Credibility of Emerging Markets, Foreign Investors' Risk Perceptions, and Capital Flows

Álvaro Aguirre
Andrés Fernández
Şebnem Kalemli-Özcan
editors

Banco Central de Chile / Central Bank of Chile

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Emerging market economies (EMEs) are constantly exposed to shocks that originate in world capital markets, posing serious challenges to policymakers. By dealing with these shocks —Covid-19 representing the most recent event— several lessons have been learned in terms of the ways they propagate as well as the various tradeoffs of policy responses available. Credibility and foreign investors' risk perceptions are central when analyzing these episodes, and they are closely associated with the design of monetary and fiscal frameworks, as well as the conduct of unconventional policies. This volume contributes to the study of these issues by focusing on the understanding of the array of challenges and policy options for EMEs' policymakers for short-run stabilization purposes as well as longer-term issues that should be on their radars, bringing together a multinational group of distinguished scholars to discuss the latest research findings.

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Table of Contents

Credibility of Emerging Markets, Foreign Investors' Risk Perceptions, and Capital Flows: An Overview
Álvaro Aguirre, Andrés Fernández, and Şebnem Kalemli-Özcan 1

The International Financial System after Covid-19
Maurice Obstfeld 9

Exchange Rate Puzzles and Policies
Oleg Itskhoki and Dmitry Mukhin 47

International Risk Spillovers: Implications for Emerging Markets' Monetary Policy Frameworks with an Application to Chile
Şebnem Kalemli-Özcan 97

Global Drivers and Macroeconomic Volatility in EMEs: A Dynamic-Factor, General- Equilibrium Perspective
Gent Bajraj, Andrés Fernández, Miguel Fuentes, Benjamin García, Jorge Lorca, Manuel Paillacar, and Juan M. Wlasiuk 111

Sovereign-Debt Crises and Floating-Rate Bonds
Mark Aguiar, Manuel Amador, and Ricardo Alves Monteiro 159

*KF*star and Portfolio Inflows: A Focus on Latin America
John D. Burger, Francis E. Warnock, and Veronica Cacdac Warnock 185

How Important Is The Commodity Supercycle?
Andrés Fernández, Stephanie Schmitt- Grohé, and Martín Uribe 205

Cross-Border Corporate Control: Openness and Tax Havens
Gur Aminadav, Luís Fonseca, and Elias Papaioannou 249

The Reversal Problem: Development Going Backwards
Eduardo Olaberria and Carmen M. Reinhart 301

Credibility of Emerging Markets, Foreign Investors' Risk Perceptions, and Capital Flows: an Overview

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Since the work of economist Carlos F. Diaz-Alejandro in the early 1980s, it is well known that emerging market economies (EMEs) are constantly impacted by shocks that emanate from world capital markets, posing many challenges for policymakers in these economies. Examples range from the debt crisis in the 1980s and the Asian and Russian crises of the late 1990s—both having as epicenter EMEs—, to the Global Financial Crisis (GFC) that originated in the United States and, more recently, the Covid-19 crisis and the inflationary consequences of the recovery amid raising interest rates in advanced economies (AEs).

Throughout these nearly half century of increasing financial and trade integration by EMEs with the rest of the world, several lessons have been learned in terms of the ways in which these shocks propagate as well as the various tradeoffs of the policy responses available. The role of credibility is central when dealing with international capital markets. Most EMEs have been able to gradually improve along these dimensions by enhancing their monetary and fiscal frameworks. The role of unconventional policies has also begun taking a predominant role.

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Initially implemented by EMEs in the 1990s, unconventional policies took center stage at the onset of the GFC. They proved to be very effective also during Covid-19 across both AEs and EMEs.

In parallel to these developments, academic research has helped guide policymakers in navigating these challenges. Such growing body of research has documented how international capital market frictions have become markedly more prominent together with the dollar dominance, despite the larger share of EMEs in the world economy, with sharp implications for the conduct of monetary and exchange rate policies. The legacy of the GFC also led many to study the factors that trigger sovereign default crises and what policymakers can do to prevent them or at least mitigate their macro impact.

Despite the prominence of stabilization policies, it is important not to lose sight of long-term growth and its determinants. There has been much progress in advancing economic and social indicators in EMEs. However, there is still lots of ground to cover, and a possible stagnation or even a reversal of the achievements made thus far—as the data seems to indicate—is something that ought to be in every policymaker's radar, no matter how relevant the short-run stabilization concerns are. Surely, a stable macroeconomic environment is a necessary—but not sufficient—condition for growth. Understanding the long-term determinants of capital flows is a must. Likewise, a good understanding of the process of globalization at the corporate level and its ramifications for the transfer of knowledge is key.

It is with these considerations that the volume is divided into two blocks of papers. A first block gathers works aimed at understanding the array of challenges and policy options for EMEs' policymakers for short-run stabilization purposes. A second block deals with longer term issues that should be in EME's policymakers' radars, along the lines mentioned above.

The first block begins with the chapter **The International Financial System after Covid-19**, by Maurice Obstfeld. The author points out that, once again, the world economy appeared on the brink of collapse—until it was pulled back by monetary and fiscal interventions that outstripped even those of the 2008–2009 GFC. The Covid-19 crisis originated in a totally different type of shock—one coming exogenously from outside the financial system rather than from within—, and it provided a kind of stress test for the amended international financial system. But a collapse in 2020 was averted only thanks to unprecedented policy support, previously unthinkable in magnitude and scope, which would be rash to rely on for the future.

The paper reviews the evolution of global financial markets since the GFC, changes in academic thinking about the domestic impacts of these markets, the strains seen during the Covid-19 crisis, and perils that may lie ahead. A key theme is that stability will be enhanced if the global community embraces reforms that elevate market resilience, rather than depending on skillful policymakers wielding aggressive but ad hoc policy interventions to come to the rescue again.

In the second chapter, titled **Exchange Rate Puzzles and Policies**, Oleg Itskhoki and Dmitry Mukhin zoom in on the problem of designing optimal exchange rate regimes, providing a fresh look at the long-standing debate of benefits and costs of pegged vs managed or free-floating regimes. Indeed, a perennial question in the mind of EMEs' practitioners relates to the extent to which their economies face a trilemma constraint in choosing between inflation and exchange rate stabilization, unlike the divine coincidence observed in closed economies. The authors address these questions by developing a general equilibrium policy analysis framework with nominal rigidities and financial frictions that are both central for equilibrium exchange rate determination and result in an empirically realistic model of exchange rates. Building on their earlier work, the new model is consistent with the exchange rate disconnect properties across floating and fixed regimes allowing for explicit policy analysis, using both monetary policy and foreign exchange (FX) interventions in the financial market. The model features a Balassa-Samuelson mechanism determining the value of the frictionless real exchange rate (departures from purchasing power parity, PPP) and segmented financial markets resulting in endogenous equilibrium, and Uncovered Interest Rate Parity (UIP) deviations. The presence of both endogenous PPP and UIP deviations is essential for the optimal exchange rate policy analysis, as exchange rate variation is at the core of both deviations. A key takeaway is that, within this environment, FX interventions surface as a valuable stabilization tool against the costly UIP deviations.

The third chapter, on **International Risk Spillovers: Implications for Emerging Markets' Monetary Policy Frameworks with an Application to Chile**, by Şebnem Kalemli-Özcan looks closely at international risk spillovers to EMEs and their implications for the design of their monetary policy frameworks. Among the factors behind international spillovers, U.S. monetary policy developments retain a major influence. Such developments drive the global financial cycle and shape global investors' risk perceptions. Drawing from her earlier work, she shows that the transmission mechanism for monetary policy spillovers to emerging market economies rests on the effect of U.S. monetary policy on investors' risk sentiments, as those sentiments are more volatile in the case of EMEs, and that capital flows to emerging markets are particularly "risk-sensitive," creating a challenge unique to the EME policymakers and their monetary policy frameworks. A clear policy advise follows: EME policymakers should smooth out this risk sensitivity by deciding not to use monetary policy rates but other policy tools instead. A good barometer of this risk sensitivity is the UIP risk premia and, if EME policymakers use policy rate to respond to U.S. monetary policy changes, the UIP risk premia increase further. The case for flexible exchange rates is stronger under international risk spillovers, since floating exchange rates help to smooth out the UIP risk premia, thus freeing domestic monetary policy's hand to focus on inflation targeting and output stabilization. Countries may want to limit exchange-rate volatility because of the negative effects of excessive volatility on balance sheets due to extensive debt denominated in foreign currency and/or a high degree of passthrough of currency depreciations to inflation. For "fear of floating" linked to foreign-currency debt, countries can limit the extent of foreign-currency debt by using countercyclical prudential policies. Macroprudential and capital-flow management policies can be used countercyclically in a transitory way, to limit unhedged foreign-currency-denominated liabilities not only in the financial sector, as typically done, but also in the nonfinancial corporate sector.

The fourth chapter, on **Global Drivers and Macroeconomic Volatility in EMEs: A Dynamic-Factor, General-Equilibrium Perspective**, by Gent Bajraj, Andrés Fernández, Miguel Fuentes, Benjamín García, Jorge Lorca, Manuel Paillacar, and Juan Marcos Wlasiuk complements this strand of literature by exploring the nature of the global forces that impact EMEs as well as their transmission mechanism. Motivated by various influential works in the literature, they consider a wide array of forces and their interlinkages, from a global financial cycle to fluctuations in commodity prices and a common growth factor. One interesting finding of their work is the preponderance of the financial factor affecting the other two. It is also noteworthy that jointly, the three factors account for more than a third of the variance in GDP in a pool of 12 EMEs. Then, to better understand how these global forces are transmitted into EMEs, they zoom in on Chile and augment a large-scale DSGE regularly used for policy analysis with the estimated global dynamic factor structure.

This allows them to document the general equilibrium channels through which shocks in these global factors are transmitted into the business cycle and, in turn, the policy challenges that they entail. Their findings indicate a preponderant role of global drivers for EMEs' business cycles, with a third of their macro variability being traced back to shocks in global dynamic factors. While the global financial cycle is a relevant force, a factor associated to global prices and commodities appears equally important, with a relatively modest role played by pure growth/productivity forces. Further, although some of the ensuing effects of shocks to the financial cycle offset each other, the opposite occurs when a shock to global prices materializes, calling for a more active monetary policy response.

In extreme cases global forces, combined with domestic factors, may drive EMEs to default on their debt. This is addressed in the fifth and last chapter of the fist block by Mark Aguiar, Manuel Amador, and Ricardo Alves Monteiro in **Sovereign-Debt Crises and Floating-Rate Bonds**, which looks at sovereign-debt crises and the policy options related to them. They argue that the choice of sovereign-debt maturity in countries at risk of default represents a complex set of competing forces. An interesting policy recommendation that the authors stress is that long-term bonds may be a useful tool for a government to hedge shocks to the cost of funds arising from business cycle fluctuations. They show that having a coupon on a long-term bond indexed to one-period-ahead default probabilities provides all the incentive properties of one-period bonds, without the vulnerability to rollover risk. In terms of implementation, the authors argue that such policy can be implemented by indexing the coupon to the auction price of a small amount of one-period bonds.

The second block, which deals with longer term issues, begins with the sixth chapter by John D. Burger, Francis E. Warnock, and Veronica Cacdac Warnock, entitled **KFstar and Portfolio Inflows: A Focus on Latin America**. Their work studies the natural level of capital flows, denoted by KF*, as a useful tool available to policymakers faced with volatile capital flows who may desire a method to identify the level of flows likely to persist in the medium run. KF* is a supply-side measure in that it is derived from the supply of rest-of-the-world savings, and it is computed by using lagged portfolio weights from portfolio liabilities data multiplied by current rest-of-the-world savings. In that sense, KF* is an easy-to-construct, slow-moving supply-side benchmark derived from the supply of rest-of-the-world savings approximating the level of flows to which they should converge over a medium-term horizon.

In an interesting application to Latin American portfolio inflows, the authors show how deviations from KF* help predict sudden stops in the region. Furthermore, they document the ability of KF* to act as an indicator of vulnerability in the face of global shocks. Case studies of the Global Financial Crisis (GFC), the post-GFC surge, and the Covid-19 pandemic each indicate that, for Latin American countries, KF* provides useful real-time information on the vulnerability of flows. Finally, they analyze the drivers of short-run deviations of flows from KF* and document interesting heterogeneity as flows to Brazil, Chile, and Mexico appear closely linked to the commodity prices, while flows to Argentina, Costa Rica, and Peru are linked to global risk tolerance.

Another source of medium-to low-frequency variability for various EMEs is the so-called "commodity supercycle". This is explored in the seventh chapter, entitled **How Important is the Commodity Supercycle?** by Andrés Fernández, Stephanie Schmitt-Grohé, and Martín Uribe, who identify global disturbances that cause regular cycles and supercycles in world commodity prices and estimate their contribution to aggregate fluctuations across many emerging and developed countries. The commodity price supercycle, which has a periodicity of 20 to 30 years, is identified as a common permanent (nonstationary) component in all commodity prices. An important advantage of the method applied in this work is that it allows for the joint estimation of transitory and permanent domestic and world disturbances affecting aggregate activity in individual countries. The authors show that the common permanent component explains on average across commodities between 67 and 91 percent of the forecast error variance of commodity prices at horizons between five and thirty years. Estimates using quarterly and annual data from 1960 to 2018 indicate that world shocks that affect commodity prices and the world interest rate are major drivers of aggregate fluctuations in developed and emerging small open economies, explaining more than half of the variance of output growth on average across countries. However, most of this contribution (i.e., more than two thirds) stems from stationary world shocks, even at forecasting horizons typically associated with the supercycle. These results suggest that world disturbances that are responsible for the commodity price supercycle did not play a dominant role in driving fluctuations in aggregate activity at the country level.

