{\Delta A}{A_0} \cong \left[\frac{1}{\lambda_0} - \frac{1}{\lambda_1}\right]\theta$.

6. Concluding Remarks

Policymakers and observers seem to have a nagging feeling that the microeconomic structure of the Latin American economies is rather inflexible and that this is a significant obstacle to growth. Not surprisingly, structural reforms aimed at heightening flexibility are extensive in most of the countries in the region. Despite this widespread belief, formal and systematic evidence on the extent of inflexibility and its costs is scarce. The data and methodological obstacles to producing this evidence are significant.

For this paper, we collected extensive data sets for several Latin American countries. We then developed a methodology suitable for extracting an answer to the inflexibility questions from these data sets.

Our estimates confirm the general fears. Microeconomic inflexibility is significant and very costly in our region. Moreover, the trend does not seem to be pointing in the right direction in Chile, the only country where we could measure the time path of flexibility with some precision. Our initial estimates suggest that if the decline in flexibility observed at the end of the 1990s were to persist, it could shave nearly half of a percent off Chile's potential growth rate.

Appendix A
Estimating the Speed of Adjustment

Our starting point is equation (1) in the main text; for simplicity we ignore sectors and time variation in the target's drift:

$$\Delta e_{it} = \psi_{it}(e^{*}_{it} - e_{i,t-1}) \ , \qquad \text{(A1)}$$

where ψ_{it} are i.i.d. with mean λ, variance $\alpha\lambda(1 - \lambda)$, and $\alpha \in [0, 1]$. We denote by z_{it} the gap after period t adjustments; that is, $z_{it} \equiv e_{it}{}^{*} - e_{it}$. We assume

$$\Delta e^{*}_{it} = \Delta e^{*}_{At} + \varepsilon_{it} \ ,$$

where Δe^{*}_{At}'s are i.i.d. with mean μ_A and variance $\sigma_A{}^2$ and where ε_{it} are i.i.d. independent from the Δe^{*}_{it}'s, with zero mean and variance $\sigma_I{}^2$.

Given an integer, M = 2, 3,… , we define

$$z^{M}_{it} = \frac{1}{M}\sum_{K=1}^{M} z_{it-k} \ . \qquad \text{(A2)}$$

The central idea is that with plant-specific fixed effects (for example, systematic differences in labor force composition), we do not observe the z's implicit on the right-hand side of equation (A1), but only observe the difference, $z_{it} - z_{it}{}^{M}$ (since the fixed effects cancel out once we subtract z^M). We therefore fix t and estimate equation (A1) with $z - z^M$ on the right-hand side instead of z. One advantage of this approach is that the estimated values of λ_t do not vary with the length of the time period considered in the sample, as is the case when fixed effects are estimated using the time average over the whole sample.

Denote $\sigma_t{}^2 \equiv \text{var}\ (z_{it})$, where the variance is calculated over i, keeping t fixed. Also denote by $\hat{\lambda}_t$ the OLS estimator of λ_t, again keeping t fixed and regressing over i. A calculation from first principles then shows that for M = 2 we have

$$E\left[\hat{\lambda}_t\right] = \lambda_t \left\{ 1 + \frac{\sigma^2_{t-1} - \sigma^2_{t-2}}{4Var\left(z_{it} - z^{\mathrm{M}}_{it} + \Delta l_{it}\right)} \right\} , \qquad \text{(A3)}$$

with

$$\sigma^2_t = \frac{1-\lambda_t}{\lambda_t\left[\alpha + (1-\alpha)\lambda_t\right]} \left\{ \left[1 - (1-\alpha)\lambda_t\right] Var(\Delta e_{it}) + \frac{\alpha(2\lambda_t - 1)}{\lambda_1}(\Delta e_{At})^2 \right\}, \qquad \text{(A4)}$$

where Δe_{At} denotes the average (over i) of Δe_{it}.

It follows from equation (A3) that the time average of the estimates for λ_t will be unbiased, since $\sigma_{t-1}{}^2$ is equal to $\sigma_{t-2}{}^2$, on average. The estimator may be biased for any particular *t*, but the expression in equation (A4) can be used to correct the bias in equation (A3), since it expresses the bias in terms of observables. We calculated the actual bias for the Chilean data, and it is rather small for all periods.

Expressions analogous to equation (A3) can be obtained for values of *M* larger than 2. Surprisingly, the result of an unbiased average described above holds only for $M = 2$.[32] An additional advantage of the $M = 2$ case is that if the fixed effect changes slowly over time, then the added precision associated with larger values of *M* comes at the expense of a larger bias stemming from time-varying fixed effects. In this sense, $M = 2$ provides a good compromise.

32. As M tends to infinity, the estimator is (asymptotically) unbiased, and there is thus no need to average over time.

APPENDIX B
Gauging the Costs

Here we show that for the model in section 5,

$$\frac{\Delta A}{A_0} \cong \left[\frac{1}{\lambda_0} - \frac{1}{\lambda_1}\right]\theta \ , \qquad \text{(B1)}$$

with

$$\theta = \frac{\alpha\gamma(2-\alpha\gamma)}{2(1-\alpha\gamma)^2}(\sigma_1^2 + \sigma_A^2) \qquad \text{(B2)}$$

and

$$\gamma = (\eta - 1)/\eta \, .$$

The intuition becomes evident on considering the following equivalent problem. The economy consists of a very large, fixed number of firms (no entry or exit). Production by firm *i* in period *t* is $Y_{it} = A_{it}L_{it}^{\alpha}$, where A_{it} denotes productivity shocks, which are assumed to follow a geometric random walk.[33] Consequently,

$$\Delta \log A_{it} \equiv \Delta a_{it} = v_t^A + \mathrm{v}_{it}^I$$

where v_t^A is i.i.d. $N(0, \sigma_A^2)$ and v_{it}^I is i.i.d. $N(0, \sigma_I^2)$. Hence, $\Delta\alpha_{it}$ follows $N(0, \sigma_T^2)$, with $\sigma_T^2 = \sigma_A^2 + \sigma_I^2$. The (inverse) demand for good *i* in period *t* is

$$P_{it} = Y_{it}^{-1/\eta} \ .$$

Finally, we assume the wage remains constant throughout.

In what follows, lower case letters denote the logarithm of upper case variables, while an asterisk on a variable denotes the frictionless counterpart of the same variable without an asterisk.

Solving the firm's maximization problem in the absence of adjustment costs leads to

$$\Delta l_{it}^* = \frac{\gamma}{1-\alpha\gamma}\Delta a_{it} \ , \qquad \text{(B3)}$$

33. We ignore hours in the production function.

and hence

$$\Delta y_{it}^{*} = \frac{\gamma}{1-\alpha\gamma}\Delta a_{it} \ . \tag{B4}$$

Denote by Y_t* aggregate production in period t if there were no frictions. It then follows from equation (B4) that

$$Y_{it}^{*} = e^{\tau\Delta a_{it}} Y_{it-1}^{*} \ , \tag{B5}$$

with $\tau \equiv 1/(1 - \alpha\gamma)$. Taking expectations (over i for a particular realization of v_t^A) on both sides of equation (B5) and noting that both terms being multiplied on the right-hand side are, by assumption, independent (random walk), yields

$$Y_t^{*} = e^{\tau v_t^A + \frac{1}{2}\tau^2\sigma_I^2} Y_{t-1}^{*} \ . \tag{B6}$$

We then average over all possible realizations of v_t^A, since these fluctuations are not the ones we are interested in for the calculation at hand. This leads to

$$Y_t^{*} = e^{\frac{1}{2}\tau^2\sigma_T^2} Y_{t-1}^{*} \ ,$$

and therefore, for $k = 1, 2, 3, \ldots$,

$$Y_t^{*} = e^{\frac{1}{2}k\tau^2\sigma_T^2} Y_{t-k}^{*} \ , \tag{B7}$$

where $Y_{t,t-k}$ represents the aggregate Y that would attain in period t if firms had the frictionless optimal levels of labor corresponding to period t - k (this is the average Y for units that last adjusted k periods ago) and where $Y_{it,t-k}$ is the corresponding level of production of firm i in t.

From the expressions derived above it follows that

$$\frac{Y_{it,t-1}}{Y_{it}^{*}} = \left(\frac{L_{it-1}^{*}}{L_{it}^{*}}\right)^{\alpha} = e^{-\alpha\gamma\tau\Delta a_{it}} \ ,$$

and therefore

$$Y_{it,t-1} = e^{\Delta a_{it}} Y_{it-1}^{*} \ .$$

Taking expectations (with respect to idiosyncratic and aggregate shocks) on both sides of this last expression (here we use the fact that Δa_{it} is independent of $Y_{t,t\text{-}1}$*) yields

$$Y_{t,t-1} = e^{\frac{1}{2}\sigma_T^2} Y_{t-1}^*$$

which, combined with equation (B7), leads to

$$Y_{t,t-1} = e^{\frac{1}{2}(1-\tau^2)\sigma_T^2} Y_t^* \; .$$

A derivation similar to the one above leads to

$$Y_{it,t-k} = e^{\Delta a_{it} + \Delta a_{it-1} + \ldots + \Delta a_{it-k+1}} Y_{t-k}^* \; ,$$

which, combined with equation (B7), gives

$$Y_{t,t-k} = e^{-k\theta} Y_t^* \; , \qquad \text{(B8)}$$

with θ defined in equation (B2).

Assuming Calvo-type adjustment with probability λ, we decompose aggregate production into the sum of the contributions of cohorts:

$$Y_t = \lambda Y_t^* + \lambda(1-\lambda) Y_{t,t-1} + \lambda(1-\lambda)^2 Y_{t,t-2} + \ldots$$

Substituting equation (B8) into the expression above yields

$$Y_t = \frac{\lambda}{1-(1-\lambda)e^{-\theta}} Y_t^* \; . \qquad \text{(B9)}$$

It follows that the production gap, defined as

$$\text{Prod.Gap} = \frac{Y_t^* - Y_t}{Y_t^*},$$

is equal to

$$\text{Prod.Gap} = \frac{(1-\lambda)(1-e^{-\theta})}{1-(1-\lambda)e^{-\theta}} \; . \qquad \text{(B10)}$$

A first-order Taylor expansion then shows that when $|\theta| << 1$,

$$\text{Prod.Gap} \cong \frac{(1-\lambda)}{\lambda}\theta \; . \qquad \text{(B11)}$$

Subtracting this gap evaluated at λ_0 from its value evaluated at λ_1, and noting that this gap difference corresponds to $\Delta A / A_0$ in the main text, yields equation (B1). This concludes the proof.

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Designing Labor Market Institutions

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There is fairly wide agreement among economists on what constitutes optimal—or at least good—product market and financial market institutions. There is much less agreement on what constitutes optimal—or at least good—labor market institutions. As a result, the public debate is too often dominated by clichés and slogans. "Get rid of labor market rigidities" is one of the most frequent. Meanwhile, policymakers focus on politically feasible, incremental reforms, with little sense of the ultimate goal.

Economists could play a more useful role here. This is why Jean Tirole and I decided to explore the optimal design of labor market institutions. This paper represents a progress report.[1] It gives a sense of the general architecture we see coming out of our analysis.

It is probably best to start with three warnings. First, our research so far has focused on unemployment insurance and employment protection, which we see as the two pillars of labor market institutions. Many other dimensions are relevant, including minimum wages, negative income taxes, labor laws, and collective bargaining. We have not analyzed these issues in detail, so I only touch on them here. Second, one size does not fit all. The economic principles we derive are quite general, but the specifics are likely to differ across countries, according to income level and institutional development. What may be optimal for Sweden may not be optimal for Chile, for example. Finally, having a sense of the ultimate goal is only half of

This paper is based on joint work with Jean Tirole.

1. An earlier report, with the same title as this paper, was given in Blanchard (2002). Blanchard and Tirole (2004) develop the theory behind the informal arguments presented below; Blanchard and Tirole (2003) apply the framework to the issue of the reform of employment protection in France.

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what is needed. The other half involves how to go from here to there, how to improve existing institutions. As governments have learned, labor market reforms face many political constraints. These constraints are very relevant, but I ignore them here. Characterizing the goal comes first, and that is the focus of this paper.

1. A Benchmark

Let me start with a stark, simplistic benchmark. Consider an economy in which workers are risk averse and firms are risk neutral. Firms hire workers and put them into jobs. All jobs look the same ex ante, that is, they have the same probability distribution. Ex post, productivity differs across jobs. If the productivity of a job is sufficiently low, the firm lays the worker off, and the worker becomes unemployed.

What happens in this economy is straightforward. Firms, which are risk neutral, insure workers, who are risk averse. They pay workers a constant wage, independent of their realized productivity. If they decide to lay some workers off, they pay them unemployment benefits so as to fully insure them against unemployment. (Note that, by offering such insurance to the risk averse workers, firms are able not only to reduce wages, but also to reduce their expected labor costs.)

The payment of unemployment benefits to the workers makes the firms fully internalize the cost of unemployment for the workers they lay off. The decision as to whether to lay off a worker is therefore socially efficient.

Now think of a different way of achieving the same outcome. Suppose that, instead of making payments directly to the workers they lay off, firms make these payments to an unemployment agency. Call these payments unemployment contributions or layoff taxes; the terminology does not matter. Also, let the unemployment agency pay unemployment benefits to the laid-off workers.

This clearly leads to the same outcome as before. By construction, payments from firms to the agency are equal to the payments from the agency to the unemployed. Firms face the same costs as before, so they make the same decisions; workers receive the same payments, so they get the same utility.

Given this equivalence, why introduce such an agency? Why not let firms handle the payments themselves? There is, in fact, a good reason. I assumed implicitly above that firms could fully insure workers by paying unemployment benefits directly. In reality, individual firms cannot easily provide unemployment insurance to laid-off workers. A one-time payment at the time of the layoff provides very poor insurance against unemployment, since the main source of uncertainty when becoming unemployed is how long one will remain unemployed. A one-time payment offers no insurance against uncertain duration.

Good unemployment insurance therefore requires payments of unemployment benefits over time, conditional on whether the worker is still unemployed and searching for another job. Individual firms are not equipped to do this. Whether a laid-off worker is still unemployed or has found another job is difficult enough for an individual firm to verify. Monitoring the unemployed worker's search effort goes far beyond what a firm can do.

Hence, there is the need for an unemployment agency to check on employment status, monitor search activity, and deliver the benefits to the unemployed. The agency need not be a state agency. The state, however, probably has to be involved, given that it already has much of the infrastructure needed to check, monitor, and distribute benefits.

If the agency pays unemployment benefits to the unemployed over time, how does one ensure that firms still make contributions to the agency equal to what the agency pays to the laid-off workers? This can be achieved in one of two ways. It can be done ex ante: at the time the layoff takes place, firms can pay the expected value of unemployment benefits that the agency will pay to the worker who is laid off. This expected value is likely to depend on the age, skills, and geographic location of the worker, and it may be difficult to assess. This suggests doing it ex post instead, while the worker is unemployed: whenever the unemployment agency sends a benefit check to an unemployed worker, these benefits are charged to the firm that laid that worker off.

1.1 Taking Stock

The purpose of the benchmark was to convey a basic message. The architecture of labor market institutions must be built on two pillars: unemployment insurance and employment protection, in the form of layoff taxes. To the extent that workers receive unemployment benefits, it is essential, for efficiency purposes, that firms take this cost into account when deciding whether to layoff workers. This requires the use of layoff taxes, a form of employment protection.

This argument is straightforward, but it is at odds with the often-heard position that the less employment protection the better. Absent layoff taxes, in our benchmark, firms would lay off more workers than is socially efficient, in effect free-riding on the unemployment benefits paid by the unemployment agency.

This basic message is general and important. The benchmark is too simple, however, in that it rules out a number of relevant imperfections in the labor market. I now consider a number of these imperfections, with an eye to refining and modifying the basic message.

2. Four Complications

The benchmark made at least four implicit assumptions. First, the unemployed can be fully insured. Second, firms can pay the layoff taxes. Third, because workers are risk averse, they are willing to accept lower wages in exchange for insurance, leading firms to offer this insurance either directly (through direct payments) or indirectly (through the unemployment agency). Fourth, all firms and all workers are the same ex ante. All four assumptions are too strong. Let me take each one in turn.

2.1 Limits to Insurance

Even if it were feasible to fully insure the unemployed, it would not be desirable to do so. The reason is well understood: if unemployment implied no loss of utility, there would be no incentives for the unemployed to search for jobs. The question is how such limits to insurance affect my earlier conclusions. A formal analysis yields two conclusions. First, laid-off workers should receive the highest feasible level of unemployment insurance consistent with search incentives. This may sound obvious, but it has practical implications for the design of unemployment insurance to which I return below. Second, layoff taxes paid by firms should exceed the unemployment benefits paid to laid-off workers. In other words, employment protection should be higher than in the benchmark. Why? Given the utility loss in becoming unemployed, it is optimal to distort the layoff decision of firms so as to decrease layoffs and, hence, the incidence of unemployment.

In short, if there are limits on unemployment insurance, it is then optimal to have higher employment protection. This inverse relation between unemployment insurance and employment protection fits the facts surprisingly well. Boeri, Conde-Ruiz, and Galasso (2003) document a clear negative relation between the generosity of unemployment insurance and the strictness of employment protection across European countries. A likely explanation is a political economy story that parallels the optimality argument above: the less generous the unemployment insurance (for whatever reason), the stronger the political pressure to put in place restrictions on layoffs and the stronger the degree of employment protection.

High employment protection is a partial substitute for unemployment insurance. It is a very poor substitute, however, because it entails strong distortions and a potentially large efficiency loss. It impedes reallocation, decreasing output and perhaps even affecting growth. This has an important practical implication. Any reform of the unemployment insurance system that delivers better insurance while maintaining search incentives is useful not only on its own, but also indirectly: it allows for a decrease in layoff taxes and thus reduces distortions.

A number of recent reforms of unemployment insurance systems offer increased benefits in exchange for stronger penalties for unemployed workers who either do not search or do not accept job offers. These efforts are promising. They relax the limits on insurance and offer the hope of reducing employment protection to a more efficient level.

2.2 Limits to Layoff Taxes

The benchmark assumed that firms were risk neutral and able to pay the layoff taxes. This assumption is too strong. Many small firms have a single owner, who is likely to be risk averse and unable to diversify the firm's risk. Even larger firms may be facing financial constraints. Layoffs, by their very nature, tend to take place when firms are not doing well. The firm may thus be unable to pay the layoff taxes. Even if the firm can pay, this may come at a high cost, perhaps forcing the firm to close other operations or preventing investment crucial to its future.

One of the things the state can do, instead of forcing firms to pay layoff taxes at the time layoffs take place, is to shift payments to times when the firm is in better financial shape. This is the principle behind the system in place in the United States. The details vary from state to state, but essentially the unemployment agency keeps a running balance for each firm, registering the benefits that the agency pays to workers laid-off by the firm on the debit side and the firm's payments to the agency on the credit side. At regular intervals, the firm pays a proportion of the remaining balance. The lower this proportion, the longer the implied average time between the payments of benefits to the workers and the payment of contributions by the firm.

Such a system may alleviate the problem, but it is unlikely to eliminate it. Some firms may still not be in a strong financial position when the tax comes due. It may therefore be optimal for the state to impose lower layoff taxes than in the benchmark, thus decreasing the burden on firms in difficulty. If financial constraints vary systematically across types of firms, then it is better to tailor the tax rate for different categories of firms than to decrease the layoff tax rate for all firms. It may be optimal, for example, to levy a lower tax on new and young firms, which tend to be financially constrained, while leaving the rate higher for established firms.

In short, the presence of financial constraints may require a decrease in layoff taxes relative to the benchmark. The unemployment agency must still be financed, however. If layoff taxes are lower, the rest of the funds must be raised through higher payroll taxes. The overall architecture now has unemployment insurance on one side and layoff and payroll taxes on the other. Moreover, the decrease in layoff taxes implies that firms will lay off too many workers relative to the benchmark. Given the presence of financial constraints, however, this is the best that can be done.

2.3 Ex Post Wage Bargaining

The benchmark assumed that because workers were risk averse, the provision of insurance by firms (either directly or indirectly through the unemployment agency) allowed the firms to decrease wages and expected labor costs. Indeed, in the benchmark, the state did not have to force firms to join the unemployment-insurance-cum-layoff-tax system; they did so voluntarily.

This assumption raises an old issue in labor economics: namely, how wages are set. Before they are hired, risk averse workers will be willing to accept lower wages in exchange for the provision of unemployment insurance. After they are hired, however, they may want to renegotiate wages, and by then the bargaining conditions are very different. If bargaining fails and the workers are laid off, they are now entitled to unemployment benefits and the firm has to pay a layoff tax. Both factors clearly strengthen the workers' bargaining position, so the wage may well go up.[2]

Hence, if wages are at least partly determined ex post, the provision of unemployment insurance is likely to lead to higher, not lower, wages. The same applies to layoff taxes: the higher the layoff taxes, the more expensive it is for the firm to lay off workers, which weakens the firm's bargaining position and thus raises the wage.

The precise characterization of the optimal architecture in this case depends on the details of bargaining and the characterization of the rest of the economy. Based on an analysis of some simple cases, two conclusions appear to hold quite generally. To the extent that higher unemployment benefits increase wages and therefore increase firms' costs, these benefits should be lower than in the benchmark. And to the extent that higher layoff taxes strengthen the bargaining position of workers, further increasing wages and firms' costs, layoff taxes should be lower than unemployment benefits, with the difference financed by payroll taxes.

2.4 Heterogeneity

The benchmark ignored ex ante heterogeneity of firms and workers. All jobs and all workers looked the same ex ante. This is obviously not the case. For example, some firms operate in volatile markets and so are likely to have a higher layoff rate than firms in more stable sectors. Some workers, because of individual characteristics such as lack of work experience, represent a greater risk to the firm than do others, and they are more likely to be laid off. This heterogeneity has implications for the design of tax rates.

I focus here on worker heterogeneity; parallel arguments can be extended to the case of firm heterogeneity. When firms have to pay high layoff taxes, they are reluctant to hire workers whom they may have to lay off. Depending on how wages are set, these high-risk workers may have to accept lower wages in order to be hired or may simply not be hired at all.

2. This effect is well captured by the assumption of Nash bargaining in modern flow-bargaining models (as presented, for example, in Pissarides, 2000), but it is clearly more general than this particular class of models.

Decreasing layoff taxes below unemployment benefits can help alleviate the bias against these high-risk workers. Again, specific groups of workers should be targeted. For example, the tax structure may provide preferential treatment for new entrants—workers without a work history. If firms that lay off new entrants are subject to lower layoff taxes, this reduces their incentives to discriminate against new entrants in hiring. In addition, contracts might include a trial period during which either workers or firms can separate at no cost, allowing both parties to assess the quality of the match before layoff taxes enter into force.

2.5 Taking Stock

The benchmark established a simple architecture, with full unemployment insurance on one side, employment protection on the other side, and layoff taxes equal to unemployment benefits. A closer look suggests a number of amendments. Limits to insurance point to increasing layoff taxes relative to the benchmark. Financial constraints suggest instead decreasing layoff taxes relative to the benchmark. Ex post bargaining suggests decreasing both unemployment benefits and layoff taxes. Heterogeneity suggests treating different categories of firms or groups of workers differently—for example, by applying lower layoff taxes to young firms and new workforce entrants or by introducing a trial period when neither unemployment benefits nor layoff taxes apply.

All these amendments take the form of changes in the level of unemployment benefits, or in the level and composition of taxes. One issue I have not discussed is the role of judges in the process. This is an important issue, since employment protection has an important judicial component in most countries. To think about it, it is important to introduce the distinction between layoffs and quits. Not all separations are layoffs; many are quits, triggered not by a change in the productivity of the job, but by the offer of another job to the worker or by increased worker dissatisfaction with the current job. To the extent that firms only pay layoff taxes and workers only receive unemployment benefits in the case of a layoff, this opens the scope for games between firms and workers. Firms that want to layoff a worker may harass the worker into quitting in order to save on the layoff tax. Workers who want to quit may misbehave so as to be laid off, thus getting unemployment benefits. If layoff taxes are less than unemployment benefits, this opens the possibility of another set of games, this time with workers and firms on one side and the state on the other. Workers and firms may collude and declare quits to be layoffs, getting a net subsidy from the state.

In all these cases, the incentives to misbehave depend on the generosity of benefits and the level of layoff taxes. Judges must clearly be involved in cases of disagreement between firms and workers. If, however, a firm is willing to declare a separation a layoff and pay the layoff tax, there is no obvious reason for judges to become involved and potentially overturn the decision of the firm or require additional payments. I insist on this last point because in many countries, judges can and do second guess firms' decisions to layoff workers, thereby introducing substantial uncertainty and arbitrariness in the process. Layoff taxes are a much better instrument for forcing firms to face the implications of their layoff decisions.

Let me wrap up this section. The complications I have explored may lead some readers to reject the whole architecture—to give up on state-provided insurance so as not to have to confront the issues of financing. I return to the issue of self-insurance below, but I am quite sure this conclusion is wrong. Optimal tax and insurance systems are, by their nature, complicated. This is no reason to reject them in toto, just as the complexity of the tax system does not justify eliminating government spending. The goal must be to provide insurance at the smallest cost in terms of efficiency. The message from this and the previous section is that the basic architecture needed to do so is a simple combination of unemployment insurance and employment protection. The details are complex and must be carefully worked out, but this should not obscure the basic architecture.

3. Two Issues of Relevance to Latin America

This section takes up two issues that appear particularly relevant in the context of labor market reforms in Latin America. The first is the role of severance payments, while the second is the role of self-insurance by workers and of mandatory unemployment accounts.

3.1 Severance Payments

So far, I have described a system based on unemployment benefits combined with layoff (and possibly payroll) taxes. I have not mentioned severance payments—direct payments from firms to workers at the time of separation. The two central issues with regard to severance payments are whether they can serve as an alternative to the system I have described and whether they might play a role of complementing unemployment insurance within that system.

As for the first issue, severance payments are a very poor alternative to the system I have described. The basic reason was addressed in the discussion of the benchmark. Severance payments provide very poor insurance against the main source of uncertainty associated with unemployment, namely, unemployment duration. Some economists argue that lump-sum payments such as severance payments provide strong incentives for the unemployed to search for another job. Indeed they do, but this comes at the cost of very poor insurance. Any need to provide search incentives is better accomplished by a benefit schedule in which unemployment benefits decrease with the duration of unemployment.

The analysis presented above provides other arguments against severance payments as unemployment insurance. As we have seen, financial market imperfections may make it optimal to have lower layoff taxes, while providing unemployment insurance to the workers. This is easily done in a system in which unemployment insurance is financed partially by layoff taxes and partially by payroll taxes. It is impossible under severance payments, however, where by construction payments by firms are equal to the benefits received by workers. Also, some firms may simply go bankrupt. Under severance payments, workers bear the bankruptcy risk and may thus get nothing. In the presence of unemployment insurance, the risk is taken on by the unemployment agency, so workers can still receive unemployment benefits.

Nevertheless, severance payments should still be considered as an alternative to unemployment insurance plus layoff taxes when a country is at an early stage of institutional development. Running an unemployment agency—from keeping track of the employment status of workers to monitoring search activity and distributing benefits—is a complex operation. In countries with limited institutional capacity, severance payments may be the best that can be done, despite their shortcomings. As such countries develop, they should move from a system based on severance payments to one based on unemployment benefits and layoff taxes. One of the political challenges in such a transition is how to decrease severance payments while introducing unemployment insurance—an issue relevant for Chile. If unemployment insurance is introduced and severance payments are not reduced roughly in proportion, the outcome may prove very inefficient.

This brings me to the second issue, of whether severance payments have a place within the system of unemployment insurance and layoff taxes. The answer is probably yes. Losing a job involves two different costs. The first is the cost of being unemployed for some time, which depends on how long one is unemployed. The second is the cost of becoming unemployed, which is incurred even if another job is found right away. This is a psychic cost involving the loss of a network of workplace friends, the loss of self-esteem, and so on, and it can be substantial, especially for workers with high seniority.

Two characteristics of this psychic cost are relevant here. First, it can be assessed at the time of separation. This implies that in contrast to the first cost, it can be largely compensated by a one-time payment, that is, by severance payments at the time the layoff takes place. Second, it is likely to be a function of seniority. The longer the worker has been in the firm, the higher the psychic cost of losing a job. This suggests that the payment should be increasing, perhaps even convex, in seniority.

Thus while severance payments are an inferior way of delivering unemployment insurance, they may be justified as partial compensation for the loss associated with losing a long-held job. This points to a complementary role for limited severance payments, increasing in seniority, in addition to unemployment insurance.

3.2 The Role and Scope for Self-insurance

Given the distortions associated with any realistic system of state-provided unemployment benefits, one may ask whether it would not be better simply to rely on self-insurance by workers, so as to avoid all these problems. By self-insurance, I mean the accumulation of sufficient precautionary saving by workers to be used if and when they become unemployed. Jean and I have just started working on this set of issues, so what follows is speculative.

Self-insurance clearly alleviates some of the problems discussed above. Consider an economy in which some insurance comes from self-insurance by workers and some insurance is provided by the state. A strong reliance on self-insurance reduces some of the moral hazard problems discussed earlier: when workers self insure, they have strong incentives to search for jobs should they become unemployed. Self-insurance also reduces the gap between ex ante and ex post wage setting and thus reduces expected costs for firms.

Self-insurance is not sufficient on its own, however, and cannot provide a full substitute for state-provided insurance. Compare saving for retirement and saving for unemployment. The time of retirement is roughly known in advance; it is a long way away when one starts his or her working life, and it is thus easy to plan for. In contrast, unemployment is uncertain; it often comes early in working life (indeed often at the very start), when workers have not accumulated substantial funds. In short, while one may well want to rely on individual retirement saving, the arguments do not carry over to individual unemployment saving. Without state-provided insurance, some of the unemployed are likely to have insufficient funds to maintain an adequate level of consumption.

In practice, existing individual unemployment account systems always include some additional state-provided insurance, for example, allowing unemployed workers to borrow up to some ceiling, either directly from the state or from financial institutions through a state guarantee. These additional provisions raise many of the same issues discussed earlier. How much should the state provide or guarantee, and in what form? How does the state ensure that firms internalize the cost of these guarantees, thereby motivating them to take efficient layoff decisions? I do not yet know the answer to these questions, but the optimal architecture probably includes some self-insurance by workers, within a system of state-provided unemployment insurance and layoff taxes.

4. Some Conclusions

I end by stating a number of broad conclusions, probably with more conviction than is warranted.

4.1 Social Protection and Efficiency

Countries can provide high social protection to workers, without large sacrifices in efficiency. This requires three main tools:

— The provision of unemployment insurance through an unemployment agency. Benefits can be generous, but they must be conditional on active search and job-taking. The idea of requiring the unemployed to take "acceptable jobs" or lose benefits is appealing, and it underlies reforms in many European countries. In principle, it provides insurance contingent on the state of the labor market. If there are truly no jobs, the unemployed continue to receive benefits, as they should. If there are jobs and the unemployed do not take them, they lose benefits, as they should. In practice, however, it is difficult to define what constitutes an acceptable job and to enforce the conditional receipt of benefits.