A potentially important driver of growth for most EMEs relates to the ability of firms and corporations to insert themselves into the world economy. This can happen through various channels, one of which is cross-border corporate control. A first step towards a better understanding of this topic is done in the eighth chapter entitled **Cross-Border Corporate Control: Openness and Tax Havens** by Gur Aminadav, Luís Fonseca, and Elias Papaioannou, documenting some broad patterns, based on ongoing research of the drivers of the internationalization of corporate control. Until now, this has not been well-understood due to the esoteric corporate holding schemes and the complex network of equity holdings. By compiling new ownership data for almost 90 percent of the world market capitalization of listed firms in 2012, the authors provide a mapping of corporate control, zooming into the role of tax havens. Their descriptive analysis reveals considerable cross-country heterogeneity in the openness to foreign control and in the usage of tax-haven intermediate and controlling entities. Foreign control is more frequent in smaller and less developed countries, echoing the international trade and portfolio investment evidence. Residents and entities of richer countries exert a larger portion of their controlling equity stakes in firms abroad. The authors also detect that a sizable portion of controlling equity stakes is held by or via tax-haven-incorporated entities, particularly in lower-income economies. Tax-haven use is the highest in Eastern Europe (Bulgaria, Ukraine, Serbia, Latvia, Russia), Africa (Ghana, Nigeria, Ivory Coast), and some East Asian countries like Indonesia and Pakistan.

Completing the volume, the final chapter **The Reversal Problem: Development Going Backwards** by Eduardo Olaberria and Carmen M. Reinhart, identifies a setback in development for emerging and developing countries that appeared around half a decade earlier than the huge economic downturn triggered by the Covid-19 pandemic, which, on the other hand, deepened and accelerated the process. The authors argue that the turning point occurred around 2015, and it was marked by the sharpest decline in commodity prices since the early 1980s. They document these trends with an encompassing array of indicators, including poverty rates, inequality, human capital, and democratic values. Their analysis shows that the Reversal Problem is widespread in terms of geography, though more acute in Africa and other low-income regions. It is also encompassing in that it not only affects economic and social indicators but also has the potential to cause political instability both domestically and across borders, as these trends appear to coincide with a setback in democratic values. They conclude by arguing that, while the extent of the development Reversal Problem is not yet on the development reversals scale of 40 years ago, there are added new risks posed by climate change that will disproportionally harm emerging and developing economies.

The International Financial System after Covid-19

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In March of 2020, international markets seized up with a violence unequaled since the Global Financial Crisis (GFC) nearly a dozen years before. As economies around the world locked down in the face of the potentially deadly but completely novel SARS-CoV-2 virus, stock markets fell, firms and governments scrambled for cash, liquidity strains emerged even in the market for U.S. Treasurys, and capital flows to emerging and developing economies (EMDEs) reversed violently. Once again, the world economy appeared on the brink of collapse—until it was pulled back by monetary and fiscal interventions that outstripped even those of the 2008–2009 Global Financial Crisis.

The GFC erupted after five years of global financial-market expansion following the Asian crisis of the late 1990s, the dot.com collapse and Enron corporate fraud scandal, and the 9/11 attack on the United States. Following the GFC, macroeconomists questioned their earlier theoretical paradigms, financial firms altered their business

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models, and regulators rethought their oversight frameworks. Those paradigms, models, and frameworks needed to change: they had complemented each other in allowing the most severe financial crisis since the 1930s.

The Covid-19 crisis originated in a totally different type of shock—one coming exogenously from outside the financial system rather than from within—and it provided a kind of stress test for the amended international financial system. So far the system has survived tolerably well, even in the face of a global public-health response that has underperformed on many levels. But a collapse in 2020 was avoided only thanks to unprecedented policy support, previously unthinkable in magnitude and scope, which it would be rash to rely on for the future. And now, support is being withdrawn.

This paper reviews the evolution of global financial markets since the GFC, changes in academic thinking about the domestic impacts of these markets, the strains seen during the Covid-19 crisis, and perils that may lie ahead. A key theme is that stability will be enhanced if the global community embraces reforms that elevate market resilience, rather than depending on skillful policymakers wielding aggressive but ad hoc policy interventions to ride to the rescue again. Next time could be different—and not in a good way.

The plan of this paper is as follows. Section 1 surveys trends in financial market activity since the GFC, focusing especially on the huge demands that the Covid-19 shock placed on markets. Section 2 reviews the emerging evidence that global asset and commodity prices, capital flows, and intermediary leverage are driven by a global financial cycle linked to U.S. monetary policy. Section 3 summarizes measures central banks took to counteract the effects of the Covid-19 shock, focusing on the case of the Republic of Korea. For EMDE central banks, the episode stood in sharp contrast to earlier crises, in which their authorities sometimes felt forced to react procyclically. But it is too early to argue that EMDEs have entered a new world of copious policy space. Section 4 argues that with advanced economies defeating the pandemic more quickly than EMDEs, the world is having an uneven rebound in which lagging and more indebted EMDEs are likely to be hit by a contracting global financial cycle, driving them into liquidity or solvency crises.

That potential scenario is just one threat to financial stability that the Covid-19 crisis has highlighted. Accordingly, section 5 outlines several areas where reforms at both the global and national levels could improve the resilience of international financial markets.

1. Trends in International Financial Markets

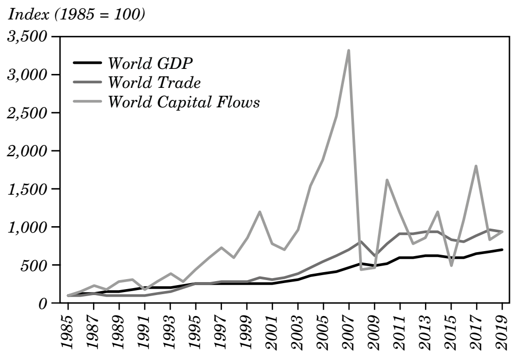
Starting in the 1990s, the scale and scope of global financial markets exploded. Eventually, additional financial vulnerabilities owing to massive and largely unregulated cross-border financial flows came to outweigh incremental gains from asset exchange, resulting in the global financial distress of 2008–09. Figure 1 shows an index of global capital flows since the mid-1980s. By the mid-1990s, growth in international financial transactions outstripped that in output or trade, even as the growth in the latter was amplified in the first decade of the new millennium by the proliferation of global value chains. The extreme bulge in capital flows in that same decade cannot be explained by a sudden rise in opportunities for mutually advantageous, socially beneficial asset trade. Instead, it reflected market distortions that came to tears before the end of the decade. Since the Global Financial Crisis, international capital flows have fluctuated wildly in response to various shocks, though never again reaching their earlier 2007 peak. Korea has not been immune to these capital-account surges and stops.

Key to these developments has been the regulatory regime around international financial flows: the set of guardrails governments maintain to manage the volume and character of cross-border finance, as well as its uses within the domestic financial system. Figure 2 reports the Chinn-Ito (2006) measure of financial account openness, updated to 2018. This index is a de jure measure, which codes the level of official restrictions as reported by the IMF, as opposed to a de facto index of actual international capital movements. After the early 1990s, high-income countries quickly removed remaining restrictions, approaching maximum levels of financial openness by the early 2000s.[1] Like other high-income countries, Korea has for several years been characterized by nearly complete de jure financial openness. Lower-income countries also began a liberalization process around the early 1990s, but it has been slower and has remained incomplete, even backtracking slightly after the Global Financial Crisis. Accordingly, flows between advanced economies account for the bulk of the early-millennium surge seen in figure 1. In general, middle- and low-income countries with current-account surpluses invest them in advanced markets, which then recycle them to developing markets with current-account financing needs. However, in the past two decades, the volume

1. For a discussion of this process, see Obstfeld (2021a).

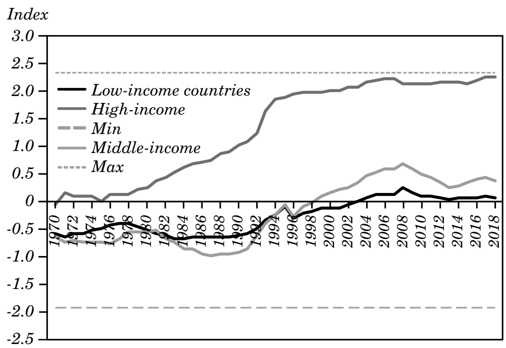
of direct flows between developing markets has risen[2]), also supporting rising gross levels of external assets and liabilities on the part of the less prosperous economies.

Figure 1. Comparing the Growth of World GDP, World Trade, and World Capital Flows
(nominal U.S. dollars, all series rebased to 1985 = 100)



Sources: IMF, *World Economic Outlook* database, April 2021, IMF Balance of Payments and International Investment Position Statistics database, and UN Comtrade database. World trade is measured as world imports.

Figure 2. Chinn-Ito Index of Financial Account Restrictions, 1970–2018
(simple country-group averages)

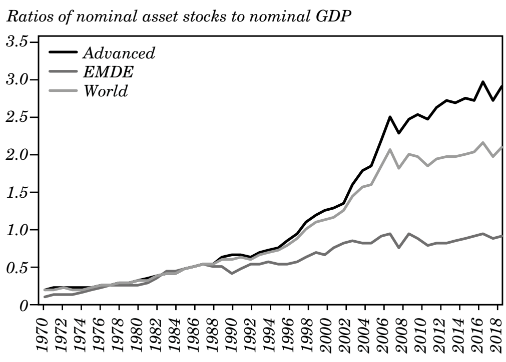


Source: Chinn and Ito (2006) data, updated by authors through 2018.

2. See Broner and others (2020) and CGFS (2021).

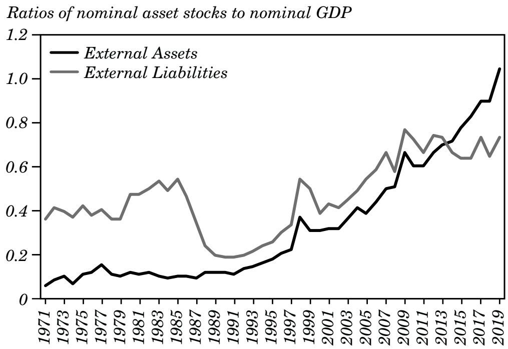
A country's level of gross external assets and liabilities relative to GDP furnishes one possible de facto measure of international financial integration. For the EMDE-country groups, figure 3 shows the average of external assets and liabilities as a ratio to GDP, based on the data of Lane and Milesi-Ferretti (2018). The rapid run-up in advanced economies, starting in the early 1990s but slowing sharply after the Global Financial Crisis, is evident and quite consistent with figure 2. The very high numbers (recently around three times GDP) reflect in part the extreme sizes of the balance sheets of financial centers, including offshore havens. Also consistent with figure 2, EMDEs show a less extreme (though still pronounced) increase after the early 1990s. However, that trend has pretty much stalled relative to the de facto openness levels reached just before 2008, in contrast to the continuing slow rise seen for advanced economies. Figure 4 shows the external assets and liabilities of Korea, also as a share of GDP. The magnitudes are similar to those for the EMDE grouping in figure 3. In Korea's case, however, while the growth of gross external liabilities (relative to output) has stalled since the Global Financial Crisis, external assets have continued to grow, consistent with Korea's ongoing current-account surpluses (which in 2015 reached 7.2 percent of GDP, falling to a still substantial 4.6 percent in 2020).

Figure 3. Ratios of External Financial Exposure to GDP for Advanced Economies and EMDEs, 1970–2019
(average of gross external assets and liabilities)



Source: Lane and Milesi-Ferretti (2018) data, updated by authors through 2019.

Figure 4. Korea: Ratios of External Assets and Liabilities to GDP, 1971–2019



Source: Lane and Milesi-Ferretti (2018) data, updated by authors through 2019.

Extreme as they may seem compared with world trade, the capital-flow numbers graphed in figure 1 far understate true gross levels of international transactions in financial instruments. To see why, note that figure 1 shows the sum of all countries' capital (or financial) inflows (which equals the sum of global capital outflows apart from errors and omissions in the official data). By definition, a country's capital (or financial) inflow equals foreign purchases of assets issued by domestic residents less foreign sales of assets issued by domestic residents, that is, net foreign purchases of domestic assets. Capital outflows are defined analogously as domestic residents' purchases of foreign assets less their sales of the same. However, reported capital inflows and outflows—often referred to as 'gross' capital flows because their difference is the *net capital inflow* or current-account deficit (again, apart from errors and omissions)—actually are themselves the result of netting the purchases and sales carried out on the same period by a particular set of actors. In principle, such 'gross' capital flows thus understate the absolute levels of two-way flows.[3] To get an accurate assessment, we need the gross 'gross' numbers, that is, purchases and sales of domestic and foreign residents before netting.

Such data are hard to come by, but at least for the United States, we can calculate a workable lower bound from the U.S. Treasury's Treasury International Capital (TIC) System data and compare

3. See Koepke and Paetzold (2020).

those numbers both with the net capital flow required to offset the current account and the conventionally defined gross capital inflow and outflow. The TIC data are monthly and report:

(a) Gross U.S. resident sales to foreign residents of U.S. stocks and U.S. long-term bonds (for example, excluding Treasury bills, but including long-term corporate bonds). These necessarily equal foreign purchases of the U.S. assets.