— Employment protection, in the form of layoff taxes rather than judicial intervention.

— Reliance on a negative income tax rather than on a minimum wage to ensure that even low-productivity workers have an adequate level of income. A minimum wage should be set to avoid the worst cases of exploitation by firms, but it should be a true minimum, rather than a living wage. If the productivity of the lowest-productivity workers is less than is needed for them to live decently, the difference must be made up by the state, not through the imposition of a minimum wage.

4.2 The Sins of Europe

In light of the characterization of good labor market institutions described in this paper, many European countries committed three sins. First, they often chose open-ended unemployment benefits or assistance, and, for some categories of workers, chose very high replacement rates (defined as the ratio of after-tax benefits to after-tax wages); this creates few incentives for some of the unemployed to look for work. Second, they established heavy judicial and administrative employment protection. Nearly all European countries finance unemployment benefits through payroll rather than layoff taxes, which, by itself, would lead to excessive layoffs. This is offset, however, by high judicial and administrative employment protection. In many countries, judges can second-guess and overturn the decision of a firm to layoff workers. This should not be the case. Third, they relied too heavily on the minimum wage rather than a negative income tax.

In most of these countries, reforms are taking place at all three margins. The highest replacement rates have been reduced. New labor contracts have been introduced, subject to simpler and more limited employment protection.[3] Many countries have introduced some form of a negative income tax. There is still a long way to go, but the movement is in the right direction.

3. The existence of two types of contracts, some with limited protection and some with full protection, raises other issues, but this is a topic for another time. See, for example, Blanchard and Landier (2002).

4.3 Lessons from the Unemployment Miracles

In some countries, unemployment has either remained low (in Sweden, for example, except for a sharp cyclical upturn in the early 1990s) or declined dramatically after increasing in the 1970s and 1980s (for example, in the Netherlands). These countries have achieved low unemployment without dramatic labor market reforms. They have eliminated excesses, but continue to offer high levels of social protection, even relative to the European average. Institutional reforms probably played some role in the decrease in unemployment. Wage moderation was a major factor, however, and this is not easily explained by changes in institutions. It seems mostly to reflect the attitudes of unions in collective bargaining.[4] In countries where collective bargaining is important, good labor relations, trust between unions and firms, and some form of wage coordination both seem essential to maintaining low unemployment in the face of major adverse shocks.

4. I draw here on Blanchard and Phillipon (2003), with apologies for yet another self reference.

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The Dynamics of Earnings in Chile

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Uncertainty is a key dimension of individual decisionmaking. Individuals cannot insure against certain contingencies under incomplete markets. Uncertainty thus influences the life-cycle evolution of consumption and savings, labor supply and asset allocation, and education and occupation choices. Uncertainty and risk also determine income and consumption inequality. Individuals who are identical ex ante will have different lifetime paths of consumption ex post, as some individuals are lucky and get good draws of income, employment, and health, whereas others get bad shocks and end up with lower levels of consumption over the life cycle. Income mobility and the persistence of income inequality and poverty depend on the dynamics of earnings, health outcomes, investment opportunities, and general earnings capacity.

In this paper, we measure the earnings uncertainty faced by individuals. Most of the existing empirical literature focuses on the dynamics of income and wages using data from developed countries (Abowd and Card, 1989; Pischke, 1995; Meghir and Pistaferri, 2004). Our data are drawn from a survey of Chilean households—namely, the *Encuesta Suplementaria de Ingresos* (ESI) carried out by the National Institute of Statistics (INE) (see INE, various years). Whether consumers in an emerging economy face levels of uncertainty similar to those in developed economies is an empirical matter that is

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addressed in this paper. The welfare consequences of uncertainty may be much larger than our findings indicate, however: individuals operating in underdeveloped markets have fewer opportunities than their counterparts in developed countries to share risks through the marketplace, and the public welfare system is much smaller in developing countries, offering little possibility of offsetting negative shocks.

Our modeling structure allows us to distinguish between a predictable and an unpredictable component of income. We further decompose the unpredictable part into permanent and transitory shocks to income. Specifically, we model the unexplained portion of individual earnings as the sum of a permanent and a (persistent) transitory disturbance. We also allow for time-varying variances of permanent and transitory disturbances, and we evaluate whether they correlate with the business cycle. Since the ESI data set is a repeated cross-section, we construct synthetic panels to perform our estimations. Our synthetic panels contain annual observations from 1990–2000, based on five-year birth cohorts.

Our results for men between the ages of twenty-five and sixty indicate that the age profile of labor income has the typical hump shape and that there are very large educational effects. At age fifty, a college-educated individual expects to earn 2.5 times the earnings of a person who finished high school and 3.8 times the earnings of an individual who only completed eight years of schooling. We also find that married men earn more than their single and divorced counterparts and that household size has a negative impact on earnings.

Our decomposition of the unexplained portion of income yields a high persistence but low variance of permanent shocks, together with a negligible variance of the transitory shock. These low variances may be an artifact of our synthetic panel technique, since averaging reduces the observed variability. We investigate this hypothesis by comparing our results to those obtained using U.S. data from the Panel Study of Income Dynamics (PSID) (see SRC, various years). When we replicate our cohort estimation procedure with U.S. data, we find a similar process for the dynamics of income, but the variance of earnings is significantly higher in the United States than in Chile. We interpret this as evidence of the relative rigidity of the Chilean labor market vis-à-vis the U.S. labor market. Another result is that averaging within cohorts reduces the estimated variance of the permanent shock by one order of magnitude. We cannot provide an estimate of the variance of the transitory shock, as all our benchmark estimates turn out to be insignificant.

If markets are complete, individuals can perfectly share their good and bad fortunes. The measurement of individual uncertainty then becomes irrelevant. However, vast evidence shows that in practice, many important events are not insured and markets do not fully pool risks (Attanasio and Davis, 1996; Dynarski and Gruber, 1996). A number of mechanisms help individuals insulate their consumption from income shocks (changes in their labor supply, spousal income, public and private transfers, and the progressivity of the income tax). In this paper we ask whether government transfers allow consumers to partly offset persistent shifts in earnings capacity. To answer this question, we reestimate our basic model using labor income plus the receipts from public welfare programs as our measure of individual earnings. We find that the inclusion of government transfers hardly affects the estimated income process, although earned income has a negative effect on the likelihood that any given individual receives a transfer.

The paper also provides a number of applications. We analyze income inequality and earnings mobility simulating the life-cycle income profiles of individuals who face the process we have estimated. Since income is estimated to be highly persistent, we should expect to observe little mobility of individuals across the distribution of income. An individual who starts off at the lowest quintile of the earnings distribution will stay in that same quintile for a year with a 0.77–0.84 chance. Furthermore, the likelihood that the same individual will still be in the lowest quintile ten years ahead ranges between 0.40 and 0.58. A similar pattern is found at the top of the distribution. That is, we find that the Chilean income distribution is highly persistent because the underlying earnings process is also highly persistent. Finally, we find that a large portion of income inequality can be explained by the underlying variability of the earnings process.

The paper is organized as follows. The next section describes the data and compares the ESI to the *Encuesta de Caracterización Socioeconómica Nacional* (or CASEN survey), which is the most widely used survey for the analysis of Chilean household behavior (see MIDEPLAN, various years). Section 2 then presents the model and estimation techniques. In section 3, we provide our estimates of mean income and then use the unexplained portion of income to fit different dynamic processes. We also compare the results on Chile to a similar sample of U.S. workers. Section 4 provides a number of applications of our results, and section 5 concludes.

1. Data

The data used in this paper are drawn from the *Encuesta Suplementaria de Ingresos* (ESI), which is a supplement to the national employment survey conducted monthly by the INE. The main goal of the ESI is to provide information on individual and household income. The ESI collects information over the last quarter of every year on a sample of roughly 36,000 households. These households are representative of the Chilean population. The survey gathers information on all household members that are at least fifteen years old. Data are registered on all types of income perceived during the previous month, as well as on a number of individual characteristics such as educational attainment, marital status, gender, and employment status. Population weights are also provided. Data are available for the 1990–2000 period, with the exception of 1994 when the survey was not conducted. The use of the ESI as a source of income data has been fairly limited; one exception is Granados (2001).

Our analysis considers men between the ages of twenty-five and sixty who are not self-employed. We deflate all nominal variables using the consumer price index (CPI) of the corresponding month of the interview. Real variables are reported in December 1999 Chilean pesos. Table 1 reports the sample's basic statistics. On average, individuals in our sample earn almost 170,000 pesos each month. The median is just above 100,000 pesos, reflecting the skewness of the Chilean income distribution. About 17 percent of individuals report income below the monthly minimum wage. The typical individual in the sample is thirty-eight years old, married, and has completed nine years of education (corresponding to an education level of one year of high school).

Table 1. Sample Descriptive Statistics: ESI

Variable	*Mean*	*Standard deviation*	*Minimum*	*Maximum*	*Median*
Monthly labor income[a]	168,534	191,520	6	4,189,475	107,729
Age	38.8	9.4	25	60	38
Years of schooling	9.3	4.2	0	20	9
Household size	4.6	2.0	1	26	4
Married	0.70	0.46	0	1	1
Geographical location (fraction)					
Metropolitan Region	0.22	0.42	0	1	0
V Region	0.11	0.31	0	1	0
VIII Region	0.13	0.34	0	1	0

Source: ESI survey data (INE, 1990–2000).
a. In December 1999 Chilean pesos.

Finally, the median household has four residents, and most individuals live in the V, VIII, and Metropolitan administrative regions.

Figure 1 plots the distribution of personal labor income. The distribution shows the extent of income inequality in Chile, which is analyzed extensively elsewhere. The figure also shows the distribution of income in the 1996 CASEN (taken from Baytelman, Cowan, and De Gregorio, 1999).[1] These two distributions are not directly comparable, as the CASEN figures include transfers and represent different sample years.[2] Furthermore, the ESI distribution is built from our data set on men. Even so, the graph shows that the distributions are quite alike, especially for the middle deciles. Most of the differences are concentrated at the bottom and top of the distribution: the ratio of the income share of the 20 percent of individuals with the highest income to the share of the bottom 20 percent is 7.9 in the ESI versus 13.8 in the CASEN, while the ratio of the share of the highest decile to the share of the lowest decile is 13.2 in the ESI and 29.5 in the CASEN.

Figure 2 shows the evolution of the mean of the log of real earnings over the sample period. On average, real earnings grew at an annual rate of 4.7 percent in the period 1990–2000.[3] The figure breaks the sample down by educational attainment: individuals identified as having a primary education have completed eight years of schooling; those with a high school education have completed twelve years of schooling; and those with a college education have completed seventeen years of education.[4] In all cases the path of mean earnings is

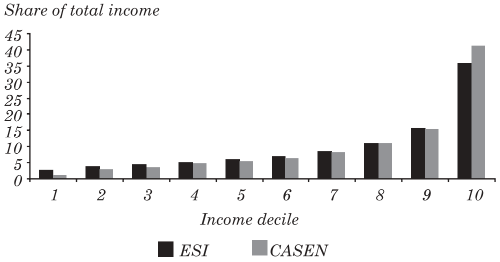
1. The CASEN is the most widely used survey for the analysis of Chilean household and individual income and earnings. The survey began in 1985, and it has been carried out almost every two years since then. The CASEN measures household and individual income for a representative sample of the Chilean population; it sampled 48,107 households in 1998. Like the ESI, the survey gathers information on all types of income and on a number of demographic characteristics. It also collects information on in-kind transfers, such as public programs in education, housing, and health, and on housing and durable goods ownership. This type of information supports detailed studies of poverty, and the survey has mainly been used to examine income inequality and the role of social policies in reducing it (Anríquez, Cowan, and De Gregorio, 1998; Larrañaga, 1994; Contreras and others, 2001). We are precluded from the use of the CASEN because of its two-year frequency. As we show below, the dynamics of earnings is highly persistent, and thus the CASEN misses most of the action in the two-year lag. The appendix provides a formal demonstration of this point.

2. Nevertheless, the distribution of income has hardly changed over the last decade. See Baytelman, Cowan, and De Gregorio (1999).

3. Per capita gross domestic product (GDP) grew at about 5 percent over this period.

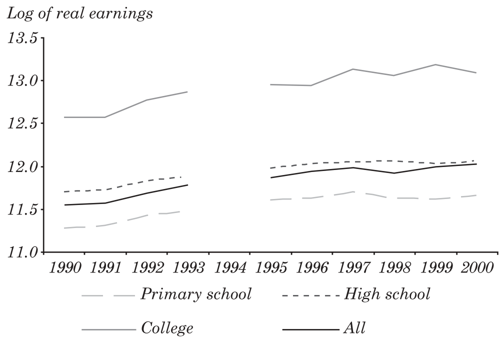
4. Universities in Chile simultaneously grant a college degree and a professional title. Most programs last about five years.

Figure 1. Distribution of Labor Income: ESI versus CASEN



Source: Authors' calculations, based on ESI and CASEN survey data (INE, 1990–2000; Baytelman, Cowan, and De Gregorio, 1999).

Figure 2. Mean of Log Real Earnings



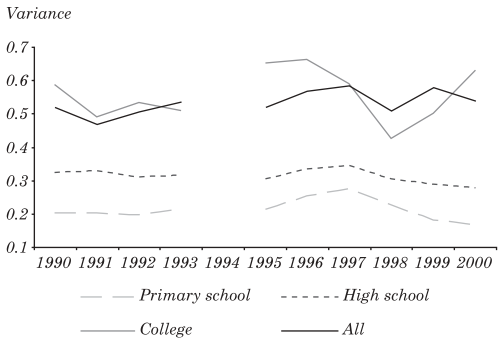
Source: Authors' calculations, based on ESI survey data (INE, 1990–2000).

upwardly trended, although the annual growth rates correlate positively with educational attainment. Chile, like other countries, has thus experienced a widening of the earnings distribution.[5]

Figure 3 plots the evolution of the variance of log earnings over the sample period, for all individuals and for three different education groups. Highly educated workers face a much larger variance of earnings; that is, there are large private returns on university education at the cost of increased earnings risk. The variance of the earnings of low education groups is quite stable over the period. This stability contrasts with the behavior of the earnings variance of those

5. For U.S. evidence, see Bound and Johnson (1992), Katz and Murphy (1992), and Murphy and Welch (1992); for Chilean evidence, see Bravo and Marinovic (1997), and Beyer and Le Foulon (2002).

Figure 3. Variance of Log Real Earnings



Source: Authors' calculations, based on ESI survey data (INE, 1990–2000).

who have attained college education, which experienced large swings over the ten-year span.

In the next subsections we use the ESI dataset to estimate mean income profiles over the life cycle for the typical Chilean individual. We then use the unexplained portion of income to estimate the dynamic process of earnings. The use of the ESI survey has a major shortcoming: the analysis of income dynamics requires following the same individuals over time. Since the ESI survey represents cross-sections of households, we build synthetic panels based on five-year birth cohorts. Table 2 presents the number of observations available for each cohort and year.

2. The Earnings Model

In this paper, we consider models in which all individuals within an educational category have identical income processes, but face different realizations of this process.[6] Income consists of the sum of a predictable component and a stochastic component. Let y_{it} represent the logarithm of the real measured income of individual *i* in year *t*. Let $\mathbf{Z}_{it}$ represent a vector of demographic characteristics, and η_{it} the stochastic component of income. We assume that the unexplained component can be decomposed into a permanent shock, y_{it}^{p} (such as health shocks that affect earnings capacity in a long-lasting way and

6. Recent literature modeling earnings processes allows for heterogeneity between agents. See Alvarez, Browning, and Ejrnæs (2002).