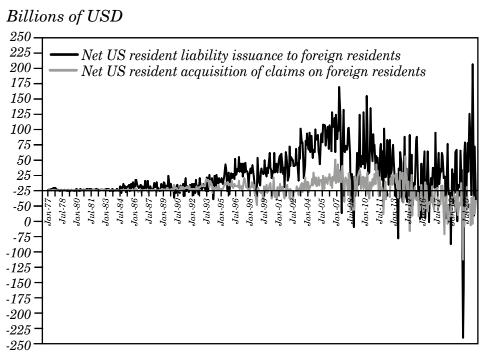
(b) Gross U.S. resident purchases of U.S. stocks and U.S. long-term bonds from foreign residents. These necessarily equal foreign sales of the U.S. assets.

(c) Gross U.S. resident purchases of foreign stocks and bonds from foreign residents.

(d) Gross U.S. resident sales of foreign stocks and bonds to foreign residents.

These data therefore capture much of portfolio capital flows; they exclude, in addition to transactions in short-maturity U.S. Treasury bills, foreign direct investment flows, and flows of bank loans. In conventional balance-of-payments accounting, U.S. capital inflows relate closely to (a) less (b), whereas U.S. capital outflows relate closely to (c) less (d).

Figure 5. U.S. Conventional 'Gross' Monthly Long-Term Portfolio Inflows and Outflows



Source: U.S. Treasury, Treasury International Capital System, Monthly Transactions in Long-term Securities.

Figure 5 graphs these two proxies for the U.S. 'gross' capital inflow and outflow. In terms of overall magnitude, the absolute values of the series stay below USD 250 billion, which is just slightly more than 1 percent of projected 2022 annual U.S. GDP. Because these are monthly flows and not expressed at an annual rate, however, the correct comparison is with one-twelfth of annual GDP. So we are looking at monthly inflows and outflows that can be on the order of 10 percent of GDP. If the TIC data offered a comprehensive picture of U.S. international financial flows, the U.S. current-account deficit would equal the difference between capital inflows (a) less (b) and capital outflows (c) less (d).[4] The deficit was about three percent of GDP over 2020—roughly one-third the magnitude of 'gross' capital inflows and outflows. Also notable in figure 5 are the abrupt contractions in international positions—with foreign residents selling U.S. assets and U.S. residents selling foreign assets—around the Lehman shock in 2008 (see figure 3) and the Covid-19 shock in the early spring of 2020. U.S. recovery and fiscal stimulus early in 2021 bring a surge of capital inflows.

Figure 6 graphs the true gross capital-account transactions (gross 'gross' flows)—the sales and purchases considered separately. Often these may be legs of a single transaction, corresponding to offsetting bookkeeping entries in the balance of payments, but nonetheless, the magnitudes of transaction volumes are breathtaking.[5] The numbers have tended to grow over time, falling after Lehman but then rising back up and reaching very high levels in the volatile market conditions of the Covid-19 crisis. Transaction volumes for U.S. long-term assets have recently approached USD 7 trillion per month, which would exceed monthly U.S. GDP by a factor between three and four (and these numbers exclude trade in short-term assets.) One interesting (if unsurprising) feature of the data is that in trades involving U.S. residents, transaction volumes for U.S. assets are consistently much higher than those for foreign assets. This is a reflection of continuing "home bias" by U.S. residents, of the outsized role of the dollar in global

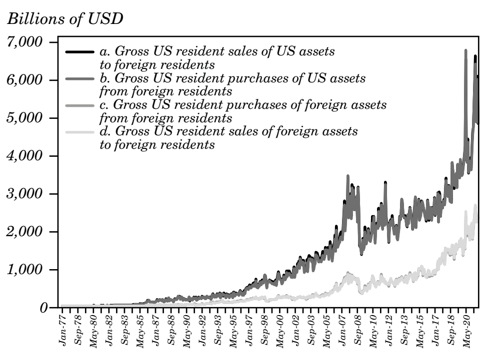
4. Thus, if the financial flow data were comprehensive, the current-account deficit would also equal (a) + (d) – [(b) + (c)]: gross U.S. resident sales of all assets to foreigners (whether claims on the United States or a foreign country) less gross U.S. resident purchases of all assets from foreigners.

5. That is why the series are so highly correlated. Suppose a foreign resident holder of a U.S. brokerage account shifts from U.S. bonds to U.S. stocks. The U.S. is selling them a stock but buying back a bond in payment. The trade gives rise to offsetting items in category (a) and (b) above, with no net impact on U.S. capital inflows (a) – (b).

financial markets, and of the United States' big net debtor position. Moreover, the gap between transaction volume in U.S. assets and in foreign assets appears to be secularly widening over the 2000s.

Net capital flows (the current account) matter as a component of aggregate demand. Conventionally defined gross capital flows matter as a measure of the net global demand for country assets. A general collapse in gross flows may signal a global risk-off episode, while a collapse in gross inflows (a sudden stop) can leave an economy with depressed asset prices as well as an inability to pay maturing debts.[6] The enormous volume of truly gross two-way asset trade indicates how small are the asymmetric proportional changes that can potentially spark crises. The same is true of foreign portfolio shifts between a given country's asset classes. Such shocks could be amplified if the financial system's plumbing leads to liquidity shortages, fire sales, failed settlements, or other dysfunction. The volume of global financial transactions seems disproportional to any fundamental economic need or activity, yet produces a system prone to fragility.[7] Like the Global Financial Crisis, the Covid-19 shock in the spring of 2020 illustrated the need for massive central-bank intervention as a backstop to market stability.

Figure 6. Gross U.S. Resident Monthly Long-Term Portfolio Asset Sales to and Purchases from Foreign Residents



Source: U.S. Treasury, Treasury International Capital System, Monthly Transactions in Long-term Securities.

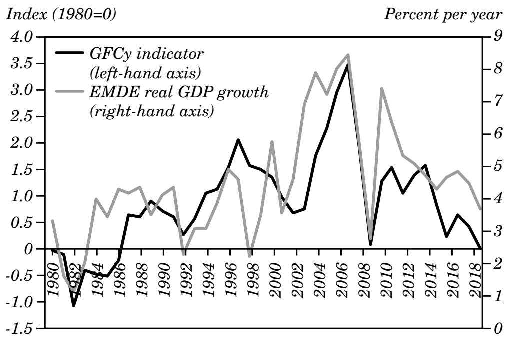
6. See Forbes and Warnock (2012).

7. Trading levels may be socially inefficient for several reasons, among them: tax arbitrage schemes or money-laundering motives, investor overconfidence (Odean, 1999), externalities from liquidity management (He and Kondor, 2016), or the design of fund managers' incentive contracts (Kashyap and others, 2020).

2. Global Cycles in Asset Prices, Commodity Prices, and Asset Flows

The last section described the distinct upward trends in international financial integration and transaction volumes. But what forces underlie the fluctuations around trend that the data also show? Recent research points to a pattern of synchronized international movements in financial conditions such that asset prices, commodity prices, capital flows, and intermediary leverage tend to surge and ebb together across a range of national markets (Miranda-Agrippino and Rey, 2021). Given the central role of U.S. financial markets and the dollar in global markets, U.S. financial conditions and Federal Reserve monetary policy are key drivers of the global cycle. Financial conditions and monetary policies in other developed markets also play roles, and global fluctuations in risk aversion certainly correlate with the cycle, partly as cause and partly as effect. Figure 1 suggests a cyclical behavior in global capital flows, most notably in the run-up to the Global Financial Crisis.[8]

Figure 7. Growth in Emerging and Developing Economies and the Global Financial Cycle



Source: *GFCy* variable with data updated through 2019 is available at http://silviamirandaagrippino.com/code-data. The raw monthly data are averaged to derive annual observations. Real GDP growth is from IMF, World Economic Outlook database, April 2020.

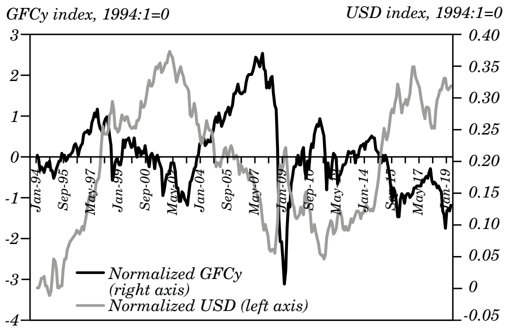
8. There is disagreement over the scope of the global financial cycle. For example, Cerutti and others (2019) argue that the cycle encompasses asset prices but not capital flows.

For countries with some degree of integration into world markets, these cycles reflect global financial-market impulses with potentially powerful effects on exchange rates, growth, prices, and financial stability. Researchers have therefore sought to measure the global financial cycle and to ascertain its effects and the variables that drive it.

Miranda-Agrippino and Rey (2020) use a monthly dynamic factor model of equity, bond, and commodity prices spanning five continents to estimate a single-global factor accounting for 20 percent of the common variance of the asset prices. Scheubel and others (2019) develop alternative measures based on a latent factor model that includes not only asset prices, but also non-price indicators including portfolio inflows to EMDEs, global credit volume, and the leverage of broker-dealers. Davis and others (2021) apply a related approach to explain net and gross capital flows (gross being defined in the conventional sense). They find that two factors, a global financial cycle factor and an energy-price factor, have high explanatory power for gross and net flows across advanced economies and EMDEs. Both the Scheubel-Stracca-Tille factor and the Davis-Valente-van Wincoop financial factor correlate well with the factor of Miranda-Agrippino and Rey, which I denote by *GFCy*.

Figure 7 illustrates the close relationship between the global financial cycle index GFCy and real output growth in EMDEs, which are especially vulnerable to the vicissitudes of international capital flows. For the annual data in the figure, changes in EMDE growth rates track broadly the swings in *GFCy*.

Figure 8. *GFCy* Index versus BIS Broad Nominal Dollar Index



Source: *GFCy* variable with data updated through 2019 is available at http://silviamirandaagrippino.com/code-data. Exchange-rate data from Bank for International Settlements, available at https://www.bis.org/statistics/eer.htm.

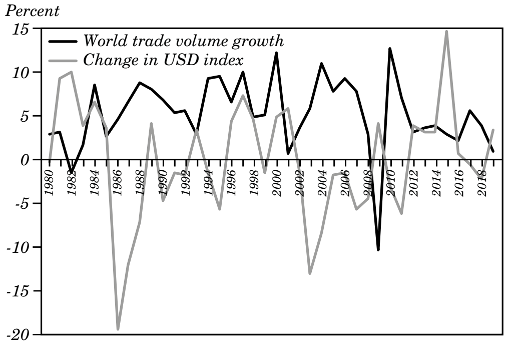
Several studies identify the U.S. dollar nominal effective exchange rate as a bellwether for global financial conditions, operating through international banking activity—as in Bruno and Shin (2015) and Shin (2019)—and possibly other channels. This association likely reflects, in part, the impact of U.S. monetary-policy shocks on the dollar exchange rate, restrictive policy implying dollar appreciation and tighter lending conditions. In this case, cross-border U.S. dollar flows will react most strongly, yielding an especially powerful negative impulse given the dollar's centrality in cross-border transactions.

Using a vector-autoregression framework, Miranda-Agrippino and Rey (2020) show how alternative measures of U.S. contractionary monetary-policy shocks induce dollar appreciation, falls in financial intermediary leverage, credit, and banking flows, and a decline in the global cycle index GFCy. As to the mechanisms at work, Cesa-Bianchi and others (2018) present evidence to support a model in which currency and house-price appreciation inflate collateral values, thereby amplifying the expansionary effect of capital inflows. The association could also reflect dynamics in which causality flows from exogenous shifts in global risk appetite into simultaneous movements of the dollar (through a safe-haven effect) and the global asset prices that underpin *GFCy*.

Looking at the data from 30,000 feet, the unconditional negative correlation between the dollar's strength and the Miranda-Agrippino and Rey global cycle factor is striking. Figure 8 shows the relationship since 1994: the correlation coefficient between the two monthly series is –0.35. More impressive than the negative month-to-month correlation, however, is the strong negative relationship between low-frequency swings in the series. The figure thus suggests that the dollar foreign exchange value is indeed a powerful inverse indicator of the global financial cycle.[9]

9. Figure 8 should be interpreted with caution, as the GFCy index is based on asset prices measured in dollars. However, Miranda-Agrippino and Rey (2020, online appendix) state that its general behavior is robust to estimation based on assets' local-currency prices.

Figure 9. U.S. Dollar Appreciation Correlates with Lower Growth in the Volume of World Trade



Source: International Monetary Fund, World Economic Outlook database, April 2021, trade volume of goods and services; FRED, dollar exchange-rate series TWEXMANL, trade-weighted based on goods trade with major-currency trading partners (Euro area, Canada, Japan, United Kingdom, Switzerland, Australia, and Sweden).

The mechanisms linking the dollar and the cycle affect EMDEs with special force, which helps to explain figure 7. One factor is the prevalence of foreign-currency borrowing in some countries, which implies that a depreciation of local currency against the dollar will batter domestic balance sheets with contractionary macro effects. Even where sovereigns have largely graduated to domestic-currency borrowing and banks avoid currency mismatch, duration mismatches in foreign currency matter, and EM corporates borrow extensively in foreign currency. Moreover, foreign holders of domestic-currency debts may be especially sensitive to prospective exchange-rate movements, creating outsized capital-flow responses that can destabilize domestic financial markets unless the domestic investor base is deep (Carstens and Shin, 2019). Two additional mechanisms follow from the dollar's impact on global trade and commodity prices.

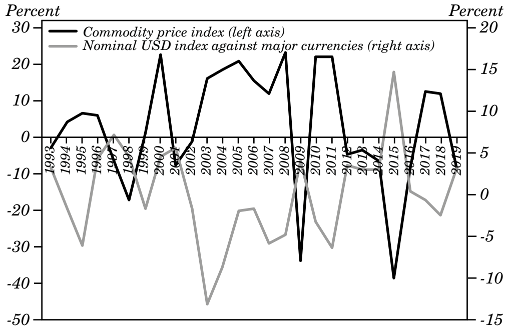
A striking relationship in the data is the strong negative association between nominal dollar appreciation and world trade volume. Figure 9 shows this relationship in annual data from 1980. This relationship is not fully understood, but likely owes to at least five primary (and complementary) mechanisms. First is a direct effect of dollar-induced financial tightening, operating through the need for trade finance credit. This effect has likely become stronger with the proliferation of global value chains since the 1990s (Bruno and Shin, 2021). A second potential mechanism works through the dollar's safe-haven tendency to strengthen in global crises, when risk appetite falls sharply as trade contracts. A third mechanism would be a contractionary effect of a stronger dollar on export demand when export prices are invoiced in U.S. dollars and sticky.[10] Gopinath and others, (2020) show how dollar appreciation reduces ex-U.S. world merchandise export growth, even controlling for global GDP growth and risk aversion (as proxied by the VIX). A fourth possible mechanism is a global decline in investment when the dollar strengthens and funding conditions tighten, insofar as international trade is particularly sensitive to investment (IMF, 2016). Finally, a fifth mechanism is driven by the fall in real commodity prices that tends (as I document next) to accompany a stronger dollar.[11]

Trade fluctuations have disproportionate effects on smaller and more open economies, especially EMDEs. Another channel through which dollar exchange-rate movements affect many of them is the dollar's association with commodity prices. (In 2019, about 20 percent of world trade consisted of primary commodities, but the exports of poorer countries were disproportionately concentrated on commodities.) Figure 10 shows the strong negative correlation between nominal dollar appreciation and changes in dollar commodity prices. The simple correlation coefficient is –0.72 over 1993–2019. Part of the strong negative correlation between the GFCy index and the dollar comes through the dollar's negative association with commodity prices. It may not be immediately obvious that commodity-price declines due to a stronger dollar harm the real incomes of the exporting countries. Let $E_{lc/\$}$ be the local-currency price of the U.S. dollar, let $P_{\$}^{comm}$ be the world dollar price of commodities, and let P_{lc}^{GDP} be the local GDP deflator in terms of domestic currency. Then the price of commodities in terms of exporter GDP equals $E_{lc/\$} P_{\$}^{comm} / P_{lc}^{GDP}$. If a stronger dollar means that all nominal dollar prices fall in proportion—as in the case of a purely monetary shock in a flexible-price world—then $E_{lc/\$}$ rises (local currency depreciates) in the same proportion as $P_{\$}^{comm}$ falls. With the local price level unchanged, the real price of the commodity export in terms of local output would remain unchanged, as would local real incomes.

10. As Bruno and Shin (2021) point out, dollar invoicing of exports likely increases the demand for dollar-denominated trade credits (since the short dollar position is naturally hedged), thus accentuating the impact of dollar appreciation through the previous mechanism.

11. See also Druck and others (2018).

Figure 10. Dollar Commodity Prices Tend to Fall when the U.S. Dollar Appreciates in Nominal Terms



Source: International Monetary Fund, World Economic Outlook database, April 2021, trade volume of goods and services; FRED, dollar exchange-rate series TWEXMANL, trade-weighted based on goods trade with major-currency trading partners (Euro area, Canada, Japan, United Kingdom, Switzerland, Australia, and Sweden). The price index covers fuel and non-fuel commodities.

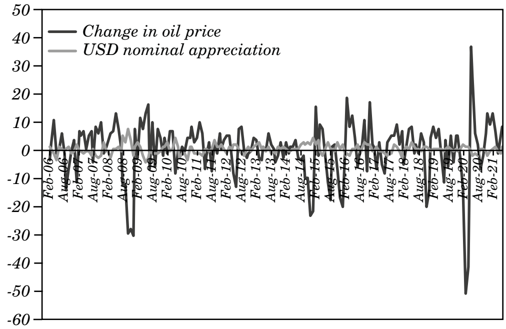
Table 1. Monthly Correlation Between Change in Nominal Dollar Index Against Major Currencies and Change in Real Local Commodity Price, February 2006–June 2021

Brazil	*Chile*	*India*	*Saudi Arabia*	*South Africa*	*Thailand*
−0.20	−0.35	−0.44	−0.58	−0.21	−0.45

Source: U.S. nominal effective exchange rate against advanced country currencies from Federal Reserve Board of Governors. Monthly dollar commodity price index from IMF Primary Commodity Prices website. Monthly local CPI data and country exchange rates against the U.S. dollar from FRED. For Thailand, CPI from national sources via Macrobond.

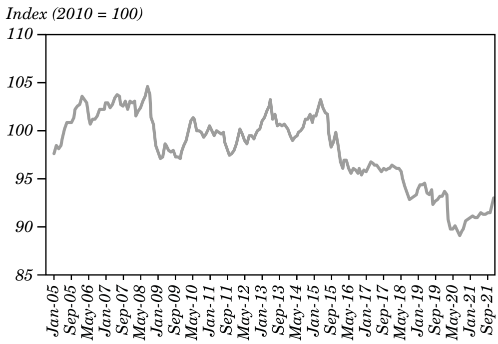
But this is far from what happens in practice when the dollar becomes stronger. When the dollar appreciates by x percent in nominal effective terms against other advanced-country currencies, $E_{lc/\$}$ may well rise by less than x percent: some commodity exporters intervene in foreign exchange to limit exchange-rate movements (“fear of floating”), while others may peg their currencies to the dollar outright. More importantly, $P_{\$}^{comm}$ will tend to fall by more than x percent, as is evident from the much larger scale of the left axis in figure 10. Both factors result in a fall in the relative price $E_{lc/\$}\, P_{\$}^{comm} / P_{lc}^{GDP}$ when the dollar appreciates, and a consequent fall in exporter real income. A stronger dollar, if not accompanied by a rise in global commodity demand, will hammer primary exporters' terms of trade and real incomes. For six emerging markets, table 1 shows the negative correlations between monthly movements in the real local value of the IMF commodity price index (using CPIs to stand in for GDP deflators) and the Federal Reserve effective dollar index against advanced-country currencies.

Figure 11. U.S. Dollar Appreciation and Change in Dollar Oil Price, Monthly Data



Source: IMF dollar oil price index from IMF Primary Commodity Prices website. Exchange rate is effective nominal dollar rate against advanced economies, as reported by the Board of Governors of the Federal Reserve System.

Figure 12. Emerging-Market Nominal Broad Effective Exchange-Rate Index, 2005–2021



Source: Monthly data from J.P. Morgan via Macrobond.

Figure 11 focuses on the case of oil prices, showing their outsized fluctuations compared with those in the dollar nominal effective rate. The correlation coefficient between the price changes for the dollar and oil is −0.39 over the period shown.[12]

Rey's (2013) important paper on the global financial cycle focused attention on the degree to which more flexible exchange rates can help countries, and especially EMDEs, steer an independent policy course amid the monetary and financial shocks arriving through global capital markets. An earlier "fear of floating" literature (Calvo and Reinhart, 2002) pointed out that with faster passthrough of exchange rates to domestic prices and more dollarized domestic debts, EMDEs faced a harsher policy tradeoff between stabilization and inflation in responding to adverse foreign shocks with currency depreciation, and would therefore opt for more limited exchange-rate flexibility.[13] Even earlier, Cooper (1999) argued that exchange-rate movements driven by capital flows could be a source of discomfort for policymakers.

The "trilemma versus dilemma" description of this problem is simplistic. Even among the most ardent proponents of flexible exchange rates, few have contended that they would provide perfect insulation against all shocks. Countries may well face more difficult tradeoffs owing to fluctuations in global financial conditions: this happens when some instruments become less effective at promoting desired macroeconomic responses while simultaneously inflicting more unintended consequences. Yet, exchange-rate flexibility still affords a precious degree of freedom for policy, without which macro outcomes would be worse overall.[14] The need for flexibility may be greatest during crises, when exceptional policies can be brought to bear to mitigate the adverse side effects of large exchange-rate movements, for example, allocating foreign exchange reserves to the economy's systemically important foreign-currency debtors. In both the Global Financial Crisis and the crisis associated with the outbreak of Covid-19, many EMDEs allowed the currencies to depreciate sharply (figure 12).

Recent studies affirm that policy tradeoffs are indeed worse for EMDEs, but that exchange-rate flexibility mitigates the negative impacts of various shocks. Klein and Shambaugh (2015) conclude that

12. Simple ordinary least squares (OLS) regression of the oil-price change on dollar appreciation (both in natural logarithms) yields a coefficient of −2.45 (standard error of 0.42, $R^2 = 0.15$).

13. Gourinchas (2017) presents a notably clear account of this tradeoff.

14. See Obstfeld (2015).

for EMDEs, capital controls afford relatively little policy autonomy unless they are extensive, whereas policy autonomy (in the sense of independence of short-term interest rates) rises with more exchange-rate flexibility. Looking in detail at the case of Chile, Gourinchas (2017) estimates a dynamic model in which a conventionally responsive domestic monetary policy will help mitigate spillovers from foreign shocks, so that "flexible exchange rates remain the primary line of defense against foreign monetary policy and global financial cycles alike."[15] Based on quarterly 1996–2018 data for 55 emerging markets and 14 advanced economies, Kalemli-Özcan (2019) finds that tighter U.S. monetary policy propagates powerfully to EMDEs (though not to advanced economies) through capital flows and increases in interest-rate risk premia. However, she also finds that exchange-rate flexibility can moderate the impact on economic activity. In data for a quarterly panel of 40 emerging market economies over 1973–2016, Ben Zeev (2019) finds that countries with pegs fare significantly worse (in terms of output, exports, asset prices, and other key variables) in the face of contractionary Gilchrist-Zakrajsek credit shocks than countries with more flexible regimes. Using a large global set of monthly data spanning 30 advanced and emerging economies over 1990–2018, Degasperi and others (2021), reaffirm the Kalemli-Özcan result that U.S. monetary policy affects emerging markets through higher term premia regardless of exchange-rate regime, but conclude (pp. 3–4) that "both real and nominal spillover effects are larger in countries with more rigid exchange-rate regimes." This relatively short list of studies is selective rather than complete, but it stands in for a much larger body of evidence pointing in the same direction.

The Global Financial Cycle impacts all countries in some way, whether advanced, emerging, developing, or a high-income emerging market like Korea that is nonetheless subject to volatile capital flows. Higher-income economies seem to absorb the resulting shocks more easily, due to deeper and more fluid financial markets, their wealth, their productive diversity in many cases, the generally greater credibility of their policy frameworks, and elements of the global financial safety net from which they benefit disproportionately. Nonetheless, the initial phase of Covid-19 indicated that emerging market economies too had policy space to address the crisis—in part by exploiting exchange-rate flexibility, and with an assist from macroeconomic support policies in advanced economies.

15. Gourinchas (2017, p. 282).

3. Emerging Market Policy Responses to the Initial Covid-19 Shock

The appearance of the global pandemic inflicted massive external real and financial shocks on EMDEs. Global trade collapsed in the first quarter of 2020, to a degree comparable with 2008's trade collapse. Korea of course did not escape this shock but suffered to a degree less than the global average. The financial shock manifested in a sharp reversal of capital inflows in March 2020. Figure 13 shows the pattern of portfolio capital inflows for a group of 26 mostly middle-income countries, including Korea. Figure 14 shows the Korean data, which suggests a March-2020 capital-flow reversal comparable with that around the Lehman event.

Korea is a high-income economy with a very flexible exchange rate, credible policies, and an evolved macroprudential framework including measures targeting foreign-currency liabilities (IMF, 2017a; Lee, 2017). Its monetary and financial policy reactions to the Covid-19 crisis parallel those successfully used elsewhere in many economies and notably in EM economies.

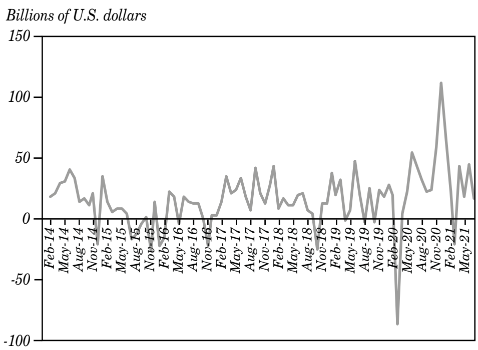
English and others (2021) offer an excellent compendium on central banks' responses to the initial phase of the Covid-19 crisis, with the chapter by Céspedes and De Gregorio (2021) focusing on emerging economies. While the details differ among EMs—indeed, Indonesia went so far as to allow temporarily direct financing of the fiscal deficit by Bank Indonesia—a partial list of measures undertaken by EM central banks often included the following:

- Interest rate cuts.
- Large-scale central-bank purchases of domestic assets, mostly sovereign debt.
- Foreign exchange intervention.
- Looser reserve requirements (including loosening those discouraging capital inflows).
- Liquidity enhancing operations.
- Measures to promote bank loans to businesses.
- Macroprudential easing (e.g., relaxed capital requirements).
- Market functioning enhancements.

EMDEs benefited, however, from the massive monetary and financial stimulus provided by advanced economies early in the crisis and especially from the easing actions of the U.S. Federal Reserve. These actions underpinned the sudden reversal of negative capital inflows after March 2020, evident in figures 13 and 14. Although capital flows have continued to be volatile since then, even turning negative again in a few months, the financial environment has generally remained benign for EMDEs so far, as a new expansive phase of the global financial cycle has set in. In particular, the generalized wave of EMDE sovereign defaults that some predicted at the outbreak of the crisis did not materialize in 2020–2021, despite those countries' aggressive use of their monetary and fiscal policy space.