Table 2. Number of Available Observations by Cohort and Year

	Year										
Cohort[a]	*1990*	*1991*	*1992*	*1993*	*1995*	*1996*	*1997*	*1998*	*1999*	*2000*	*Total*
56–60	808	652	503	323	0	0	0	0	0	0	2,286
51–55	1,226	1,186	1,071	979	843	589	451	312	148	0	6,805
46–50	1,642	1,640	1,524	1,492	1,179	1,181	1,167	983	958	972	12,738
41–45	2,045	2,032	1,945	1,832	1,615	1,607	1,617	1,545	1,478	1,437	17,153
36–40	2,471	2,370	2,346	2,215	1,994	2,274	1,881	1,782	1,697	1,860	20,890
31–35	3,054	2,952	2,745	2,680	2,620	2,436	2,640	2,462	2,404	2,435	26,428
26–30	3,520	3,370	3,134	3,091	2,767	2,950	2,847	2,819	2,783	3,030	30,311
21–25	702	1,333	1,955	2,671	2,928	3,004	2,849	2,729	2,715	2,818	23,704
16–20	0	0	0	0	573	1,101	1,691	2,114	2,530	2,873	10,882
Total	15,468	15,535	15,223	15,283	14,519	15,142	15,143	14,746	14,713	15,425	151,197

Source: ESI survey data (INE, 1990–2000).
a. Age in 1990.

long-term unemployment), and a transitory innovation, μ_{it} (such as bonuses and overtime pay). We also allow for classical measurement error, ω_{it}. Finally, we assume that y^p and μ_{it} are uncorrelated at all leads and lags. We thus propose the following model for individual income:

$$y_{it} = \mathbf{Z}_{it}\beta + \eta_{it} \text{ and}$$

$$y_{it} = \mathbf{Z}_{it}\beta + y_{it}^p + \mu_{it} + \omega_{it} \ .$$

We allow for different assumptions on the process that both the permanent and transitory innovation follow. In the benchmark case, for instance, we assume that the permanent component is a random walk, whereas the transitory shock has some persistence:

$$y_{it}^p = y_{it-1}^p + \upsilon_{it} \text{ and}$$

$$\mu_{it} = \varepsilon_{it} - \theta\,\varepsilon_{it-1} \ .$$

We allow for persistence in transitory shocks to account for innovations, such as overtime pay and bonuses, that may last for a while but do not have long-lasting effects.

Alternatively, we explore a model in which permanent shocks follow a first-order autoregressive, or AR(1), process whereas the transitory component is independent and identically distributed (i.i.d.); that is,

$$y_{it}^p = \rho y_{it-1}^p + \upsilon_{it} \text{, for } 0 < \eta < 1, \text{ and}$$

$$\mu_{it} = \varepsilon_{it} \ .$$

We estimate our complete model in two stages. In the first stage, we use individual-level data to estimate η and to compute

$$\hat{\eta} = y - \mathbf{Z}\hat{\beta}$$

for each observation in our sample. In the second stage, we classify all observations on the basis of the year of birth and take averages of $\hat{\eta}$ to build a synthetic panel of cohort/year means.[7] That is,

7. We use the survey's population weights to build the means.

$$\hat{\eta}_t^c = \frac{\sum_{i \in c,t} \hat{\eta}_{it}^c}{n_t^c},$$

where the superscript c indexes birth-year cohorts and n_t^c represents the number of available observations in cohort c in year t. We use this synthetic panel to estimate the variances of the permanent and transitory components of income shocks ($\sigma_{\upsilon t}$ and $\sigma_{\varepsilon t}$, respectively) and the persistence of the transitory innovation, θ. Our modeling structure allows for time-varying variances. We estimate these parameters using equally weighted generalized method of moments (GMM) by minimizing the distance between the theoretical and the empirical autocovariances of the first difference of the stochastic component of income.[8]

Assume that there is no measurement error and that the dynamics of earnings are characterized by a random walk plus a first-order moving average, or MA(1), transitory shock.[9] Then,

$$\Delta \eta_{it} = \eta_{it} - \eta_{it-1} = \upsilon_{it} + \varepsilon_{it} - (\theta + 1)\varepsilon_{it-1} + \theta\varepsilon_{it-2} .$$

The theoretical autocovariances are thus given by

$$\text{var}(\Delta\eta_{it}) = \sigma_{\upsilon t} + \sigma_{\varepsilon t} + (\theta+1)^2 \sigma_{\varepsilon t-1} + \theta^2\sigma_{\varepsilon t-2} ,$$

$$\text{cov}(\Delta\eta_{it}, \Delta\eta_{it-1}) = -(\theta+1)\sigma_{\varepsilon t-1} - \theta\ (\theta+1)\sigma_{\varepsilon t-2} ,$$

$$\text{cov}(\Delta\eta_{it}, \Delta\eta_{it+1}) = -(\theta+1)\sigma_{\varepsilon t} - \theta\ (\theta+1)\sigma_{\varepsilon t-1} ,$$

$$\text{cov}(\Delta\eta_{it}, \Delta\eta_{it-2}) = \theta\sigma_{\varepsilon t-2} ,$$

$$\text{cov}(\Delta\eta_{it}, \Delta\eta_{it+2}) = \theta\sigma_{\varepsilon t} ,$$

$\text{cov}(\Delta\eta_{it}, \Delta\eta_{it-j}) = 0$, for $j > 2$, and

$\text{cov}(\Delta\eta_{it}, \Delta\eta_{it+j}) = 0$, for j > 2.

8. See Altonji and Segal (1996) for an analysis of alternative weighting procedures.

9. In our procedure, we assume that measurement error cancels out when we collapse our individual data set into cohort means. We thus ignore measurement error in what follows.

We follow a similar procedure to estimate the underlying parameters when we assume alternative dynamic specifications.

The fact that we construct a synthetic panel and follow cohorts, but not individuals over time, implies that our analysis is based on averages. We thus expect to underestimate the true uncertainty level that individuals face in Chile. In the analysis below, we provide estimates from a comparable sample taken from U.S. data to show how much the estimated process changes when we follow cohorts instead of individuals.

3. Results

We report our first-stage estimation results in table 3. In the regression, we control for age, education, marital status, and household size, as well as for interaction terms and nonlinear effects of these variables. We also control for the region of residence and the year and month of the interview.

Our results show that the age profile of labor income has the typical hump shape found for other countries. We also find very large educational effects. Figure 4 plots the estimated age profiles for three different educational groups identified above. All the other variables have been set at their average sample levels. The magnitude of the educational effect becomes evident when we consider three individuals who are identical except for their level of schooling. At age twenty-five, an individual who has completed eight years of schooling earns,

Table 3. Mean Income[a]
(Dependent variable: log of monthly labor income)

Explanatory variable	*Coefficient*	*Robust standard error*
Age	0.030152	0.002192
Age^2	–0.000312	0.000025
Years of schooling	–0.024810	0.004483
Years of $schooling^2$	0.003484	0.000285
Age*Years of schooling	0.000641	0.000062
Years of $schooling^4$	0.000006	0.000001
Household size	–0.009955	0.001391
Married	0.190891	0.010351
Household size*Married	–0.002730	0.001902
Constant	10.93174	0.052274
R^2	0.56	

Source: Authors' calculations, based on ESI survey data (INE, 1990–2000).
a. The dependent variable is the log of monthly labor income. The regressions include a full set of dummies for the year and month of the interview and the region of residence.

on average, about 85,000 pesos per month, whereas an individual with twelve years of education earns almost 120,000 pesos per month—that is, a difference of 40 percent. A college-educated individual earns, on average, about 280,000 pesos at age twenty-five, or 2.3 times the earnings of a high-school-educated individual. These differences increase with age. At age fifty, a college-educated individual earns 2.5 times the earnings of a person who finished high school and 3.8 times the earnings of an individual who only completed eight years of schooling. These differences widen further when we take into account the findings that education and household size are negatively related and that household size has a negative impact on earnings. Educated people are less likely to be married than individuals with incomplete schooling, but this correlation is quite small in the sample.

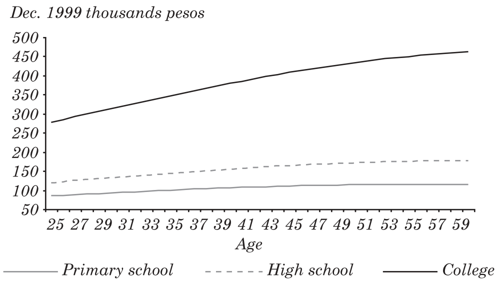
3.1 The Dynamics of Income

Since we do not follow the same individuals over time, we estimate the income process using a synthetic panel approach. For each individual in the sample, we take the unexplained component of the log of income as

$$\hat{\eta} = y - \mathbf{Z}\,\hat{\beta}\ .$$

We then classify all observations according to birth cohort to form our synthetic panel. Figure 5 tracks the variance of the unexplained portion of earnings within each cohort observed from 1990 through 2000. The variance clearly increases with age, which reflects the fact that individuals who are identical ex ante may follow quite different paths of income. In other words, in a sample of ex ante identical

Figure 4. Mean Monthly Labor Income



Source: Authors' calculations, based on ESI survey data (INE, 1990–2000).

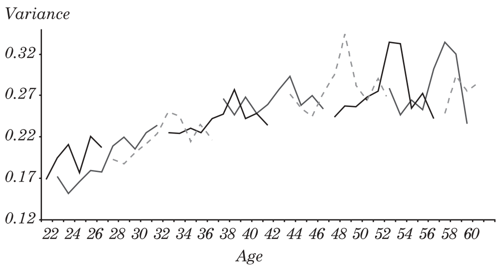
agents, income inequality increases over time whenever there is a permanent component in uncertainty. If all shocks were i.i.d., the distribution of income would be age independent. Furthermore, individuals actually start off at very different levels, as the initial variance is quite high.

The figure does not show important differences across cohorts. Except for the younger cohorts, the time path of the variance of earnings for any two consecutive cohorts typically cross, with no clear pattern. This means that individuals born in different years should not expect different levels of uncertainty at a given age. For all cohorts, the variance tends to peak around 1996, indicating the presence of time effects in the cross-sectional variance of income—perhaps aggregate fluctuations that change the dispersion of income.

Table 4 displays the sample autocovariance matrix of the residual of log income changes. The upper right triangle shows the covariances; the lower left triangle shows the correlations. We find high autocorrelations at the first order, followed by a steep decline at higher orders. These patterns suggest that income changes may be modeled as an MA(1) process.

Our benchmark estimates are reported in table 5, where we define income as annual individual earnings (without government transfers). Three cases are analyzed depending on whether the permanent component follows a random walk or an AR(1) stationary process and whether the transitory shock is i.i.d. or an MA(1) process. In all cases, the transitory component is not significant at the 5 percent level. The transitory shock does not show any persistence, either. These findings are consistent with the hypothesis that the transitory component is i.i.d at the individual level and that this component becomes negligible

Figure 5. Residual Variance across Cohorts: ESI, 1990–2000[a]



Source: Authors' calculations, based on ESI survey data (INE, 1990–2000).

a. The nine cohorts are defined by their age in 1990; from left to right, they are sixteen to twenty years of age; twenty-one to twenty-five years of age; twenty-six to thirty years of age; thirty-one to thirty-five years of age; thirty-six to forty years of age; forty-one to forty-five years of age; forty-six to fifty years of age; fifty-one to fifty-five years of age; and fifty-six to sixty years of age.

Table 4. Covariance Matrix of Log Income Changes: ESI[a]

Year	*1991*	*1992*	*1993*	*1996*	*1997*	*1998*	*1999*	*2000*
1991	0.000075	5.40E–06	2.20E–06	–0.000051	–3.10E–06	–7.40E–06	0.00053	–0.000029
1992	0.70860 (0.0491)	0.00005	–0.00004	–0.00007	–0.00002	0.00004	0.00001	0.00003
1993	–0.54160 (0.1656)	–0.87000 (0.005)	0.00022	–0.00019	–0.00009	0.00027	–0.00011	–0.00006
1996	–0.07240 (0.8774)	–0.53340 (0.2175)	–0.03530 (0.9402)	0.00042	0.00010	–0.00041	0.00011	–0.00001
1997	0.2951 (0.5205)	–0.27810 (0.5459)	0.41300 (0.3571)	0.32160 (0.4373)	0.00017	–0.00019	–0.00004	0.00011
1998	–0.10130 (0.8289)	0.31320 (0.4939)	0.23720 (0.6086)	–0.86830 (0.0052)	–0.55230 (0.1558)	0.00050	–0.00014	–0.00003
1999	0.28160 (0.5406)	0.13170 (0.7784)	–0.53580 (0.2152)	0.21680 (0.6061)	–0.34050 (0.4091)	–0.33580 (0.4162)	0.00019	–0.00015
2000	–0.19110 (0.7169)	0.26450 (0.6124)	–0.23820 (0.6494)	–0.11660 (0.8033)	0.49230 (0.2617)	–0.01620 (0.9725)	–0.59620 (0.1577)	0.00032

Source: Authors' calculations, based on ESI survey data (INE, 1990–2000).
a. Correlations are below the diagonal; covariances are above the diagonal. Statistical significance is in parentheses.

when we average within cohorts. In other words, the transitory component is indistinguishable from classical measurement error. The permanent component follows an AR(1) process, as the autocorrelation coefficient is statistically smaller than 1. The estimated variance of the permanent component is much larger than the variance of the transitory shock, but it is an order of magnitude smaller than the variance estimated by several authors using a panel of individual U.S. data from panel sets such as the PSID.[10] This large difference can also be explained by the fact that we track cohorts and not individuals over time.[11] We further investigate this hypothesis below.

Table 5 also estimates the dynamics of Chilean earnings using labor income plus government transfers. This exercise identifies the extent to which the government provides insurance through its monetary

10. See Meghir and Pistaferri (2004) for the most recent results.
11. See Pischke (1995) for a comparison of the variability and persistence of aggregate and individual income.

Table 5. The Dynamic Process of Labor Income: ESI[a]

Definition of income and income process[b]	*Permanent component*		*Transitory component*	
	Variance	*Autocorrelation*	*Variance*	*MA(1) coefficient*
Without transfers				
Permanent AR(1)	0.00395	0.93095	–0.00028	
Transitory i.i.d.	(0.00062)	(0.02830)	(0.00028)	
Permanent random walk	0.00326		0.00014	0.15868
Transitory MA(1)	(0.00080)		(0.00067)	(2.64100)
Permanent random walk	0.003028		0.000303	
Transitory i.i.d.	(0.00067)		(0.00264)	
With transfers				
Permanent AR(1)	0.00394	0.93077	–0.00026	
Transitory i.i.d.	(0.00062)	(0.02900)	(0.00028)	
Permanent random walk	0.00327		0.00014	0.15871
Transitory MA(1)	(0.00081)		(0.00068)	(2.6169)
Permanent random walk	0.00304		0.00031	
Transitory i.i.d.	(0.00069)		(0.00027)	

Source: Authors' calculations, based on ESI survey data (INE, 1990–2000).

a. Standard errors are in parentheses.

b. Income is defined as annual individual earnings, with and without government transfers. The assumed income processes vary depending on whether the permanent component follows a random walk or an AR(1) stationary process and whether the transitory shock is i.i.d. or an MA(1) process.

transfers. A number of papers analyze the role of government transfers in alleviating poverty and reducing income inequality in Chile (Baytelman, Cowan, and De Gregorio, 1999; Engel, Galetovic, and Raddatz, 1999). We analyze whether public transfers reduce the uncertainty faced by individuals.