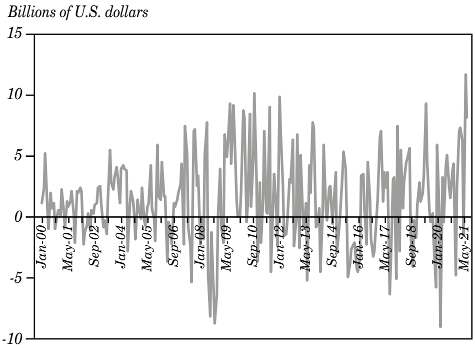
Providing important support to the global economy, the Fed extended dollar swap lines to 14 central banks, reducing the cost and lengthening the tenor of its offerings. Although only two emerging economies—Brazil and Mexico—were offered swap lines, as they were in 2008, the facilities offered to advanced economy authorities can help stabilize conditions in a broader region that includes emerging markets (for example, the impact on emerging Europe of swap lines to Nordic central banks). In the current crisis, the locus of swap line usage shifted geographically compared with the Global Financial Crisis, from Europe to Asia. This time, drawings by the Bank of Japan exceeded those by the ECB, and the Bank of Korea and the Monetary Authority of Singapore also participated (Gislén and others, 2021).

Figure 13. Capital Inflows to 26 Emerging Market Economies, 2014–2021



Source: Data for a group of mostly middle-income countries assembled by Koepke and Paetzold (2020).

Figure 14. Capital Inflows to the Republic of Korea, 2000–2021



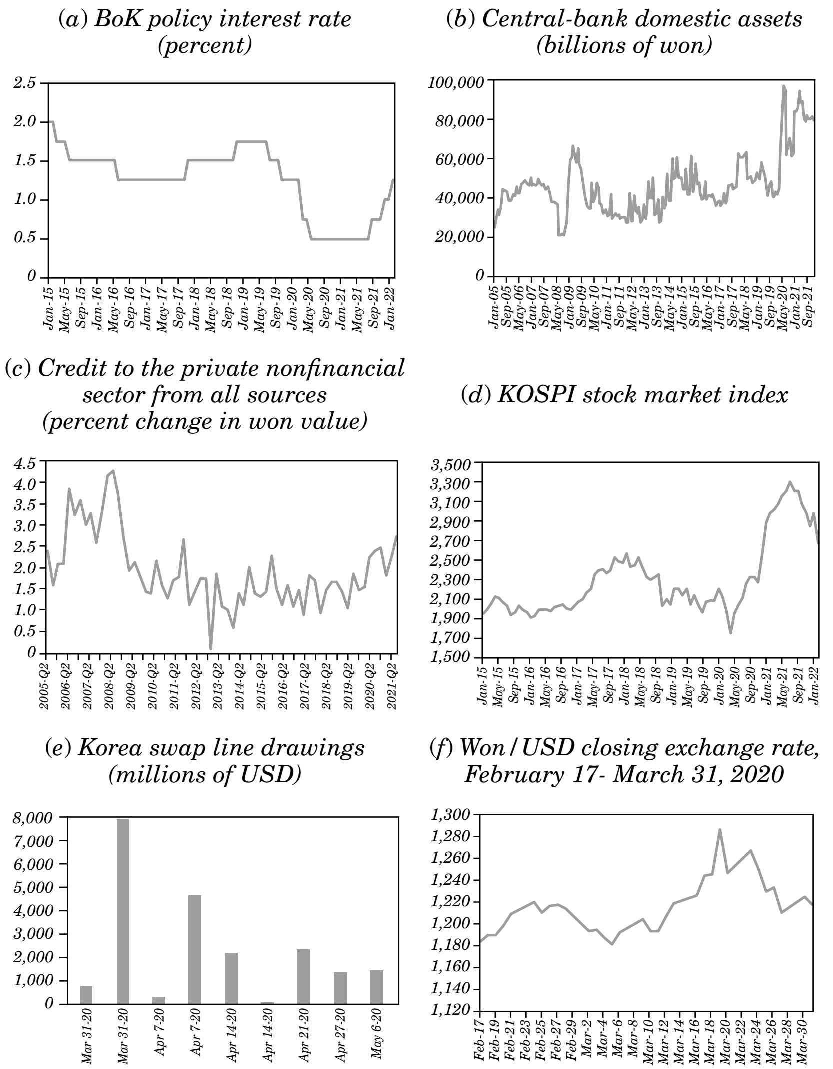
Source: Data for a group of mostly middle-income countries assembled by Koepke and Paetzold (2020).

Figure 15 summarizes aspects of Korea's response. The BoK promptly cut its policy interest rate, though not all the way to zero [panel (a)]. It also expanded its balance sheet [panel (b)]. (The BoK has already raised the rate three times more recently in the face of inflation and financial stability concerns.)

Céspedes and De Gregorio (2021) emphasize how emerging-market central banks were able to maintain domestic credit growth in 2020, unlike the experience in the Global Financial Crisis. For Korea, this pattern is evident in panel (c). Credit growth rose once the Covid-19 crisis hit, unlike its decline in 2008–2009 (albeit then, from very high levels that were symptomatic of the forces generating the previous crisis). In line with lower interest rates and the growth in domestic credit, panel (d) indicates that Korea has participated in the current expansive phase of the global financial cycle, with a sharp increase in its equity prices, as in the United States and other countries following the initial crash in March 2020.

Korea drew several times on its \$60 billion swap line with the Fed [panel (e)], auctioning these dollars to domestic banks with dollar funding needs. Even the announcement of the swap agreement had a dramatic impact on the foreign exchange market. Korean authorities allowed the won to depreciate sharply during the generalized panic after the WHO's March 11, 2020 declaration of a global pandemic [panel (f)]. The won/dollar exchange rate reached a high point on March 19; later that day the Fed announced the Korean dollar swap line, prompting an immediate reversal in the won's depreciation.

Figure 15. Korea Responses to the Covid-19 Crisis



Sources: Federal Reserve Bank of New York; accessed via Macrobond: Bank for International Settlements, Bank of Korea, and Korea Stock Exchange.

4. Continuing Vulnerabilities for EMDEs

EMDEs' ability to use monetary (as well as fiscal) policies to mount strong counter-cyclical responses was a positive surprise at the start of the Covid-19 crisis. In general they built on the accumulated capital of monetary-policy credibility (which had reduced EMDE inflation rates to low levels compared with past decades), on the increasing intellectual sophistication and operational expertise of their policymakers, on a comparatively strong cyclical position at the start of 2020, and on a strong lift from expansionary policies in advanced economies in the face of a shock with initially deflationary consequences. They departed from past practice also in more fully exploiting exchange-rate flexibility, cutting interest rates even as their currencies depreciated in the face of a capital-flow sudden stop.[16] This response suggests that the trilemma has not collapsed to a simple dilemma: open capital account without monetary autonomy, or closed capital account with monetary autonomy—regardless of the exchange-rate regime.

Nonetheless, EMDEs could be vulnerable to sudden stops in the near-term future as the next contractionary phase of the global financial cycle is getting underway.[17] Two current factors make this more likely.

First, the rollout of vaccines has been slower in most EMDEs than in advanced economies, and in many cases much slower. Moreover, some EMDEs are using less effective vaccines (notably less effective against the Delta variant of SARS-CoV-2), while often even vaccines that are available can go to waste due to underdeveloped infrastructures for getting shots into arms. The Covax mechanism has failed to meet even its modest targets as rich countries have effectively hoarded vaccine doses. In the longer run, this imbalance will threaten even highly vaccinated countries because unvaccinated regions will remain breeding grounds for new resistant variants; but in the near term, it implies a more rapid recovery in the advanced world than in EMDEs, with a consequent rise in global interest rates while EMDEs are still struggling.

Second, EMDE fiscal responses to the crisis have made them more vulnerable to hikes in advanced-economy interest rates—which could set off a contractionary phase of the global financial cycle. In

16. See also Aguilar and Cantú (2020).
17. Kalemli-Özcan (2021), IMF (2021), and Obstfeld (2021b) voice similar concerns.

advanced and less prosperous countries alike, fiscal deficits grew in 2020 as governments intervened to support firms and households during lockdowns, raised public-health spending, and lost revenues due to compressed economic activity levels. In many EMDEs, public revenue fell even as a percent of their lower levels of GDP. While fiscal responses in EMDEs were not as extensive as those of advanced economies, the EMDEs have historically been constrained to lower debt levels due to their less-developed revenue capacities and capital markets. Being able to fund sovereign debt in domestic currency is no panacea because higher debt levels undermine inflation credibility more quickly for EMDEs and raise their vulnerability to capital-flow reversals (Carstens and Shin, 2019).

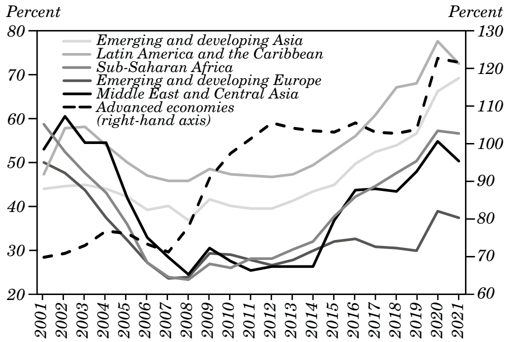
Figure 16 shows the development of general public debt-GDP ratios in advanced economies and in emerging and developing regions. (Figures for 2021 are IMF projections as of October 2021.) While the 2020 runup in advanced economies (tracked on the right-hand axis) is bigger in absolute terms, all EMDE regions also show significant jumps for that year. Moreover, in all regions, debt-GDP ratios had already been rising since the early 2010s. Figure 17 offers a more relevant comparison of the percent increases in debt-GDP ratios in the country groupings. Here, advanced economies are in the middle of the pack for 2020. Broadly speaking, EMDEs' changes in debt-GDP ratios were comparable to those of advanced economies, conditional on the lower debt capacity of the former group. The improvement in EMDE debt ratios the IMF assumes for 2021 relies on relatively optimistic growth forecasts and also reflects less ambition in fiscal support policies—although greater fiscal support might be needed to generate the assumed growth.[18]

In short, higher interest rates in advanced economies will put greater stress on public finances in EMDEs. They will also harm the fortunes of EMDE corporates that borrowed more since the crisis began, a downside legacy of the continuing domestic credit growth that supported EMDE economies in 2020. The same observations apply to the macroprudential easing policies that were positive for growth in 2020.[19]

18. The sharp 2021 reduction in debt ratios for the Middle East and Central Asia is the result of elevated energy prices in that year, driven by global recovery and a fairly restrictive policy by OPEC+.

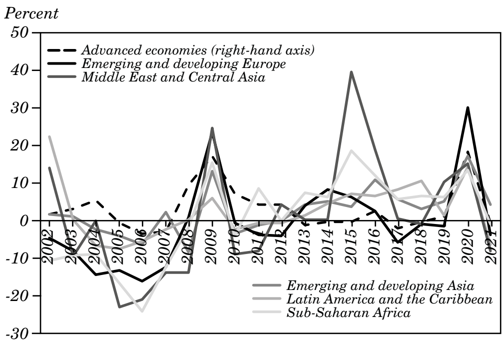
19. See Bergant and Forbes (2021).

Figure 16. General Government Debt-GDP Ratios, Advanced and EMDE Economies



Source: IMF, *World Economic Outlook* database, October 2021.

Figure 17. Percent Changes in General Public Debt-GDP Ratios, Advanced and EMDE Economies



Source: IMF, *World Economic Outlook* database, October 2021.

Figure 18 focuses on one particular source of potential fragility, the concentration of new sovereign-debt issuance on domestic bank balance sheets in a number of EMDEs.[20] This pattern sets up the possibility of a sovereign-bank doom loop. As Kalemli-Özcan (2019) shows, U.S. monetary tightening transmits to EMDEs via a rise in longer-term bond premia, and therefore a fall in bond prices. By weakening EMDE bank balance sheets, that development could set up

20. See Sachdeva and Harvey (2020), and IMF (2021).

destabilizing expectations of government fiscal intervention to support the banking sector, higher deficits, more accommodative monetary policy, and yet lower bond prices. Figure 18 also indicates that in the first year of the Covid-19 crisis, foreign investors on the whole reduced their sovereign exposures. Higher domestic saving due to the lockdowns facilitates the domestic placement of sovereign debt, but with recovery, higher saving rates will not persist. A further challenge, facing advanced and less prosperous economies alike, comes from the inflationary pressures that supply-chain disruptions are exacerbating.

We should therefore expect heightened financial fragility as an uneven rebound unfolds in the world economy. Apart from the homegrown problems that advanced economies may face emerging from a period of prolonged policy accommodation, they could face significant spillovers from EMDE woes. How resilient will global financial markets prove in the face of these pressures?

5. Enhancing the Resilience of Global Financial Markets

Reforms in several directions could strengthen the global financial system to face the turbulence that may lie ahead. Most of these proposals reflect long-standing needs, although the experience in the recent Covid-19 crisis underscores the urgency of action.[21]

In the spring of 2020, banks avoided the widespread distress of the Global Financial Crisis. In large part this success owed to the origin of the Covid-19 shock being outside of the banking sector. But some credit is also due to the national and international banking sector reforms that followed the 2008–2009 crisis and the euro-area crisis, which augmented bank capital, enhanced the liquidity of balance sheets, and upgraded prudential regulatory frameworks in many countries.