The estimated income processes with and without transfers are very much alike. This is due to the fact that very few individuals report having received transfers in our data set.[12] However, a probit regression of a dummy indicating whether the individual received a positive transfer on the level of real earnings, as well as year, month, and regional dummies, yields a highly significant negative effect of

12. Only about 1.5 percent report a positive level of transfers. This underreporting of transfers might be explained by the fact that most monetary subsidies are paid through the worker's paycheck, which may lead individuals to incorrectly report transfers as part of their labor earnings.

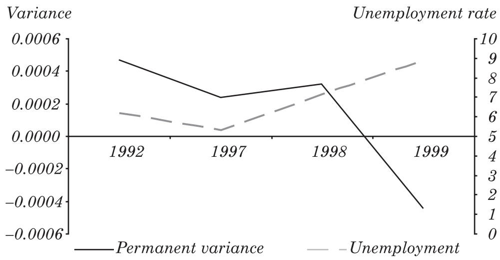
perceived income on the probability of receiving a transfer. Hence, public transfers do play a limited redistributive role in our sample.[13]

We have thus far restricted the variances of the shocks to be constant over time, although our modeling structure allows for time-varying variances. Following Meghir and Pistaferri (2004), we can identify the evolution of the variance of the permanent shock using the following moment condition:

$$E\left[\Delta\eta_t^c\left(\sum_{j=-(1+q)}^{(1+q)}\Delta\eta_{t+j}^c\right)\right]=E\left(\sigma_{\upsilon t}\right),$$

where c indexes cohorts, t indexes years, and q is the order of the MA (transitory) component. Figure 6 plots the estimated evolution of the variance of the permanent component, along with the unemployment rate. Unfortunately, our sample is short and was interrupted in 1994, so we are only able to estimate the variance for 1992, 1997, 1998, and 1999. Except for 1999, the unemployment rate and permanent variance behave in a synchronized manner. Our point estimate of the 1999 variance is negative, but not statistically significant. Our results thus show that income uncertainty correlates with the business cycle.

Figure 6. Variance of Permanent Shock and Unemployment



Source: Authors' calculations, based on ESI survey data (INE, 1990–2000).

13. The marginal effect is $-2.08 \bullet 10^{-8}$, so each additional 200,000 pesos of income (about one standard deviation in the sample) reduces the chance of receiving a public transfer by 0.42 percentage points.

3.2 Synthetic Cohorts and Variance Underestimation: Comparing Chile and the United States

To further investigate the hypothesis that cohort averaging leads to a large underestimation of the variances, we study whether our estimation process leads to similar results when we use data from the United States. Specifically, we compare our results to those obtained from a comparable sample taken from U.S. data, using a synthetic panel and individual-level data. Our source of information is the Panel Study of Income Dynamics (PSID), which is a representative longitudinal survey of nearly 8,000 households. The PSID started collecting data on individuals and households in 1968, and it has followed the same households and their split-offs on a yearly basis since then. The survey has rich data on a large number of economic and demographic variables. Below we exploit the fact that the PSID has a panel structure, which allows us to estimate the dynamics of income using individual data directly. We then reestimate the process using cohort data to analyze the way estimated parameters are affected by using a synthetic panel technique. We use the surveys from 1988 to 1997. Table 6 reports some sample descriptive statistics.

Our analysis of the U.S. data replicates the analysis of Chilean data. We first restrict our samples to men between the ages of twenty-five and sixty. We deflate wage income by the consumer price index for all urban consumers (CPI-U) and then estimate the predictable component of labor income using the same variables and functional form reported for Chile in the previous subsection. We construct a series for the unexplained portion of labor income for every individual in our sample. We use the sample weights to perform our estimates.

Table 6. Sample Descriptive Statistics: PSID

Variable	*Mean*	*Standard deviation*	*Minimum*	*Maximum*	*Median*
Annual labor income[a]	39,714	40,119	1,076	1,274,859	32,617
Age	39.3	8.5	25	60	38
Years of schooling	13.1	2.8	0	21	12
Household size	3.2	1.5	1	13	3
Married	0.8	0.4	0	1	1

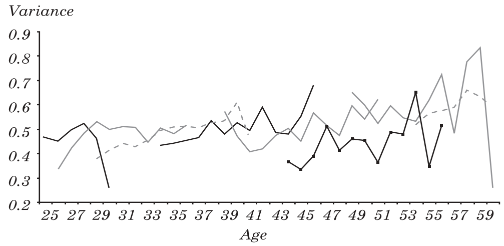
Source: PSID survey data (SRC, 1988–97).
a. In 1996 U.S. dollars.

Figure 7 plots the behavior of the U.S. within-cohort residual variance over the sample period. As in the Chilean case, the figures do not reveal the presence of a cohort effect. Two properties are not shared by the Chilean and U.S. profiles. First, the variance is quite flat over most of the life-cycle in the United States, whereas it increases with age in Chile. Second, the variance is much larger in the United States than in Chile—almost 2.05 times larger on average. This result seems counterintuitive, and it is not an artifact of the different currency denominations used to measure income, as income is measured in logs.[14]

This striking gap in earnings risk may be the result of institutional rigidities that reduce wage dispersion in the Chilean labor market relative to the U.S. labor market. As shown by Bertola and Ichino (1995), wage inequality is reduced in markets where workers move across firms, occupations, and regions in response to productivity and demand shocks. Wage dispersion differences may thus reflect different labor market institutions, labor reallocation costs, and wage contract structures.

Alternatively, once we allow for the endogeneity of earnings, a pos-

Figure 7. Residual Variance across Cohorts: PSID, 1988–97[a]



Source: Authors' calculations, based on PSID survey data (SRC, 1988–97).
a. The eight cohorts are defined by their age in 1988; from left to right, they are nineteen to twenty-four years of age; twenty-five to twenty-nine years of age; thirty to thirty-four years of age; thirty-five to thirty-nine years of age; forty to forty-four years of age; forty-five to forty-nine years of age; fifty to fifty-four years of age; and fifty-five to sixty years of age.

14. Survey methodologies and the extent of measurement error might partly explain these differences. However, the U.S. variance of log real earnings is higher than the Chilean variance even before conditioning on worker characteristics. Furthermore, the R^2 of the regression is much higher for the Chilean data than for the U.S. data, so a larger portion of total income variability in Chile is explained by the predictable part of earnings. Both facts are consistent with the notion that Chilean workers face much less income uncertainty than do their American counterparts. We obtained similar results using the 1990–2000 Current Population Survey (CPS); these results are available on request.

sible explanation is that U.S. workers are willing to face a much larger level of uncertainty than their Chilean counterparts, as they have more opportunities to share risks through the marketplace given their better-developed markets. Furthermore, the public welfare system is much larger in the United States than in Chile, and female labor force participation is much higher. Both of these circumstances provide insurance against negative shocks. Therefore, our results are consistent with the hypothesis that workers in the United States can afford to take more risks than workers in Chile, and they thus choose riskier occupations and jobs.

The gap between the variance in Chile and the United States is reduced as individuals age. This fact might also have an institutional explanation: minimum wage laws might have a larger effect on young workers in Chile than in the United States. Indeed, the Chilean real minimum wage rose 72 percent over the sample years, whereas the U.S. real minimum wage rose only 18 percent.

In table 7 we estimate the dynamics of income using the information on U.S. workers, assuming the process is described by a random walk plus an i.i.d. transitory disturbance. For comparison, the table repeats the results obtained using the ESI and then presents the results using synthetic cohorts from the PSID. As with the Chilean case, we find

Table 7. The Dynamic Process of Labor Income: Chile and the United States[a]

Data source and level of aggregation	*Permanent variance*	*Transitory variance*
ESI, cohorts	0.00303 (0.00067)	0.00030 (0.00026)
PSID, cohorts	0.01181 (0.00362)	0.00080 (0.00157)
PSID, individuals	0.08150 (0.00839)	0.11173 (0.00644)

Source: Authors' calculations, based on ESI and PSID survey data (INE, 1990–2000; SRC, 1988–97).
a. The simulations assume a random walk plus an i.i.d. transitory disturbance. Standard errors are in parentheses.

that the process can be described solely by a random walk, as the transitory shock averages out in the aggregate. Moreover, the variance of the permanent shock is much larger in the United States than in Chile, which confirms the results in figures 5 and 7.[15]

Finally, table 7 reports the estimated parameters using individual-level data from the PSID. We find that the variance of the permanent shock is one order of magnitude larger than the variance estimated using cohort data. We also find a significant variance of the transitory shock.[16] Our results are consistent with other analyses. For instance, Meghir and Pistaferri (2004) use a similar sample from the PSID and find that the variance of the permanent shock is 0.0313, whereas the variance of the transitory shock is between 0.008 and 0.03.[17]

If we extrapolate the information in the PSID exercises to the Chilean case, the variance of the permanent shock is one order of magnitude larger than the variance estimated using the panel of cohorts, that is, about 0.0209. The variance of the transitory shock is also likely to be different from zero. These results have important behavioral implications. First, if innovations are permanent and individuals are prudent, then precautionary savings can become quantitatively very important (Deaton, 1992). Second, the distribution of labor income can be persistently very unequal. Finally, the position of an individual on the income distribution is also highly persistent, as good and bad fortunes last forever. Below we provide simulation exercises to illustrate these points. We first simulate life-cycle paths of income using our estimated processes. We then use the simulated outcomes to build income distributions and estimate the likelihood that an individual will move along the income distribution.

4. Application to Earnings Mobility

Our application refers to income inequality and earnings persistence over the life cycle. We provide two sets of exercises that illustrate the effects of shock variance and persistence, using our benchmark estimates. The first set estimates transition matrices—the conditional probability that an individual will move along the income distribu-

15. The qualitative results are similar if we assume either an AR(1) plus an i.i.d. shock or a random walk plus an MA(1) shock.

16. Since we estimated the processes using a nonlinear methodology, we should not expect to find that the cohort-level variance is equal to the individual-level variance divided by the number of individuals in the cohort.

17. Meghir and Pistaferri (2004) allow for measurement error and show that the process for measurement error and for the transitory shock cannot be identified without external information. They find that the variance of the error in measurement must be between 0.01 and 0.03, assuming an MA(1) transitory shock. They estimate that the MA(1) coefficient is bounded between –0.18 and –0.25.

tion—that results from the estimated persistence of the dynamics of income. The second set of exercises estimates the distribution of income that results from the estimated variances.

To compute the transition matrices and the income distributions, we first generate 5,000 lifetime income streams based on our estimates. We assume a life-cycle of thirty-five years (ages twenty-five through sixty) and set the parameters of the model equal to the values estimated in table 5. We assume all individuals are identical at age twenty-five. Table 8 shows the simulated one-year transition matrices that result from assuming a random walk and an AR(1) process with first-order autocorrelation coefficient equal to 0.93095. The high persistence of the shocks implies that earnings mobility is very limited. For instance, an individual who starts off at the lowest quintile of the distribution has a 0.77–0.84 chance of staying there for another period. The likelihood of an individual at the richest quintile staying at that same quintile is quite similar. As expected, mobility is concentrated at the middle of the distribution, but the persistence is still quite high. Table 9 shows the simulated ten-year transition matrices (that is, the chance that an individual starting off at any given quintile will be at the same or at another quintile ten years ahead). Since the processes we estimate are highly persistent, mobility is rather limited even over a ten-year horizon. Figures 8 and 9 show the chance that an individual at any given point in life will be at the third and lowest quintiles, respectively, under the two alterna-

Table 8. Simulated Income Mobility: One-year Transition Matrix, ESI[a]

Income process and income quintile in period t	*Income quintile in t + 1* 1	2	3	4	5
Random walk					
Quintile 1	0.84	0.14	0.01	0.00	0.00
Quintile 2	0.14	0.65	0.19	0.02	0.00
Quintile 3	0.01	0.19	0.60	0.18	0.01
Quintile 4	0.00	0.02	0.19	0.65	0.15
Quintile 5	0.00	0.00	0.01	0.15	0.84
AR(1)					
Quintile 1	0.77	0.20	0.03	0.00	0.00
Quintile 2	0.20	0.50	0.25	0.05	0.00
Quintile 3	0.03	0.25	0.45	0.25	0.03
Quintile 4	0.00	0.05	0.25	0.50	0.20
Quintile 5	0.00	0.00	0.03	0.20	0.76

Source: Authors' calculations, based on ESI survey data (INE, 1990–2000).

a. The simulations assume a random walk and an AR(1) process with first-order autocorrelation coefficient equal to 0.93095.

Table 9. Simulated Income Mobility: Ten-year Transition Matrix, ESI[a]

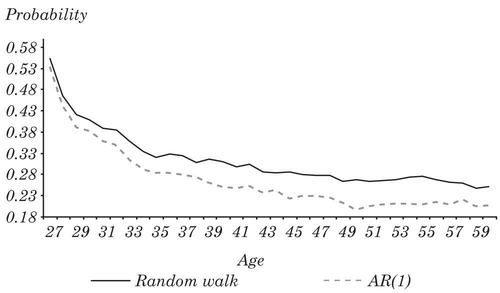
Income process and income quintile in period t	*Income quintile in t + 10*				
	1	*2*	*3*	*4*	*5*
Random walk					
Quintile 1	0.58	0.25	0.11	0.05	0.02
Quintile 2	0.25	0.32	0.24	0.13	0.05
Quintile 3	0.11	0.24	0.29	0.24	0.11
Quintile 4	0.04	0.14	0.24	0.32	0.26
Quintile 5	0.01	0.05	0.12	0.25	0.56
AR(1)					
Quintile 1	0.40	0.24	0.18	0.12	0.06
Quintile 2	0.25	0.24	0.21	0.18	0.12
Quintile 3	0.18	0.22	0.22	0.21	0.18
Quintile 4	0.12	0.18	0.21	0.24	0.25
Quintile 5	0.06	0.12	0.18	0.25	0.39

Source: Authors' calculations, based on ESI survey data (INE, 1990–2000).
a. The simulations assume a random walk and an AR(1) process with first-order autocorrelation coefficient equal to 0.93095.

Figure 8. Income Mobility and Lifetime Persistence: Third Quintile

Source: Authors' calculations, based on ESI survey data (INE, 1990–2000).

Figure 9. Income Mobility and Lifetime Persistence: Lowest Quintile



Source: Authors' calculations, based on ESI survey data (INE, 1990–2000).

tive levels of persistence. A major implication of our results is that poverty and the distribution of income in Chile should be quite persistent.

The variance level of the income process has important implications for the skewness of the income distribution. Table 10 shows the share of total income of individuals at different positions of the income distribution and at different ages. The first column reports the simulation results using the estimated variance. The second column uses the scaled ESI variance to account for the underestimation implied by our cohort technique. For instance, the richest 20 percent of individuals at age thirty receive a share of income that is 1.41 times the share of the poorest 20 percent, assuming our uncorrected benchmark estimates. This ratio increases to 2.48 when we scale the variance according to our PSID results. Because of the persistence of our estimates, simulated inequality increases with age. Nevertheless, our simulations cannot match actual income disparities. According to the CASEN, the richest quintile receives a share of about 13.8 times the share of the lowest quintile. Our underestimation is the result of the

Table 10. Simulated Income Distributions

Age of individuals	*Income shares*		*Share V / Share I*	
	ESI	*Scaled ESI*	*ESI*	*Scaled ESI*
30 years	0.17	0.12	1.41	2.48
	0.19	0.16		
	0.20	0.19		
	0.21	0.23		
	0.24	0.30		
40 years	0.15	0.08	1.80	4.62
	0.18	0.13		
	0.20	0.17		
	0.22	0.23		
	0.26	0.38		
50 years	0.13	0.06	2.14	7.43
	0.17	0.11		
	0.19	0.16		
	0.22	0.23		
	0.28	0.45		
60 years	0.12	0.05	2.46	10.59
	0.16	0.09		
	0.19	0.14		
	0.23	0.22		
	0.30	0.50		

Source: Authors' calculations, based on ESI survey data (INE, 1990–2000).

assumption that all individuals start off with the same characteristics. In particular, we assume the same educational attainment across workers. As discussed earlier, figure 4 shows that schooling might explain a large portion of income inequality. Still, once we take the scaled variances, a large portion of actual inequality is explained by the underlying variability of the process of individual income.