A predictable side effect, however, has been the migration of financial activity from the more constrained banking sector to unregulated or loosely regulated nonbank financial actors. In its recent report, the Committee on the Global Financial System (CGFS) of the BIS notes several changes in the structure of international capital flows, but first among them is the growing share of market-based capital flows (CGFS, 2021).[22] Since 2007, the share of bank loans in

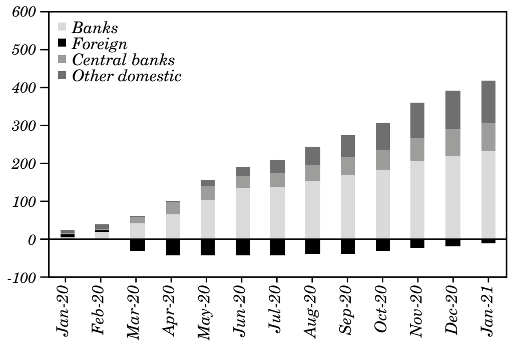
21. See also Eguren-Martin and others (2020).
22. See also Lane and Milesi-Ferretti (2018).

the external debt of advanced economies has shrunk from about 35 percent to about 22 percent, whereas the share of portfolio debt has risen from about 43 percent to 50 percent. At the same time, the share of bank loans in the external debt of emerging-market borrowers has fallen from around 52 percent to 45 percent, and the share of portfolio debt has risen from around 24 percent to nearly 40 percent. Advanced economy cross-border bank claims (which include debt securities, not just loans) declined from about 70 percent of home-country GDP at the time of the GFC to around 50 percent in 2019 (CGFS 2021, graph 1.2). Eguren-Martin and others (2020) document the dominant role of nonbank actors in the reversal of EMDE capital flows in March 2020.

Returning to the TIC data, figure 19 shows how the foreign position of U.S. banks and other financial institutions has essentially been stagnant in nominal terms since just before the GFC.

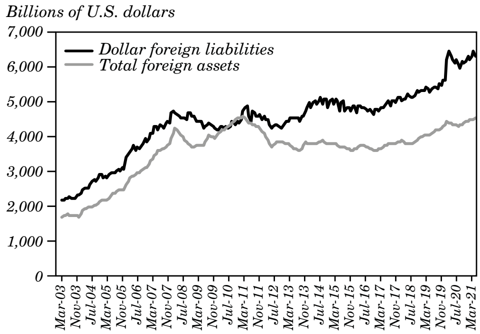
At the same time, and as noted earlier, the cross-border activity of emerging-market banks has risen—according to CGFS (2021)—from about seven to nine percent of home GDP between 2008 and 2019. However, it remains small in scale compared with advanced economies' international bank activity.

Figure 18. Domestic Sovereign-Bond Holdings in 12 Emerging Market Economies
(cumulative change, billions of U.S. dollars)



Source: Updated data for figure 1.6.6 in IMF (2021), provided by IMF Money and Capital Markets Department.
Note: The emerging market economies include Brazil, Colombia, Mexico, Indonesia, Malaysia, Ukraine, Turkey, Thailand, South Africa, Poland, Hungary, and Romania.

Figure 19. U.S. Banks' and Other Financial Institutions' Foreign Assets and Liabilities, 2003–2020



Source: U.S. Treasury, Treasury International Capital System, U.S. Banking Data. Monthly asset data are interpolated quarterly data. The dollar liability data cover about 95 percent of total liabilities (that is, liabilities in all currencies). The liability series also encompasses all non–U.S. holdings of short-term Treasury securities.

From a policy perspective, this evolution points to the need for more thinking about financial stability risks coming from the nonbank sector, for example, through increasingly complex intermediation chains that may ultimately also impinge on the banks. The spread of innovative fintech platforms only increases the risks, including from cybersecurity breaches, and may render prudential oversight more difficult. All along, climate-related risks are only rising. The challenges that the international dimension raises are particularly big, owing to the seams between national regulatory systems. The Financial Stability Board (FSB) has outlined an extensive program to assess the risks from nonbank financial institutions in light of the Covid-19 market turmoil of spring 2020 (FSB, 2020). However, it seems fair to say that even bank regulation now needs to encompass an even broader set of potential systemic risks than were envisioned in the immediate post-GFC reforms. The trend of emerging-market banks increasingly venturing abroad into other emerging markets only raises the stakes for those countries.

Another part of the financial market infrastructure in need of strengthening is the global financial safety net (GFSN). Bilateral swap lines have become increasingly important in the GFSN (Perks and others, 2021). Federal Reserve swap lines were essential in stabilizing global markets in the spring of 2020 in light of the dollar's continuing dominance as a funding and investment currency. But the geographic coverage and market reach of those swap lines were limited, especially because dollar funding activity has tended to migrate from the European theater that was dominant in the GFC to Asia and emerging markets (CGFS, 2020).

The need to extend central-bank swap lines multilaterally, especially the Fed's, has long been apparent,[23] though it remains unclear what institutional structure would be most politically acceptable to the issuers of funding currencies, and what lending safeguards would be necessary. At the least, building trust would demand a higher degree of coordination in financial regulatory policies than now exists. In 2017, IMF staff developed a proposal for a Short-term Liquidity Swap facility to "provide liquidity support for potential balance-of-payments needs of a short-term, frequent, and moderate nature, resulting from volatility in international capital markets" (IMF, 2017b). The facility was meant to be available to countries with "strong fundamentals," and without ex-post conditionality. The IMF Executive Board divided on the proposal, which some major shareholders opposed, and turned it down. Amid the market disruption in April 2020, however, the Fund Board approved a similar Short-term Liquidity Line (SLL) facility intended to address some of the gaps in the network of bilateral swaps. Unfortunately, potential beneficiaries seem not to view the SLL (or the Fund's two other precautionary credit lines originating in the GFC period) as equivalent to central-bank swaps, and indeed, not a single country has drawn on the SLL so far. Plant and Rojas-Suárez 2021 provide an excellent discussion of the likely reasons, as well as of ways the IMF could encourage take-up of the facility. The IMF declined to adopt the pandemic support facility that Fisher and Mazarei (2020) proposed, but such a policy instrument would also strengthen the GFSN during the current pandemic and could be mobilized in future contagious outbreaks. Also relevant is the proposed Resilience and Sustainability Trust, which would provide an IMF umbrella for richer countries to lend SDRs for investments in climate adaptation, health, and other areas of vulnerability.[24] The upcoming Sixteenth General Review of IMF quotas will provide another opportunity to strengthen the GFSN through enhanced non-borrowed lending resources.

The U.S. market for Treasury securities showed unexpected dysfunctionality in March 2020, notably during a "dash for cash" later in the month when Treasurys became temporarily illiquid as

23. For example, see Obstfeld (2009).
24. See also G30 Working Group on Sovereign Debt and Covid-19 (2021).

domestic and foreign holders rushed to sell them for money (Duffie, 2020; FSB, 2020). The dollar remains by far the central currency in the international financial system (CGFS, 2020) and, for better or worse, no serious competitor is yet in view. At the same time, central-bank dollar reserves play a key role in the overall resilience of the GFSN. If central banks or sovereign wealth funds cannot rely on converting their Treasury holdings at par, those reserves become less effective in providing insurance to their holders. Thus, the health of the Treasury market is vital to that of the GFSN, and measures that strengthen its functioning also strengthen the GFSN.[25]

To enhance the liquidity of Treasurys amid the turmoil, on March 31, 2020, the Fed established the Foreign and International Monetary Authorities (FIMA) repo facility for converting official foreign Treasury into cash. It became a standing facility on July 28, 2021. (Reflecting ongoing tensions in domestic markets, in June 2020, the BoK floated an analogous facility to allow domestic banks, insurance companies, and brokerages to swap U.S. Treasurys into dollar cash.)[26] Several changes would enhance the plumbing of the U.S. Treasury market, the most far-reaching of which would be central clearing of transactions in the market, including repo.[27]

For EMDEs, improved defensive policies can bolster resilience—and thereby global resilience. Their vulnerability to the global financial cycle makes it understandable why so many less affluent economies, even emerging market economies, have stopped short of full financial opening (recall figure 2). In 2012 the IMF officially recognized this reality by developing an "institutional view" (IV) on capital controls that allows for their use in some circumstances, notably when financial flows threaten economic or financial stability and the capital-flow measures (CFMs) do not substitute for necessary adjustments in macroprudential, monetary, or fiscal policies (IMF, 2012).[28] The Fund's acceptance of CFMs as a legitimate policy tool was a huge shift in approach: an aversion to exchange control resides deep within the institution's DNA, and even an attempt to focus surgically on cross-border financial transactions could spill over to the current account.

25. Euro reserves are also an important component of global international reserves and, in the spring of 2020, euro bond markets also experienced liquidity problems.

26. See Roh and Park (2020).

27. For reform proposals, see Duffie (2020), G30 Working Group on Treasury Market Liquidity (2021), and Hubbard and others (2021).

28. Even before the IV, however, IMF staff accepted and even recommended capital controls in some individual country cases. For the case of Iceland in 2008, see Honohan (2020).

Nonetheless, the IV is in several ways too restrictive. Research shows that CFMs are rarely imposed in the temporary manner the IV envisions, in response to cyclical tides in the global capital market. Instead, they are generally structural and thus long-lived in nature. Notwithstanding the IV, many Fund members feel that global markets might stigmatize them if they vary CFMs reactively. Thus, the Article IV surveillance process has regularly featured disagreements between Fund staff and country authorities as to whether particular policy measures should be labeled as CFMs or MPMs (macroprudential measures), with the authorities often advocating for the latter designation (Everaert and Genberg, 2020).[29] A particular cause of disagreement has been policy in some countries (including some richer countries such as Canada) to limit foreign speculative purchases of property in soaring real-estate markets. Finally, the IV is asymmetric with respect to inflow and outflow controls, restricting the use of the latter to situations of imminent or ongoing crisis. The Fund's internal Independent Evaluation Office (2020) recognized these criticisms in a comprehensive review and recommended rethinking the IV.

Recently the Fund has proposed an Integrated Policy Framework that conceptualizes the use of CFMs, foreign exchange intervention, monetary policy, fiscal policy, and macroprudential policy as distinct instruments that may all be needed to reach multiple policy goals in a small open economy (IMF, 2020).[30] Importantly, the approach has the potential to place capital control and foreign exchange intervention policies on an equivalent plane with monetary, fiscal, and macroprudential policies, and thereby remove some of the stigma that currently adheres to CFMs. In light of this work and the limitations of the IV, the Fund is currently reconsidering its advice on CFMs and could go further in the direction of regularizing their use in a wider set of circumstances.[31] This approach would also be in line with the

29. CFMs can play a macroprudential role—for example, when they limit foreign funding of imprudent domestic investments—but they can also play other policy roles that IMF rules proscribe—for example, preventing adjustment of an undervalued exchange rate. In contrast, a hypothetical 'pure' MPM would not discriminate in its implementation between domestic and foreign residents. The overlap in the roles of MPMs and CFMs has sometimes blurred the distinction between them, as has the difficulty smaller countries face in counteracting the global financial cycle through MPMs without the support of measures that could be characterized (at least partially) as CFMs.

30. See Jeanne (2021) for a related framework.

31. As Honohan (2020, p. 25) aptly puts it, the 2012 IV approach "is quite different from seeing [capital-flow] measures as a tool to be actively integrated with monetary, exchange-rate, and macroprudential measures."

recent recommendations of a group of ASEAN central banks (ASEAN WC-CAL, 2019). Following a 2016–2019 review, the revised OECD Code of Liberalisation of Capital Movements addresses some of the same criticisms IMF member countries have raised concerning the IV (OECD, 2020).

If a future sudden stop in capital flows to EMDEs is protracted, and especially if the pandemic lingers on, liquidity support may not be enough to stave off solvency problems. Despite some recent improvements, however, the current international architecture for external debt restructuring is inadequate to handle a rash of sovereign defaults, some potentially affecting systemic countries.[32] Earlier hints by the Group of Twenty pointing toward mandatory private-sector participation in debt restructurings have fallen by the wayside as global financial conditions have remained easy. It should not take a renewed financial crisis to revive those ideas.

32. See G30 Working Group on Sovereign Debt and Covid-19 (2021).

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Exchange Rate Puzzles and Policies

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What is the optimal exchange rate policy? Should exchange rates be optimally pegged, managed, or allowed to freely float? What defines a freely floating exchange rate? Do open economies face a trilemma constraint in choosing between inflation and exchange rate stabilization, unlike divine coincidence in a closed economy? These are generally difficult questions, as the exchange rate is neither a policy instrument, nor a direct objective of the policy, but rather an endogenous general-equilibrium variable tied by equilibrium relationships in both goods and financial markets. At the same time, equilibrium exchange rate behavior features a variety of puzzles from the point of view of conventional business-cycle models typically used for policy analysis in open economy.

We address these questions by developing a general policy analysis framework with nominal rigidities and financial frictions that are both central for equilibrium exchange rate determination and result in an empirically realistic model of exchange rates. The model builds on Itskhoki and Mukhin (2021a,b) and is consistent with the exchange rate disconnect properties across floating and fixed regimes allowing for explicit policy analysis using both monetary policy and foreign exchange (FX) interventions in the financial market. The model features Balassa-Samuelson mechanism determining the value of the frictionless real exchange rate (departures from purchasing power

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parity, PPP) and segmented financial markets resulting in endogenous equilibrium Uncovered Interest Rate Parity (UIP) deviations. The presence of both endogenous PPP and UIP deviations is essential for the optimal exchange rate policy analysis, as exchange rate variation is at the core of both deviations. We show that this framework is easily amenable to normative analysis and characterize the optimal exchange rate policies following Itskhoki and Mukhin (2022).