5. Concluding Remarks

In this paper, we have estimated the dynamic process of individual income using the *Encuesta Suplementaria de Ingresos*. We found the income process to be characterized by high persistence but low variability. We also showed that the low variance is an artifact of our cohort technique. Using data for U.S. workers, we found that averaging over cohorts leads to variance underestimation of one order of magnitude. Future work should directly address the issue of underestimation, which requires long panel sets of data on individual income. Different organizations in Chile have collected rich panel data sets that follow workers over long periods of time and on a monthly basis. Unfortunately, these data sets currently are not publicly available.

APPENDIX
Estimating Dynamics Using the CASEN Data

Most of the existing analyses of Chilean household income are based on data from the CASEN survey. Unfortunately, the CASEN only gathers information every other year. Based on the ESI, our results show that the dynamics of earnings is highly persistent. The CASEN thus misses most of the action in the two-year lag. We formally show in this appendix that it is not possible to estimate our income processes using this survey.

Our benchmark model assumes that the stochastic component of income can be decomposed into permanent and transitory shocks:

$$\eta_{it} = y_{it}^{p} + \mu_{it} + \omega_{it} \ .$$

If the permanent shock follows a random walk and the transitory shock follows an MA(1) process, then

$$y_{it}^{p} = y_{it-1}^{p} + \upsilon_{it} \ \text{and}$$

$$\mu_{it} = \varepsilon_{it} - \theta\varepsilon_{it-1} \ .$$

We identify the model's parameters using GMM by matching population and sample moments. To use the CASEN, we would then need to match the covariances of second differences; that is,

$$\Delta_2\eta_{it} = \eta_{it} - \eta_{it-2} = \upsilon_{it} + \upsilon_{it-1} + \varepsilon_{it} - \theta\varepsilon_{it-1} - \varepsilon_{it-2} + \theta\varepsilon_{it-3} \ .$$

The relevant population moments are thus as follows:

$$\text{var}\left(\Delta_2\eta_{it}\right) = \sigma_{\upsilon t} + \sigma_{\upsilon t-1} + \sigma_{\varepsilon t} + \theta^2\sigma_{\varepsilon t-1} + \sigma_{\varepsilon t-2} + \theta^2\sigma_{\varepsilon t-3} \ ,$$

$$\text{cov}\left(\Delta_2\eta_{it}, \Delta_2\eta_{it-2}\right) = -\sigma_{\varepsilon t-2} - \theta^2\sigma_{\varepsilon t-3} \ , \ \text{and}$$

$$\text{cov}\left(\Delta_2\eta_{it}, \Delta_2\eta_{it-j}\right) = 0 \ , \ \text{for } j \geq 4 \ .$$

Assume that variances are time independent. The identification conditions then reduce to

$$\text{var}\left(\Delta_2\eta_{it}\right) = 2\sigma_{\upsilon} + 2\left(1+\theta^2\right)\sigma_{\varepsilon} \ ,$$

$\text{cov}\left(\Delta_2\eta_{it}, \Delta_2\eta_{it-2}\right) = -\left(1+\theta^2\right)\sigma_\varepsilon$, and

$\text{cov}\left(\Delta_2\eta_{it}, \Delta_2\eta_{it-j}\right) = 0$, or $j \geq 4$.

This model does not have a unique solution: only the first two equations provide useful information, but there are three unknown parameters to solve for. Thus, the CASEN does not have enough information to estimate the underlying parameters of our model.

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Trade Orientation and Labor Market Evolution: Evidence from Chilean Plant-level Data

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Many developing and developed economies consider structural reforms to trade and fiscal policy that are designed to lower taxes and tariffs and stimulate investment and production of the manufacturing sector. A good example of such a country is Chile, which went through a series of structural reforms in the late 1970s and early 1980s. The labor and financial markets were deregulated and price controls eliminated. Two major tax reforms were put into operation in 1975 and 1984, and a social security reform was introduced in 1980. In addition, Chile was one of the first countries in Latin America to begin a gradual but deep trade liberalization process. In 1967 the average effective protection rate was over 100 percent. Between 1973 and 1979, Chile eliminated the quantitative restrictions and reduced the import tariff to a uniform level of 10 percent. A debt crisis in 1982 led to the delay or partial reversal of some reforms (the import tariffs were temporarily increased to 35 percent in 1984), but by 1992 all of them were successfully in place.

Such dramatic changes in the free-trade environment should have first-order implications for labor markets in Chile. The standard Heckscher-Ohlin model predicts that a low-labor-cost country like Chile trading with high-labor-cost developed economies such as the United States will experience a fall in the capital-labor ratio and a reduction in demand for skilled workers relative to unskilled workers once trade barriers are reduced. More recent theories lead to the opposite conclusion, however, if trade liberalization is associated with

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the adoption of new technologies or a shift toward importing high-technology capital goods that complement skilled labor. In this case, trade liberalization may lead to rising capital-labor ratios and a shift toward skilled labor relative to unskilled labor. Trade liberalization may also imply increased wage inequality owing to such skill-biased technological change.

Existing research provides strong support for the notion that technological change is indeed skill biased, and that this skill bias is transmitted across countries following trade liberalizations. Empirical evidence for member countries of the Organization for Economic Cooperation and Development (OECD) suggests that unskilled workers experienced a deterioration in their wages over the last two decades despite their increasing relative scarcity. Most industries in these countries saw an increasing participation of skilled workers in the labor force despite the fact that their wages have increased or remained stable relative to unskilled workers.

The evolution of the wage structure in developed countries has become an important research topic in the last few years.[1] The empirical evidence is consistent with a considerable rise in wage inequality and demand for skilled workers in the United States and the United Kingdom and only a moderate increase in countries like Japan, Sweden, and Germany (Machin and Van Reenen, 1998).

This literature advances several hypotheses to explain the increased demand for skilled relative to unskilled workers in developed economies, including skill-biased technological change and Stolper-Samuelson effects of exposure to trade. While there is no consensus, researchers tend to agree that the main force behind the behavior of relative wages and the relative demand for skilled versus unskilled workers in developed economies is the presence of pervasive skill-biased technological change.

Only a few studies analyze changes in wage and labor structure in developing economies. For the case of Mexico, the findings suggest that returns to higher education increased between the late 1980s and mid-1990s (Meza, 1999) and that the shifts in the relative demand for skilled workers have taken place mostly within industries. Cragg and Epelbaum (1994) find evidence to support capital-skill complementarity in explaining the increase in the wage dispersion. Hanson and Harrison (1999) explain the increase in wage inequality in Mexican firms in the late 1980s by arguing that the reduction in

1. See Katz and Autor (1999) for a survey.

trade protection that took place in 1985 affected relatively low-skilled industries, which received high trade protection before the liberalization process. Revenga (1997) finds similar results. Robbins (1994, 1999) presents evidence of higher wage inequality following trade liberalization in the case of Chile. For Colombia, the results are mixed: wage disparity initially fell following the trade reform, but the relative wage for skilled workers increased after 1987.

Cross-country analyses provide some evidence of skill-biased technical transfer. Berman and Machin (2000), using data for middle-income countries, find increasing demand for skilled workers, which is concentrated in the same industries as in developed countries and is highly correlated with indicators of OECD technical change. Robbins (1995) points to a high correlation between the increasing demand for skilled labor and imports of machinery and equipment, also known as the skill-enhancing-trade hypothesis.

In a closely related study, Pavcnik (2002b) examines the evolution of the white-collar share for Chilean manufacturing plants over the period 1979–86. Pavcnik finds evidence in favor of skill-biased technological change and capital-skill complementarity. Building on her approach, we extend the analysis over an additional nine years to cover the period 1979–95. This extended data is well suited to analyzing long-run trend issues, such as the evolution of the skill bias in the Chilean labor market following such significant economic reforms. Unlike Pavcnik, we also disaggregate the data by trade orientation, classifying firms by whether they are in export-oriented, import-competing, or nontradable sectors. In this way, we examine and compare the evolution of labor composition, wage premiums for skilled workers, the role of skill-biased technological change, and capital-skill complementarity across sectors with different trade orientations.

Skill-biased technological change is often linked to the adoption of new technologies through imported inputs following trade liberalization. The ability to adopt such technologies may be sector specific, in which case we expect to see systematic differences in the demand for skilled workers, as well as the evolution of the skilled-worker share across sectors classified by trade orientation. Dividing our sample based on trade orientation also provides a way of measuring the degree to which rising demand for skilled workers is broad-based or narrowly focused. To the extent that different sectors have different relative demands for skilled workers, our sectoral analysis also helps us understand the degree to which aggregate labor market outcomes may be influenced by economic fluctuations and economic policies that directly influence relative sectoral demands.

Our paper begins with a descriptive exercise, characterizing the broad movements in factor intensity, labor composition, and wage structure between skilled and unskilled workers over the period 1979–95. Our findings imply that rising demand for skilled workers is broad-based but somewhat uneven. The wage-bill share for white-collar workers rose in all three sectors of manufacturing. The nontradable sector shows the largest increase. The effect of a sharp rise in the white-collar share for the nontradable sector is diminished somewhat in the aggregate, as manufacturing production activity shifted away from nontradables toward exports over this period.

After completing this descriptive exercise, we consider a more formal analysis of the relation between trade orientation and labor market outcomes. We adopt a cost-minimization approach based on a restricted variable translog cost function to provide direct estimates of the relative demand for skilled workers. The same methodology has been used to study the presence of skill-biased technological change in both developed economies (Berman and Machin, 2000) and developing economies (Pavcnik, 2002b).

According to our analysis, most of the change in the relative demand for skilled workers, as well as the shifts in the share of skilled labor in the wage bill, occurred within rather than between industries. This finding provides preliminary support for the existence of skill-biased technological change. Our regression analysis provides evidence that the white-collar wage share is strongly associated with measures of technology adoption across all three sectors of manufacturing. These results offer further evidence in favor of skill-biased technological change.

Our results also suggest that capital-skill complementarity is a determinant of the wage-bill share for import-competing industries. This finding is buttressed by the fact that the import-competing sector exhibited more capital deepening than the export-oriented and nontradable sectors over this period. As a caveat, however, we note that our coefficient estimates imply that the size of the capital-skill complementarity is economically small. In addition, we find little evidence that capital-skill complementarity affects the relative demand for skilled workers in the export-oriented and nontradable sectors. Since the nontradable sector showed the sharpest rise in both the wage-bill share and the wage premium for skilled workers, our findings suggest that pure capital-skill complementarity is unlikely to be the primary driving force behind skill-biased technological change. Rather, more general forms of technological adoption are likely explanations for the rising demand for skilled workers in Chile over this time period.

The paper is organized in the following way. Section 1 gives a brief summary of the empirical evidence in favor of skill-biased technological change in developed economies. Section 2 provides descriptive statistics documenting the composition and evolution of manufacturing employment. Here we divide plants into industrial sectors based on their trade orientation. We also provide summary statistics regarding capital intensity and growth rates for value-added and factor inputs for each of these sectors. In section 3, we provide a more formal analysis of the evolution of the mix of skilled and unskilled worker: we decompose shifts in the labor share for skilled workers into within- and between-industry variations, and we use a cost-minimization approach to study the relation between labor composition, capital deepening, and technology adoption. Section 4 concludes.

1. Empirical Evidence in Favor of Skill-biased Technological Change

The empirical evidence for the United States and OECD countries suggests that less-skilled workers experienced a decrease in relative wages along with increasing unemployment rates during the 1980s. In the United States, real wages for young men with twelve or fewer years of education fell by 26 percent between 1979 and 1993, without signs of recovery since. Similar patterns characterized the evolution of relative wages of less-skilled workers in several OECD countries. It is a well-documented fact that labor market outcomes for less-skilled workers in developed economies deteriorated in the past two decades despite the increasing relative supply of skilled workers.

Two hypotheses have been analyzed to explain the evolution of labor composition in developed economies: the Stolper-Samuelson mechanism and skill-biased technological change. The standard Heckscher-Ohlin model predicts that when a skill-abundant country like the United States trades with less-skill-abundant developing countries, the skill-abundant country will experience an increase in the relative price of the skill-intensive good, which translates into an increase in the relative wage of skilled workers. An alternative explanation is pervasive skill-biased technological change, in which the economy experiences an increase in the proportion of skilled labor and relative wages, as well as within-industry skill upgrading. According to this hypothesis, the presence of pervasive, sector-neutral, skill-biased technological change is a potential explanation for the existence of an increasing skilled-unskilled wage premium even in the case of a small open economy. It is important to notice that, unlike the Stolper-Samuelson effect, pervasive skill-biased technological change induces within-industry increases in the share of skilled workers employed.

So far, empirical research suggests that the main explanation behind the evolution of relative labor and wages in the developed world is pervasive skill-biased technological change. The arguments in favor of this hypothesis can be summarized as follows: (1) the increase in skill intensity and wage premium have occurred within, rather than between, industries; (2) these observed shifts tend to be concentrated in the same industries across countries; (3) capital-skill complementarity seems to be small (Berman and Machin, 2000); and (4) employment shifts to skill-intensive sectors appear to be too small to be consistent with the notion that international trade mechanisms are the prime determinants of the changing skill mix. Table 1 presents the annualized change in employment and wage-bill shares for a group of developed economies.

For a developing economy like Chile, the Stolper-Samuelson mechanism implies the opposite effect in terms of the wage premium. In particular, a less-developed country like Chile that trades with skill-abundant developed economies should experience a decrease in the labor share of skilled workers and a change in the relative wage in favor of unskilled workers. According to the evidence presented in

Table 1. Skilled to Unskilled Labor and Wage-bill Shares: Developed Economies, 1980–90
Percent

	Country						
Indicator	*United States*	*United Kingdom*	*Japan*	*Sweden*	*Denmark*	*Austria*	*Belgium*
Annual change in labor share	0.30	0.29	0.06	0.12	0.41	0.16	0.16
—Within-industry share (% of total change)	73	93	123	60	87	68	96
Annual change in wage-bill share	0.51	0.62	0.14	0.07	0.64	0.36	−0.06
—Within-industry share (% of total change)	76	92	84	25	89	76	92
Change in wage ratio, 1980–90	7	14	3	−3	7	7	−5

Source: Berman, Bound, and Machin (1998).

this paper, the labor and wage-bill shares for Chilean manufacturing plants went in the opposite direction, an observation that is inconsistent with the classical Stolper-Samuelson explanation of exposure to trade. With this in mind, we analyze whether the presence of skill-biased technological change is a plausible explanation for the change in the labor composition and relative wages in the Chilean economy between 1979 and 1995.

2. Data Overview

Given the macroeconomic volatility and structural changes that occurred over this period, the use of a large panel data from 1979 to 1995 is particularly important for understanding both wages and employment dynamics at the plant level. Previous research on employment and productivity dynamics using information for Chilean manufacturing plants only considers information between 1979 and 1986.[2] The topics analyzed are related to the effects of trade liberalization in total factor productivity, the role of plant exit and entry on manufacturing productivity growth, the effect of trade in total employment movements, and the role of the adoption of foreign technology in explaining the evolution of the relative demand for skilled workers.[3]

Our current set of plant-level data for Chilean manufacturing plants was obtained from the World Bank and the National Statistics Institute of Chile (INE); the data were collected by INE. In the cleaned sample, we have an average of 3,906 plants per year with ten or more employees in the manufacturing sector. The data set contains annual information for the period 1979–95, and it includes a large set of variables about production, employment, investment, capital stocks, intermediate inputs, and plant entry and exit. All variables considered are in terms of 1980 prices. After the elimination of extreme outliers, this panel data contains 66,406 observations across plants and years.

2. This coincides with the years for which Chilean plant-level data were obtained and made available by the World Bank.

3. Pavnik (2002a), using information for Chilean industrial plants, concludes that productivity increased by 3–10 percent in the import-competing sector as a result of trade liberalization, but findings for the export-oriented sector are not conclusive. Liu and Tybout (1996) and Tybout (1996) use the 1979–86 sample to study productivity dynamics at the plant level while Levinsohn (1996) study job creation and destruction based on this data set.

We constructed appropriately defined capital indices using the perpetual inventory method, aggregated material inputs using industry-level deflators, and put all variables on a comparably deflated basis.

Employment is measured as the number of workers hired per year and is decomposed by skill-type: white collar and blue collar. Given that we want to study the relation between employment composition according to skill level, trade orientation, and technology adoption, we needed proxies for the technology measure. The proxies for technology use provided by the data are imported materials and expenditures on foreign technical assistance. Unfortunately, we do not have information on foreign direct investment or on expenditures on research and development, which are the variables commonly chosen as ideal proxies for technology measures.

2.1 Sectoral Classification

To classify plants based on their trade orientation, we rely on information on imports and exports from the World Trade Analyzer CD-ROM from Statistics Canada. The level of disaggregation in the information obtained from Statistics Canada allowed us to improve on previous definitions provided by Liu (1991) by updating the information between 1987 and 1995. In particular, plants that belong to a three-digit industry in which more than 15 percent of the industry's output is exported were characterized as export-oriented plants. Likewise, plants in an industry in which the ratio of total imports to total domestic output is higher than 15 percent were characterized as import-competing. The remaining plants were classified as belonging to the nontradable sector.

Table 2 summarizes the sectoral classification across three-digit industries, while tables 3, 4, and 5 document the evolution of plant size and the share of manufacturing value-added and employment accounted for by each sector. Unsurprisingly, table 2 indicates that export-oriented industries are concentrated in wood, paper, and mining, while import-competing industries are much more heterogeneous.

Table 3 provides the sample means for the number of employees per plant, for both the full-sample and the subsamples in which industries are split based on trade orientation. On average, Chilean plants are much smaller than their developed country (U.S.) counterparts. Plants in export-oriented industries are larger than other plants, and this size discrepancy increases over the sample period. In 1979, export-oriented plants were 26 percent larger than the average plant,

Table 2. Industrial Composition of Sectors by Trade Orientation
Percent

		Sector		
Code	*Description*	*Export-oriented*	*Import-competing*	*Nontradable*
311	Food	15	16	69
313	Beverage	–	–	100
314	Tobacco	–	–	100
321	Textiles	–	100	–
322	Apparel	–	100	–
323	Leather products	–	–	100
324	Footwear		100	
331	Wood products	100	–	–
332	Furniture	–	–	100
341	Paper	100	–	–
342	Printing	–	–	100
351	Industrial chemicals	–	100	–
352	Other chemicals	–	–	100
353	Petroleum refining	–	–	100
354	Misc. petroleum products	–	–	100
355	Rubber	–	–	100
356	Plastics	–	100	–
361	Ceramics	–	100	–
362	Glass	–	100	–
369	Nonmetallic minerals	–	–	100
371	Iron and steel	–	–	100
372	Nonferrous metals	100	–	–
381	Metal products	–	100	–
382	Nonelectric machinery	–	100	–
383	Electric machinery	–	100	–
384	Transport equipment	–	100	–
385	Professional equipment	–	100	–
390	Miscellaneous	–	100	–

Source: Authors' calculations.

while in 1995, this size discrepancy had increased to 46 percent. At the beginning of the sample, import-competing plants were also significantly larger than plants in the nontradable sector. This difference erodes over time, however. Using labor as a measure of size, the overall finding from table 3 is that plant size in the export-oriented sector appears to have expanded much more than plant size in other sectors.[4]

The increase in plant size for export-oriented firms occurred in conjunction with an overall expansion of the export-oriented sector relative to the other two sectors. Table 4 documents the share of value-added

4. This does not necessarily imply that total employment has increased more rapidly for export-oriented sectors relative to import-competing sectors, however.

accounted for by plants in each sector. According to our sample, the export sector's share of value-added rose from 14 percent to 19 percent over the sample period. The import-competing share increased slightly and the nontradable share fell almost 20 percent during this time. The import-competing sector, however, accounts for the largest component of manufacturing economic activity—on the order of 50 percent.

Table 5 documents the share of manufacturing employment accounted for by each sector. For the export-oriented sector, the employment share shows an increase similar to that of the value-added share. In contrast to the value-added share, however, the employment share for the import-competing sector fell somewhat over this period. The employment share for the nontradable sector also fell, though the drop is muted relative to the drop in the value-added share of this sector.

2.2 Sectoral Dynamics and Factor Intensity

Figure 1 documents the growth rates of value-added for each sector. These growth rates display similar cyclical patterns over time, with the exception that the export-oriented sector expanded rapidly during the early 1980s when the rest of the manufacturing sector was mired in recession. Over the full sample period, the import-competing sector grew faster, at 6.2 percent annually, than the export-oriented and nontradable sectors, which grew at 5.6 percent and 5.1 percent, respectively.

Table 3. Mean of Total Employment in Plants by Sector

	Sector			
Year	*Export-oriented*	*Import-competing*	*Nontradable*	*Full sample*
1979	58	52	34	46
1980	64	51	37	47
1981	69	55	37	49
1982	60	50	35	45
1983	65	49	36	45
1984	70	53	37	49
1985	73	57	39	52
1986	90	67	45	60
1987	91	64	45	60
1988	96	69	46	64
1989	101	72	48	67
1990	100	73	50	68
1991	98	69	49	66
1992	93	69	49	65
1993	93	69	50	65
1994	94	68	50	65
1995	95	67	50	65

Source: Authors' calculations.

Table 4. Value-added Share by Sector

Year	Sector		
	Export-oriented	*Import-competing*	*Nontradable*
1979	0.141	0.513	0.346
1980	0.127	0.505	0.368
1981	0.149	0.495	0.357
1982	0.177	0.470	0.353
1983	0.192	0.442	0.366
1984	0.194	0.469	0.337
1985	0.207	0.471	0.322
1986	0.177	0.486	0.336
1987	0.184	0.499	0.317
1988	0.220	0.498	0.281
1989	0.211	0.489	0.300
1990	0.195	0.503	0.302
1991	0.199	0.512	0.289
1992	0.197	0.504	0.299
1993	0.193	0.517	0.290
1994	0.189	0.530	0.281
1995	0.190	0.531	0.278

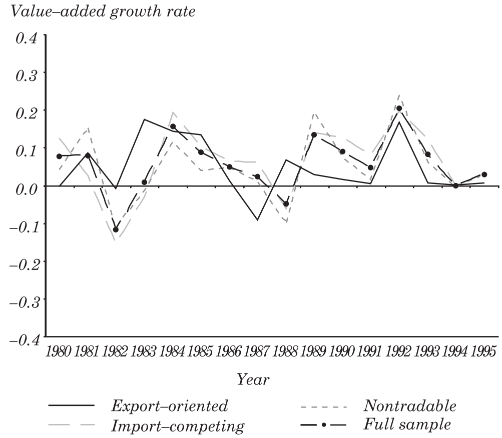
Source: Authors' calculations.

Table 5. Total Employment Share by Sector

Year	Sector		
	Export-oriented	*Import-competing*	*Nontradable*
1979	0.154	0.558	0.288
1980	0.145	0.540	0.315
1981	0.163	0.529	0.308
1982	0.156	0.515	0.329
1983	0.177	0.491	0.332
1984	0.190	0.507	0.302
1985	0.200	0.504	0.295
1986	0.164	0.526	0.310
1987	0.192	0.527	0.280
1988	0.199	0.536	0.264
1989	0.202	0.530	0.268
1990	0.200	0.524	0.275
1991	0.208	0.515	0.277
1992	0.202	0.520	0.277
1993	0.205	0.518	0.276
1994	0.209	0.518	0.274
1995	0.205	0.529	0.266

Source: Authors' calculations.

Figure 1. Value-added Growth



Source: Authors' calculations.

The overall growth rate for our manufacturing sample was 5.8 percent over this period.[5]

Figures 2 and 3 document the evolution of labor productivity (output per employee) and capital productivity (output per unit of capital) for each sector over the 1979–95 period, while table 6 provides average annual growth rates over this period.[6] Labor productivity grew most rapidly in the export-oriented sector, at an average annual rate of 3.8 percent, and least rapidly in the nontradable sector, at 2.2 percent on average. Capital productivity grew rapidly for the export and nontradable sectors, at an average annual rate of 3.3 percent and 4.2 percent, respectively. Measured by output per unit of capital, the import-competing sector saw a substantially greater increase in capital

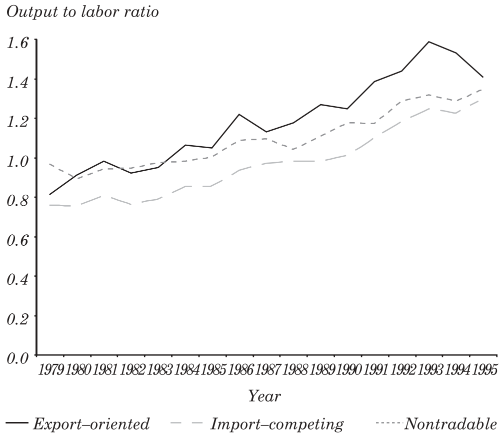
5. Because our data is a sample rather than the full universe of manufacturing plants, we measure growth rates for plants that are in the sample over consecutive periods. Let n_{t-1} denote the set of plants with observations available for both t and $t-1$. The growth rate of value-added is then computed as

$$g_t^{VA} = \log\left(\sum_{i \text{ in } n_{t-1}} VA_{it}\right) - \log\left(\sum_{i \text{ in } n_{t-1}} VA_{it-1}\right).$$

6. Because the capital stock data are not available for plants that enter the sample after 1981, there is likely some bias in the labor-intensity and capital-output ratios and the total factor productivity numbers documented in figure 3 and table 6. We thus treat these numbers as informative rather than definitive. The labor productivity numbers are not subject to such potential biases.

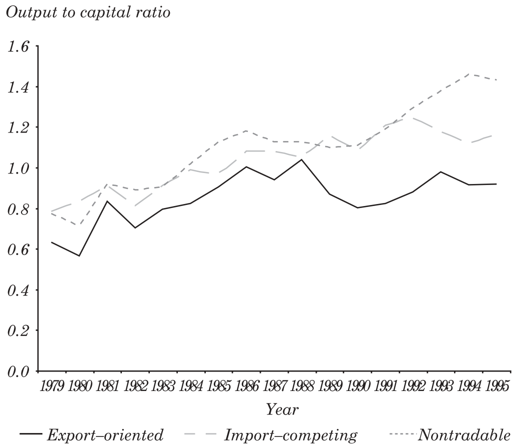
intensity relative to the other two sectors over this period. The finding that the import-competing sector became more capital intensive relative to the other two sectors over time is consistent with the notion that trade liberalization allowed import-competing firms to increase their capital intensity through the adoption of imported machinery. When

Figure 2. Labor Productivity



Source: Authors' calculations.

Figure 3. Capital Productivity



Source: Authors' calculations.

Table 6. Average Annual Growth Rate by Sector, 1979–95
Percent

Indicator	Sector: Export-oriented	Import-competing	Nontradable	Full sample
Output to capital	3.3	2.7	4.2	3.3
Output to labor	3.8	3.5	2.2	2.9
Labor to capital	–0.5	–0.8	2.0	0.4
Total factor productivity	3.6	3.2	2.9	3.1

Source: Authors' calculations.

we measure total factor productivity as a weighted average of labor and capital productivity, these numbers imply substantial gains in productivity for the export-oriented sector relative to the import-competing and nontradable sectors.

2.3 Labor Composition by Sector

We now document trends in labor composition between skilled and unskilled workers. Table 7 summarizes the evolution of the white-collar share of total employment. Here we report the ratio of white-collar employees in each sector to the total number of employees in each sector. The share of white-collar workers in total employment is higher in the import-competing sector relative to the export-oriented sector. There is no significant difference in terms of skill composition between the import-oriented and nontradable sectors. The overall share of skilled workers in total employment displays moderate increases over time, showing a rise of 8 percent for the full sample in the period 1979–95. In contrast to the import-competing and nontradable sectors, the white-collar share of employment in the export-oriented sector shows no change.

Table 8 provides further information regarding the evolution of the skill mix between white- and blue-collar workers by documenting the evolution of the wage-bill share for white-collar workers. In all sectors, the wage-bill share rose more rapidly than the labor share, implying that wage differentials between white- and blue-collar workers rose over time. The wage-bill share increased by 10 percent for the import-competing and export-oriented sectors and by 26 percent for the nontradable sector over the sample period. As summarized in table 9, these results imply an increase in the wage premium for skilled workers for the period 1979–95 on the order of 9–16 percent, depending on

Table 7. White-collar Share in Total Employment by Sector

Year	Sector: Export-oriented	Import-competing	Nontradable	Full sample
1979	0.205	0.258	0.250	0.248
1980	0.216	0.269	0.255	0.258
1981	0.223	0.271	0.254	0.258
1982	0.249	0.298	0.272	0.281
1983	0.228	0.302	0.278	0.282
1984	0.215	0.290	0.270	0.272
1985	0.207	0.281	0.267	0.265
1986	0.236	0.286	0.280	0.278
1987	0.212	0.278	0.288	0.273
1988	0.213	0.289	0.288	0.279
1989	0.194	0.275	0.291	0.270
1990	0.204	0.272	0.288	0.269
1991	0.203	0.278	0.289	0.272
1992	0.205	0.269	0.287	0.267
1993	0.204	0.268	0.288	0.266
1994	0.200	0.272	0.286	0.266
1995	0.208	0.271	0.286	0.267

Source: Authors' calculations.

Table 8. White-collar Share in Total Wage Bill by Sector

Year	Sector: Export-oriented	Import-competing	Nontradable	Full sample
1979	0.303	0.339	0.263	0.304
1980	0.298	0.324	0.260	0.296
1981	0.306	0.343	0.259	0.304
1982	0.337	0.374	0.293	0.336
1983	0.343	0.381	0.296	0.341
1984	0.335	0.377	0.293	0.336
1985	0.332	0.379	0.294	0.338
1986	0.347	0.371	0.310	0.343
1987	0.334	0.380	0.330	0.354
1988	0.331	0.381	0.327	0.354
1989	0.327	0.377	0.338	0.356
1990	0.348	0.380	0.345	0.362
1991	0.351	0.385	0.348	0.366
1992	0.340	0.380	0.344	0.361
1993	0.336	0.379	0.335	0.357
1994	0.312	0.386	0.335	0.357
1995	0.330	0.374	0.332	0.352

Source: Authors' calculations.