In section 1, we setup a simple small open economy model with a tradable and a nontradable sector. While highly stylized, this model allows us to illustrate the key mechanisms and derive the main policy insights that generalize in richer quantitative frameworks. In particular, in section 2, we show how this simple model captures the essential empirical properties of exchange rates, including the Meese-Rogoff disconnect and the Backus-Smith puzzles, in addition to PPP and UIP puzzles mentioned above. While macroeconomic aggregates are driven primarily by fundamental macroeconomic shocks such as productivity and monetary shocks, exchange rates are primarily driven by shocks emerging in international financial markets, for example, shifts in demand for different currencies that have little direct macroeconomic impact. This explains both vastly larger volatility of exchange rates relative to other macro variables—both nominal like inflation and real like consumption and GDP growth—and weak patterns of correlation between these variables and exchange rates.

More importantly, our simple model also reproduces Mussa facts on macroeconomic comovement with exchange rates associated with a switch between floating and fixed exchange rate regimes. As Mussa (1986) famously observed, the real exchange rate has changed dramatically its equilibrium behavior, along with the nominal exchange rate, immediately after the end of the Bretton Woods system of fixed exchange rates. This constitutes prime evidence in favor of non-neutrality of monetary policy regimes. At the same time, as first emphasized by Baxter and Stockman (1989), other macroeconomic aggregates, whether nominal or real, did not exhibit any comparable change in their statistical properties after the end of Bretton Woods. We argue that this set of Mussa facts requires that monetary non-neutrality emerges from the financial market, where international risk-sharing wedges endogenously respond to equilibrium exchange rate volatility. Indeed, a credible nominal exchange rate peg eliminates one of the main sources of risk in international financial transactions. As a result, financial arbitrageurs become more willing to intermediate international capital flows, resulting in smaller equilibrium UIP deviations. This, in turn, eliminates the primary source of exchange rate volatility under the float, allowing the government to achieve a credible peg without a major shift in equilibrium monetary policy. This explains why macroeconomic aggregates do not exhibit a dramatic change in their equilibrium behavior.

We describe the model of a segmented financial market with limits to arbitrage that is consistent with this Mussa mechanism. Endogeneity of international risk-sharing wedges and UIP deviations to the exchange rate regime is the key feature of the model to both explain the Mussa evidence and to provide new insights into the optimal exchange rate policy using a mix of monetary tools and FX interventions, which is the focus of section 3.

At the core of our analysis is the dual role played by the nominal exchange rate. First, it allows for adjustment of the real exchange rate when prices (or wages) are sticky. In the absence of such nominal exchange rate movements, the economy features an output gap resulting in welfare losses. Monetary policy can eliminate the output gap, but this generally requires a volatile nominal exchange rate. Second, the volatility of the nominal exchange rate limits the extent of international risk sharing in the financial market, as international financial transactions are intermediated by risk-averse market makers who need to hold the nominal exchange rate risk. This also leads to welfare losses. Financial-market interventions can redistribute the risk away from arbitrageurs, stabilizing resulting equilibrium UIP deviations and improving the extent of international risk sharing.

First, we prove a divine coincidence result in an open economy: if the frictionless real exchange rate is stable, then a fixed nominal exchange rate achieves both goals of output-gap and UIP stabilization, and thus is the optimal policy choice. Furthermore, direct nominal exchange rate targeting is favored over inflation stabilization, even though both policies have consistent goals. While the former policy ensures stable inflation as a result of exchange rate targeting, the latter policy may result in multiple equilibria in the international financial market, with and without nominal exchange rate volatility.

Second, we show that access to unconstrained monetary policy and FX interventions generally allows to implement the optimal allocation, independently of whether the frictionless real exchange rate is stable or not. The resulting equilibrium generally features volatile nominal exchange rate and inflation targeting, with financial interventions eliminating the intermediation friction and stabilizing UIP deviations. We also show that economies with segmented financial markets do not feature a conventional trilemma constraint, as market segmentation offers financial regulators an additional tool to stabilize the international financial market, even when monetary policy has an exclusive inward focus on domestic inflation and output-gap stabilization.[1]

Third, we explore various circumstances where either monetary policy is constrained (e.g., due to the zero lower bound) or financial interventions are constrained (e.g., due to non-negativity requirement on central-bank foreign reserves or value-at-risk constraints on the central bank's balance sheet). In this case, there are two independent policy goals—the output gap and the risk-sharing wedge—and only one unconstrained policy tool, thus making it generally impossible to replicate the optimal allocation. Fixing the exchange rate using monetary policy is generally feasible but is also generally suboptimal. Similarly, targeting the output gap alone is also suboptimal, and monetary policy trades off output-gap and exchange rate stabilization (a partial peg) in the absence of FX interventions. Using financial interventions to stabilize output gap is generally infeasible.

Lastly, we explore the ability of the government to extract rents in the international financial market by means of FX interventions. The government can generate expected rents for the country only in the presence of foreign noise traders by leaning against the wind of their liquidity currency demand. Arbitrageurs compete with the government for these rents, and greater equilibrium exchange rate volatility allows the government to capture a greater share of these rents by discouraging arbitrageurs from active intermediation. In general, the policymaker favors small departures from frictionless risk sharing and expected UIP deviations which result in expected incomes of the central bank against the losses of foreign noise traders. Capital controls are generally an imperfect substitute for FX interventions but could be used in combination to increase international rents of the country.

Related literature. We build on a vast literature studying the role of exchange rates in both goods and financial markets, as well as the optimal macroeconomic and financial policies in an open economy. Meese and Rogoff (1983), Mussa (1986), Backus and Smith (1993), Obstfeld and Rogoff (2001), Chari and others (2002), Engel and West

1. In other words, open market operations and sterilized interventions have a bite under financial market segmentation which is a source of departure from Wallace (1981)'s Modigliani-Miller (Ricardian) equivalence in an open economy.

(2005) are some of the most prominent papers studying exchange rate puzzles. The list of exchange rate models with frictional financial intermediation includes Kouri (1983), Jeanne and Rose (2002), Alvarez and others (2009), Gabaix and Maggiori (2015), Gourinchas and others (2019), Greenwood and others (2020), Jiang and others (2021), Bianchi and others (2021).

The normative implications of the expenditure switching channel of monetary policy is the focus of Friedman (1953), Clarida and others (2000), Corsetti and Pesenti (2001), Devereux and Engel (2003), Benigno and Benigno (2003), Gali and Monacelli (2005), Goldberg and Tille (2009), Corsetti and others (2010), Engel (2011), Farhi and others (2014), Egorov and Mukhin (2023), while the financial channel of monetary policy is studied in Farhi and Werning (2012), Rey (2013), Fanelli (2017), Basu and others (2020), Kekre and Lenel (2021), Fornaro (2021). Our analysis is also related to the recent studies of the costs and benefits of exchange rate interventions by Jeanne (2012), Amador and others (2019), Cavallino (2019), Fanelli and Straub (2021) and the optimal capital controls by Jeanne and Korinek (2010), Bianchi (2011), Costinot and others (2014), Farhi and Werning (2016, 2017), Schmitt-Grohé and Uribe (2016).

1. A Simple Model of Equilibrium Exchange Rates

We consider a simple small open economy model with a tradable and a nontradable sector. This stylized model allows us to illustrate the key mechanisms and derive the main policy insights that generalize in richer and more realistic frameworks analyzed in Itskhoki and Mukhin (2021a,b; 2022).

Households. We assume a separable log-linear utility of the households, which allows for a sharp analytical characterization of equilibrium exchange rates and optimal policies with stark policy motives:[2]

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left[\log C_t - (1-\gamma) L_t \right] \text{ with } C_t = \left(\frac{C_{Nt}}{1-\gamma} \right)^{1-\gamma} \left(\frac{C_{Tt}}{\gamma} \right)^{\gamma} \tag{1}$$

2. This assumption combined with homogenous tradables in a small open economy eliminates all markup and terms of trade motives that typically complicate the optimal policy analysis. See Corsetti and Pesenti (2001), Benigno and Benigno (2003), Egorov and Mukhin (2023).

where C_t is the final consumption good, which has a $1 - \gamma$ cost share of nontradable inputs and a γ share of tradable inputs. Without loss of generality, we assume that the household sector assembles the final good from the two inputs minimizing expenditure $P_tC_t = P_{Nt}C_{Nt} + P_{Tt}C_{Tt}$, where P_{Nt} and P_{Tt} are the respective prices. This results in optimal demand $P_{Nt}C_{Nt} = (1 - \gamma)P_tC_t$ and $P_{Tt}C_{Tt} = \gamma P_tC_t$, where the price level $P_t = P_{Nt}^{1-\gamma}P_{Tt}^{\gamma}$.

The households can borrow or lend using one-period risk-free home-currency and foreign-currency bonds (paying out one unit of respective currency next period):

$$P_{Nt}C_{Nt} + P_{Tt}C_{Tt} + \frac{B_t}{R_t} + e^{-\hat{\psi}_t}\frac{\mathcal{E}_tB_t^*}{R_t^*} = B_{t-1} + \mathcal{E}_tB_{t-1}^* + W_tL_t + \Pi_t + T_t, \tag{2}$$

where R_t and R_t^* are the gross nominal interest rates in the two currencies respectively, and $\hat{\psi}_t$ is the friction associated with holding foreign-currency bonds, which we microfound in section 3. The optimal bond holdings satisfy the Euler equations, which we write in the following way:

$$\beta R_t\mathbb{E}_t\left\{\frac{C_{Nt}}{C_{Nt+1}}\frac{P_{Nt}}{P_{Nt+1}}\right\} = 1, \tag{3}$$

$$e^{-\hat{\psi}_t}\beta R_t^*\mathbb{E}_t\left\{\frac{C_{Tt}}{C_{Tt+1}}\frac{P_{Tt}}{P_{Tt+1}}\frac{\mathcal{E}_{t+1}}{\mathcal{E}_t}\right\} = 1, \tag{4}$$

where the nominal exchange rate $\mathcal{E}_t$ is the price of foreign currency in units of home currency (an increase in $\mathcal{E}_t$ is a home-currency depreciation). The household earns labor income W_tL_t, receives profits from home firms Π_t and transfers from the government T_t. Given the log-linear utility, we write the optimal labor supply condition as:

$$P_{Nt}C_{Nt} = W_t. \tag{5}$$

Firms and production. Competitive firms produce the nontradable good using labor, $Y_{Nt} = A_{Nt}L_t$, and are endowed with homogenous nontradable output $Y_{Tt} = A_{Tt}$, where productivity (A_{Nt}, A_{Tt}) follow exogenous and possibly correlated geometric random walk processes. Combined profits of all firms are given by $\Pi_t = P_{Tt}Y_{Tt} + P_{Nt}Y_{Nt} - W_tL_t$.

The law of one price holds for tradables, $P_{Tt} = \mathcal{E}_tR_{Tt}^*$, where P_{Tt}^* is the exogenous foreign-currency world price of tradables. Finally, the prices of nontradables are fully sticky in home currency, $P_{Nt} \equiv 1$. The firms hire necessary amount of labor L_t at flexible wage rate W_t to accommodate nontradable demand $C_{Nt} = Y_{Nt}$ given $P_{Nt} = 1$. We think of this as the limiting case of a Calvo economy where probability of price nonadjustment $\nu \to 1$ and the conventional New Keynesian Phillips curve for nontradable price inflation, $\pi_{Nt} = \beta \mathbb{E}_t \pi_{Nt} + \lambda \log\left(W_t / A_{Nt}\right)$, is degenerate with $\lambda \equiv \frac{(1-\nu)(1-\beta\nu)}{\nu} \to 0$ and $\pi_{Nt} = \Delta \log P_{Nt} = 0$ independently of the level of nominal marginal cost W_t / A_{Nt}.

This combination of stark assumptions—on the functional form of the utility, the endowment of homogenous tradables with the law of one price, and the permanent stickiness of nontradable prices—yields simple closed form solutions yet does not comprise the main qualitative properties of more general models, as we confirm in the other papers.

Government. The government sets domestic interest rate R_t by trading home-currency bond B_t with the households, and it returns the revenues from financial intermediation in the foreign-currency bond back to the households, $T_t = \left(e^{-\hat{\psi}_t} - 1\right) \frac{\mathcal{E}_t B_t^*}{R_t^*}$. Combining (5) with (3), we have $\beta R_t \mathbb{E}_t \left\{ W_t / W_{t+1} \right\} = 1$, and thus the choice of R_t is equivalent to the choice of wage inflation, or the path of wages $\{W_t\}$.[3]

The first-best allocation in the nontradable sector requires $W_t / P_{Nt} = A_{Nt}$, and thus given sticky price $P_{Nt} \equiv 1$, the first-best nominal wage must track nontradable productivity, $\tilde{W}_t = A_{Nt}$. The realized wage can thus be written as $W_t = A_{Nt} X_t$, where X_t is the output gap induced by monetary policy ($X_t = 1$ corresponds to no output gap). We think of X_t as the monetary shock in the economy.