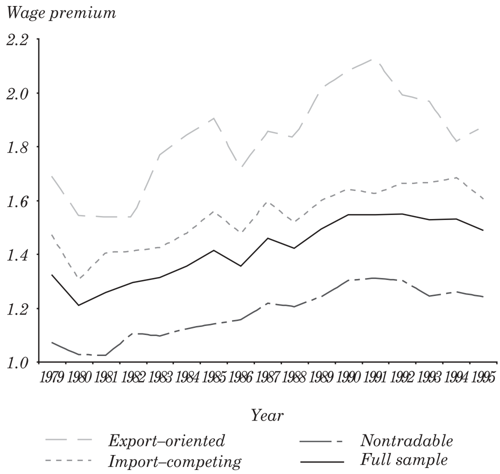
Table 9. Skilled-to-unskilled-labor and Wage Bill Shares: Chilean Manufacturing, 1979–95

Percent

	Sector			
Indicator	*Export-oriented*	*Import-competing*	*Nontradable*	*Full sample*
Annual change in labor share	0.32	0.37	0.87	0.45
—Within-industry share	87	70	75	70
Annual change in wage-bill share	0.63	0.66	1.54	0.95
—Within-industry share	74	88	40	65
Change in wage ratio, 1979–95	11	9	16	12

Source: Authors' calculations.

Figure 4. Wage Premium by Sector



Source: Authors' calculations.

the sector.[7] The evolution of the wage premium by trade orientation of the sectors is presented in figure 4. In table 9, the annualized change in the labor and wage-bill shares was higher than for developed economies, based on a comparison with table 1.

7. The decomposition of within- and between-industry shifts is presented in section 3.

3. Empirical Analysis of the Determinants of the Wage Bill Share at the Plant Level

To summarize the results so far, we have found that plants in the export-oriented sector expanded more rapidly than plants in the import-competing and nontradable sectors, both at the plant level and as a share of manufacturing output and employment. While both the export-oriented and nontradable sectors saw a higher rise in output per unit of capital than the overall increase, the import-competing sector became relatively more capital intensive by this metric. All three sectors showed increases in the demand for skilled workers relative to unskilled workers as measured by the wage-bill share, with the largest increase occurring in the nontradable sector (26 percent). The rise in the wage bill for export-oriented and import-competing firms are comparable—on the order of 10 percent for each sector. Since the import-competing sector appears to have experienced the most capital deepening, the sharp rise in the relative demand for skilled workers in the nontradable sector suggests that capital-skill complementarity is not the main driving force behind skill-biased technological change. We consider this issue further in our regression analysis.

We begin our empirical analysis of the determinants of the demand for skilled workers relative to unskilled workers by decomposing the overall change in the labor share of skilled workers relative to total workers, ΔS_t, into within- versus between-industry shifts in employment. One of the arguments in favor of skill-biased technological change is that the shifts in the share of skilled workers in total employment and in the total wage bill take place within rather than between industries. The decomposition of the change in the labor share over a period of time is given by

$$\Delta S_t = \sum_i \Delta s_{it} E_{i.} + \sum_i \Delta E_{it} s_{i.} \tag{1}$$

where s_{it} is the share of white-collar labor in total employment for industry i and year t, and E_{it} is the share of industry i's employment in total aggregate employment in year t. $E_{i.}$ and $s_{i.}$ denote industry means over time for E_{it} and s_{it}, respectively. The first term on the right-hand side of equation (1) measures the within-industry variation, while the second represents the between-industry contribution to the total change in the share ΔS_t. We compute an analogous decomposition for the wage-bill share. These results are summarized in table 10.

Table 10. Decomposition of Relative Labor Shifts, 1979–95

Sample and indicator	*White-collar share of wage bill*	*White-collar share of total employment*
Full sample		
Total share	0.049	0.020
Within-industry share	0.032	0.014
Between-industry share	0.017	0.006
Export-oriented		
Total share	0.027	0.0030
Within-industry share	0.020	0.0026
Between-industry share	0.007	0.0004
Import-competing		
Total share	0.017	0.013
Within-industry share	0.015	0.009
Between-industry share	0.002	0.004
Nontradable		
Total share	0.060	0.036
Within-industry share	0.024	0.027
Between-industry share	0.036	0.009

Source: Authors' calculations.

For the full sample, the increase in the labor share is positive (0.02) and most of this increase is accounted for by within-industry variation rather than between-industry variation, consistent with the notion of skill-biased technological change. These results also hold across sectors, with the largest increase occurring in the nontradable sector (0.036). As noted earlier, we see a much larger increase in the wage-bill share than the employment share for the full sample, though again most of the variation is explained by within-industry movements. For both the export-oriented and import-competing sectors, the within-industry variation explains the largest fraction of the change in the wage bill. In contrast to the other two sectors, however, a substantial fraction of the rise in the nontradable wage-bill share is explained by between-industry variation. With this latter result as a potential exception, these results are broadly consistent with the notion that the relative shift toward skilled workers is due to skill-biased technological change. For the full sample, shifts in the share of skilled workers are mostly explained by within-industry changes, which represent 65–70 percent of the total variation.

3.1 Regression Analysis: Cost-minimization Approach

We now consider a more structural analysis of the determinants of the wage-bill share at the plant level. In the presence of skill-biased technological change, we expect the wage-bill share to be correlated with measures of technology adoption at the plant level. To the extent that capital and skilled labor are complements in the production function, we also expect the wage-bill share to be positively related to capital intensity. This would be true, in particular, if new capital goods embodied new technologies that required highly skilled workers. We analyze the relation between labor composition, technology adoption, and capital intensity using a cost-minimization approach in which capital is assumed to be quasi-fixed and plants minimize the cost of unskilled and skilled labor. We assume constant returns to scale in production and consider a restricted variable translog cost function for plant *i* in year *t*, which results in the following expression for the share of skilled labor in the wage bill:

$$SHARE_{it} = \alpha + \beta \ln\left(\frac{w_{it}^{s}}{w_{it}^{u}}\right) + \gamma \ln\left(\frac{K_{it-1}}{VA_{it}}\right) + \delta TECH_{it} + \varepsilon_{it}\ . \tag{2}$$

In equation (2), ${w_{it}}^{s}$ and ${w_{it}}^{u}$ are wages for skilled and unskilled labor; K_{it-1} is capital, which is pre-determined; and VA_{it} is value-added. The coefficient γ measures the extent to which capital and skilled labor are complements. Plants vary not only in their wage structure and capital intensity, but also in their access to and use of technology. We therefore also include $TECH_{it}$, a vector of observable technology measures, as additional controls in the regression. Equations of this form have been estimated in other studies linking technology changes and employment structure for both developed countries (see Machin and Van Reenen, 1998; Berman, Bound, and Machin, 1998) and developing economies (Pavcnik, 2002b).

We include time, industry, and location dummies to control for unobserved shocks to the relative demand for skilled workers. Industry dummies are constructed using a four-digit industry classification. Given that relative wages are highly endogenous, they are not included in the estimating equation. Rather, relative wages are replaced by industry-specific time dummies.

The equation to be estimated is as follows:

$$\begin{aligned} SHARE_{it} = {} & \alpha + \gamma \ln\left(\frac{K_{it-1}}{VA_{it}}\right) + \delta_1 m_{it} + \delta_2 FTA_{it} + \eta YEAR_t \\ & + \theta LOCATION_i + \mu INDUSTRY_j + \varepsilon_{it} \end{aligned} \tag{3}$$

where FTA_{it} and m_{it} are the proxies for technology use. FTA_{it} measures the share of expenses in foreign technical assistance relative to total expenses in services from third parties, and m_{it} is the share of imported materials in total materials. Pavcnik (2002b) uses both of these measures in her analysis of the wage share over the period 1979–86.[8] If capital is complementary to skilled workers, γ should have a positive sign. Results are presented in tables 11 and 12.

For the full sample, the results in table 11 indicate that capital deepening is related to a higher demand for skilled workers. In particular, the coefficient on the share of capital to value-added is positive and significant. When interpreted as an elasticity, the size of the coefficient is small in absolute terms, however. It is also small relative to the elasticity for the imported materials share, which, in contrast, is of substantial economic significance—on the order of 0.1 to 0.2, depending on the sector.[9] Running the regression for each subgroup according to trade orientation produces mixed results. In the import-competing and nontradable sectors, we find that additional capital induces a higher demand for skilled workers. For the export-oriented sector, the estimated parameter is negative but not significant implying that there is no evidence of capital-skill complementarity in that sector. In summary, although we found evidence of capital-skill complementarity for import-competing and nontradable plants, the effect is estimated to be economically small, which is consistent with other findings for developed economies.

The results in table 11 also indicate that plants that use imported materials and foreign technical assistance have a higher share of skilled workers. All the coefficients are positive and significant for the subgroups as well as for the full sample. A distinction has to be made with respect to the relevance of the foreign technical assistance variable, which according to the results has a significantly stronger effect in the import-competing and export-oriented sectors relative to the nontradable sector (0.25 and 0.26 versus 0.09). At the same time, imported materials have a stronger effect for nontradable plants.

8. Other studies include R&D intensity as an additional control. Our data set does not contain such information, however.

9. It is possible that the imported materials share is a good proxy for the share of imported capital relative to domestic capital goods. If true, our results imply that complementarities between imported capital and skills are an important determinant of the relative demand for skilled workers. Our findings would still imply that the overall capital intensity is not a significant determinant of the relative demand for skilled workers, however.

The results reported in table 11 are consistent with other studies, and we are controlling for unobserved characteristics at the four-digit industry level. To determine the extent to which the effects are robust to allowing for unobserved plant-level heterogeneity, we report estimates that include plant fixed effects in table 12. For the full sample, the estimate of the coefficient on imported materials is again positive and statistically significant, and there is also evidence in favor of capital-skill complementarity. Dissimilar results are found for the different subgroups.

Table 11. Regressions for Skilled Labor Share in Wage Bill[a]

	Sector			
Explanatory variable	*Export-oriented*	*Import-competing*	*Nontradable*	*Full sample*
$\ln(K_{it-1}/VA_{it})$	–0.0007	0.002*	0.005**	0.003**
	(0.003)	(0.001)	(0.001)	(0.0009)
m_{it}	0.122**	0.168**	0.234**	0.180**
	(0.020)	(0.007)	(0.012)	(0.006)
FTA_{it}	0.262**	0.252**	0.088**	0.178**
	(0.043)	(0.032)	(0.023)	(0.019)
Summary statistic				
Adjusted R^2	0.27	0.20	0.56	0.40
No. observations	3,736	16,966	14,687	35,404

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
a. All regressions include time, location, and four-digit industry dummies. Robust standard errors are reported in parentheses.
Source: Authors' calculations.

Table 12. Plant Fixed Effects Regressions for Skilled Labor Share in Wage Bill[a]

	Sector			
Explanatory variable	*Export-oriented*	*Import-competing*	*Nontradable*	*Full sample*
$\ln(K_{it-1}/VA_{it})$	0.003	0.007**	0.0003	0.004**
	(0.003)	(0.001)	(0.002)	(0.001)
m_{it}	0.040*	0.016**	–0.009	0.010**
	(0.022)	(0.006)	(0.010)	(0.005)
FTA_{it}	0.099	0.026	–0.023	0.004
	(0.066)	(0.022)	(0.021)	(0.014)
Summary statistic				
Adjusted R^2	0.68	0.71	0.80	0.76
No. observations	3,736	16,966	14,687	35,404

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
a. All regressions include time and plant indicators. The adjusted R^2 is obtained by including the estimated group effects on the fit of the model. Robust standard errors are reported in parentheses.
Source: Authors' calculations.

The estimated coefficients on the proxies for technology use are no longer positive and significant for all subgroups. After controlling for plant heterogeneity, the import-competing sector still shows a positive relation between capital intensity and skill levels. Again, the coefficient is estimated to be small in magnitude. For the export-oriented sector, only the positive effect of imported materials on skill upgrading remains. The other two coefficients become statistically insignificant. For the nontradable sector, there is also no longer evidence to support either capital-skill complementarity or skill-biased technological change.

We now consider a specification that pools all plants but allows for sectoral-specific interaction effects. This specification also allows us to directly test for differences in coefficients across sectors. We report these results in table 13. The first column is a simple OLS regression; the second column is the within-plant estimator that allows for fixed effects. The base group is the import-competing sector.

The results in table 13 are largely consistent with the results in table 11. Relative to the import-competing sector, the evidence for

Table 13. Determinants of Wage Bill Share with Interaction Effects[a]

Explanatory variable	*Without fixed effects*	*With fixed effects*
$\ln(K_{it-1}/VA_{it})$	0.004**	0.008**
	(0.001)	(0.001)
$\ln(K_{it-1}/VA_{it})$*EXPORT$_{it}$	−0.006**	−0.005*
	(0.003)	(0.003)
$\ln(K_{it-1}/VA_{it})$*NONTRAD$_{it}$	0.000	−0.010**
	(0.001)	(0.002)
m_{it}	0.168**	0.015**
	(0.007)	(0.006)
m_{it*}EXPORT$_{it}$	−0.034*	0.011
	(0.021)	(0.019)
m_{it*}NONTRAD$_{it}$	0.059**	−0.022*
	(0.014)	(0.011)
FTA$_{it}$	0.254**	0.028
	(0.033)	(0.022)
FTA$_{it*}$EXPORT$_{it}$	0.021	0.019
	(0.053)	(0.060)
FTA$_{it*}$NONTRAD$_{it}$	−0.165**	−0.047
	(0.040)	(0.030)
Summary statistic		
Adjusted R^2	0.40	0.76
No. observations	35,404	35,404

* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
a. Robust standard errors are reported in parentheses.
Source: Authors' calculations.

capital-skill complementarity is much weaker in the export-oriented sector, regardless of whether we control for fixed effects. For the nontradable sector, we find no evidence of a differential impact of capital intensity in the regression without fixed effects, but strong evidence when allowing for fixed effects. For the measures of technology adoption, foreign technical assistance has a stronger effect for import-competing relative to nontradable plants. The effect of the share of imported materials is statistically different for the export-oriented sector without controlling for fixed effects. Consistent with the findings in table 12, these differences are substantially muted once we control for fixed effects.

4. Conclusions

This paper has documented the evolution and composition of labor in Chilean manufacturing over the period 1979–95. By sorting the data into export-oriented, import-competing, and nontradable categories, we were able to examine and compare the evolution of labor composition across sectors with different trade orientations. In particular, the share of white-collar workers in total employment is higher in the import-competing sector relative to the export-oriented sector. The average share of skilled labor in total plant employment increased by 8 percent, whereas the average wage-bill share of skilled workers rose by 16 percent during the period 1979–95. Most of the shifts in these two variables took place within industries, which is one of the arguments in favor of skill-biased technical change.

We used a cost-minimization approach to analyze the relation between the share of skilled workers in the wage bill, capital deepening, and technology adoption. Consistent with other findings (Pavnik, 2000b), our results suggest a robust link between technology adoption and the demand for skilled workers measured by the wage-bill share. We also find a statistically significant effect of capital intensity on the relative demand for skilled workers in the import-competing sector. This effect is small in economic terms, however. We find no such link for the export-oriented and nontradable sectors. Overall, these results suggest that skill-biased technological change is transmitted through mechanisms other than capital accumulation. Because one of our technology measures, the share of imported materials used in production, is probably correlated with the share of imported capital used in production, our evidence leaves open the possibility that in a free-trade environment, skill-biased technological change may be transmitted to developing countries through the adoption of new machines imported from abroad. Further research is required to say more on this question.

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