Substituting T_t and Π_t into the household budget constraint (2), and using the nontradable market clearing $C_{Nt} = Y_{Nt}$, the fact that home-currency bond is in zero net supply domestically, and the law of one price for tradables, we can write the home-country budget constraint in foreign-currency terms as follows:

$$\frac{B_t^*}{R_t^*} - B_{t-1}^* = P_{Tt}^* \left(Y_{Tt} - C_{Tt} \right), \tag{6}$$

where the right-hand side is home net exports in foreign-currency terms. $\{R_t^*, P_{Tt}^*\}$ correspond to foreign shocks in the financial and goods

3. Note from (5) that W_t corresponds to nominal nontradable expenditure $P_{Nt} C_{Nt}$, which is controlled by monetary policy.

markets. For simplicity, we shut them down and study the case with $P^*_{T_t} \equiv 1$ and $\beta R^*_t \equiv 1$, focusing on the productivity shocks (A_{Nt}, A_{Tt}) and monetary shocks X_t, as well as the risk-sharing wedge $\hat{\psi}_t$.

Equilibrium. The equilibrium in the nontradable sector is characterized by the labor supply condition (5) given sticky prices $P_{Nt} = 1$ and the market clearing $C_{Nt} = Y_{Nt} = A_{Nt} L_t$. We thus have:

$$Y_{Nt} = C_{Nt} = \frac{W_t}{P_{Nt}} = A_{Nt} X_t \text{ and } L_t = X_t. \tag{7}$$

The equilibrium in the tradable sector is an interplay of three equilibrium conditions—the expenditure switching between tradables and nontradables, the country budget constraint, and the foreign-currency Euler equation. The expenditure switching condition is the result of optimal expenditure on tradables and nontradables, and we rewrite it as:

$$\frac{\gamma}{1-\gamma}\frac{C_{Nt}}{C_{Tt}} = \frac{\mathcal{E}_t}{P_{Nt}}, \tag{8}$$

where we use the fact that $P_{Tt} = \mathcal{E}_t$ given the law of one price with the international price of tradables $P^*_{Tt} = 1$. Thus, shifts in nominal exchange rate, given sticky nontradable prices P_{Nt}, relocate expenditure between tradable and nontradable inputs of final consumption.

Finally, we rewrite the country budget constraint (6) and the Euler equation (4) as:

$$\beta B^*_t - B^*_{t-1} = Y_{Tt} - C_{Tt},$$
$$\mathbb{E}_t \frac{C_{Tt}}{C_{Tt+1}} = e^{\hat{\psi}_t},$$

where we used the facts that $\beta R^*_t = 1$ and $P^*_{T_t} = 1$. This system characterizes the solution for $\{C_{Tt}\}$, which we partition by analogy with nontradable consumption as

$$C_{Tt} = A_{Tt} Z_t, \tag{9}$$

where $\tilde{C}_{Tt} = A_{Nt}$ is approximately optimal path of tradable inputs in the absence of financial wedges $\hat{\psi}_t = 0$ (assuming $B^*_{-1} = 0$), while Z_t reflects the additional volatility in tradables due to wedges in the international financial market.

With this, we can write the equilibrium exchange rate as:[4]

$$\frac{\mathcal{E}_t}{P_{Nt}} = \frac{\gamma}{1-\gamma}\frac{A_{Nt}}{A_{Tt}}\frac{X_t}{Z_t}, \tag{10}$$

and the approximate expression for Z_t given by:

$$\Delta \log Z_t = -\frac{\beta}{1-\beta\rho}\left(\hat{\psi} - \frac{1}{\beta}\hat{\psi}_{t-1}\right), \tag{11}$$

assuming that $\hat{\psi}_t$ follows an AR(1) with persistence $\rho \in [0,1]$.[5] Equations (7)–(11) fully characterizes equilibrium in this economy where $\{A_{Nt}, A_{Tt}, X_t, \hat{\psi}_t\}$ are exogenous shocks.

Macroeconomic aggregates. We can now characterize macroeconomic aggregates in this economy—inflation (consumer price level), aggregate consumption, real GDP, employment, aggregate wage rate, and the real exchange rate. We express these macroeconomic aggregates as a function of exogenous shocks $\{A_{Nt}, A_{Tt}, X_t\}$ and the nominal exchange rate $\mathcal{E}_t$, which we characterized above.

In particular, consumer price level is given by $P_t = P_{Nt}^{1-\gamma}\mathcal{E}_t^{\gamma}$, where the two terms reflect the nontradable and tradable price inflation. Using the expenditure allocation condition and nontradable market clearing, we express aggregate consumption and real GDP as follows:

$$C_t = \frac{P_{Nt}C_{Nt}}{(1-\gamma)P_t} = \left(\frac{P_{Nt}}{\mathcal{E}_t}\right)^{\gamma}\frac{A_{Nt}X_t}{1-\gamma},$$

$$Y_t = \frac{P_{Nt}Y_{Nt} + P_{Tt}Y_{Tt}}{P_t} = \left(\frac{P_{Nt}}{\mathcal{E}_t}\right)^{\gamma} A_{Nt}X_t + \left(\frac{\mathcal{E}_t}{P_{Nt}}\right)^{1-\gamma} A_{Tt}.$$

4. See interpretation below following (16).

5. This solution relies on the fact that $Y_{Tt} = A_{Tt}$ follows a random walk and log-linearly approximates the equilibrium system around $B_t^* = 0$, which yields two dynamic equations (with $b_t^* \equiv B_t^*/Y_{T0}$):

$$\beta b_t^* - b_{t-1}^* = \mathrm{dlog}Y_{Tt} - \mathrm{dlog}C_{Tt} = -\,\mathrm{dlog}Z_t,$$

$$\hat{\psi}_t = \mathbb{E}_t\Delta\log C_{Tt+1} = \mathbb{E}_t\Delta\log Z_{t+1},$$

where we use the facts that $\log Y_{Tt} = \log A_{Tt}$ is a random walk (i.e., $\mathbb{E}_t\Delta\log A_{t+1} = 0$) and $\log Z_t = \log C_{Tt} - \log A_{Tt}$. Solving this dynamic system with $\hat{\psi}_t \sim AR(1)$ yields $\mathrm{dlog}\,Z_t = (1-\beta)b_{t-1}^* - \frac{\beta}{1-\beta\rho}\hat{\psi}_t$, which then results in (11).

This allocation is supported with aggregate employment level $L_t = X_t$ given aggregate wage rate $W_t / P_{Nt} = A_{Nt} X_t$. Finally, the real exchange rate in this economy is given by:

$$\mathcal{Q}_t = \frac{P_t^* \mathcal{E}_t}{P_t} = \left(\frac{\mathcal{E}_t}{P_{Nt}} \right)^{1-\gamma},$$

where we assume $P_t^* = P_{Tt}^* = 1$. We kept P_{Nt} in the expressions above to illustrate how the results would generalize to a model where sticky prices P_{Nt} are allowed to adjust in response to output gap X_t.

We can now rewrite this macro quantities in log changes (growth rates), which by convention we denote with corresponding small letters (with the exception of inflation denoted with π_t):[6]

$$\pi_t = (1-\gamma)\pi_{Nt} + \gamma e_t, \tag{12}$$

$$c_t = a_{Nt} + x_t - \gamma(e_t - \pi_{Nt}), \tag{13}$$

$$y_t = (1-\gamma)(a_{Nt} + x_t) + \gamma a_{Tt}, \tag{14}$$

$$q_t = (1-\gamma)(e_t - \pi_{Nt}), \tag{15}$$

where $\pi_{Nt} = 0$ under fully sticky prices and more generally satisfies the dynamic Phillips curve $\pi_{Nt} = \beta \mathbb{E}_t \pi_{Nt+1} + \lambda \log X_t$ given the path of output gap X_t chosen by monetary policy. We assume the economy is subject to random-walk productivity and monetary shocks such that (a_{Tt}, a_{Nt}, x_t) are *idd* as growth rate shocks. Finally, the nominal exchange rate in (10) follows:

$$e_t = (a_{Nt} - a_{Tt}) + (\pi_{Nt} + x_t) - z_t \text{ where } z_t = -\frac{\beta}{1-\beta\rho}\left(\hat{\psi}_t - \frac{1}{\beta}\hat{\psi}_{t-1} \right). \tag{16}$$

Since $\hat{\psi}_t \sim AR(1)$, $z_t \sim ARMA(1,1)$ with autoregressive root ρ and moving average root $1/\beta$. When β, $\rho \approx 1$, this growth rate process is arbitrary close to white noise, so that the exchange rate is close to a random walk (recall that $e_t \equiv \Delta \log \mathcal{E}_t$), consistent with its empirical properties.

6. For real GDP, we approximate around balanced trade, so that $P_{Nt}Y_{Nt}$ and $P_{Tt}Y_{Tt}$ correspond to fraction $1-\gamma$ and γ of nominal GDP respectively (like consumption expenditure shares). Given this, the effects of the exchange rate on real GDP (via inflation and relative price of tradables) cancel out.

Before using these results to analyze a range of exchange rate puzzles, we offer a brief commentary. First, the nominal exchange rate in (16) has three components: (1) Balassa-Samuelson term $\tilde{q}_t \equiv a_{Nt} - a_{Tt}$ reflecting equilibrium pressures on the relative nontradable prices;[7] (2) nominal inflationary pressure $\pi_{Nt} + x_t$, which emerges from the output gap x_t under sticky prices, and then from price inflation π_{Nt} if they adjust; (3) financial shocks captured by z_t (i.e., relative demand shocks for foreign currency $\hat{\psi}_t$ causing home-currency depreciation). The relative nontradable prices evolve with $e_t - \pi_{Nt}$, which shapes the equilibrium dynamics of the real exchange rate q_t in (15).

What concerns macro aggregates (12)–(14), domestic consumer price inflation π_t, reflects nontradable and tradable inflation π_{Nt} and et with weights $(1 - \gamma)$ and γ respectively. Aggregate consumption evolves with productivity a_{Nt} and output gap x_t, as well as responds to the expenditure switching force due to the relative nontradables price with elasticity γ. In contrast, real GDP reflects relative productivities in the two sectors with weights $(1 - \gamma)$ and γ respectively, as well as responds to the output gap, which shapes aggregate employment in the economy. These are conventional macroeconomic forces typical in standard business-cycle models, and the only unconventional feature of the model is the presence of financial shocks z_t that affect the equilibrium exchange rate.

2. Exchange Rate Puzzles

2.1 Puzzles under Floating Exchange Rate

Backus-Smith. At the core of understanding the exchange rate under floating regime is the Backus-Smith puzzle.[8] While under complete asset markets and separable utility with risk aversion σ, the real exchange rate must satisfy $q_t = \sigma\,(c_t - c_t^*)$, in the data real exchange depreciations (increases in q_t) are associated with reductions in relative home consumption $(c_t - c_t^*)$, albeit with a weak correlation

7. See Obstfeld and Rogoff (1996), chapter 4.
8. See Backus and Smith (1993) and Kollmann (1995).

(see figure 2a below). The equilibrium conditions (13) and (15) provide an insight into this puzzle, as we can calculate:

$$\frac{\operatorname{cov}(c_t, q_t)}{\operatorname{var}(q_t)} = -\frac{\gamma}{1-\gamma} + \frac{1}{1-\gamma}\frac{\operatorname{cov}(a_{Nt} + x_t, e_t - \pi_{Nt})}{\operatorname{var}(e_t - \pi_{Nt})}$$

where $e_t - \pi_{Nt}$ is given by (16) and we assume $c_t^* = 0$ in line with our small open economy approach.

The first term reflects expenditure switching—a decline in consumption driven by a real depreciation (an increase in the relative price of foreign tradables)—and its effect is proportional to the openness of the economy to foreign tradables γ.[9] The second term reflects the comovement of the domestic component of consumption with the real exchange rate and equals the combined variance contribution of productivity shocks a_{Nt} and monetary shocks x_t (output gap) to the variance of the real exchange rate $q_t = (1-\gamma)(e_t - \pi_{Nt})$. This effect does not depend on the openness of the economy γ.

The decomposition above makes it clear what features of the model result in the Backus-Smith puzzle. Note that it is not about completeness of asset markets, as we assumed incomplete markets from the get-go. In fact, if monetary shocks x_t and/or productivity shocks a_{Nt} are the key drivers of the real exchange rate, then $\frac{\operatorname{cov}(a_{Nt} + x_t, e_t - \pi_{Nt})}{\operatorname{var}(e_t - \pi_{Nt})} \approx 1$ and thus $\frac{\operatorname{cov}(c_t, q_t)}{\operatorname{var}(q_t)} \approx 1$ irrespectively of asset market incompleteness and the openness of the economy γ. As a result, the persistence of the Backus-Smith puzzle is due to the fact that international Real Business Cycle (RBC) and New Keynesian models alike robustly reproduce it independently of the many features of such models as long as productivity and monetary shocks are the key driving forces in the economy.

What is the explanation for the Backus-Smith puzzle? It requires financial exchange rate shocks z_t to be the key driver of the nominal exchange rate in (16).[10] If this is the case, and z_t is

9. We write foreign tradables here since in a more general model with imperfectly substitutable home and foreign tradables, what matters for expenditure switching is the relative price of foreign tradables in the home market and their share in total consumption expenditure. See Itskhoki and Mukhin (2021a) and Itskhoki (2021).

10. In our simple model, it is also possible to explain the Backus-Smith puzzle if the key driver of the exchange rate is the homogenous tradable endowment shock a_{Tt}. This shock, however, is at odds with other exchange rate puzzles, in particular the exchange rate disconnect puzzle that we discuss next.

largely orthogonal with monetary and productivity shocks, then $\frac{\text{cov}(a_{Nt} + x_t, e_t - \pi_{Nt})}{\text{var}(e_t - \pi_{Nt})} \approx 0$ and thus $\frac{\text{cov}(c_t, q_t)}{\text{var}(q_t)} \approx -\frac{\gamma}{1-\gamma}$, consistent with the weak negative correlation in the data. Quantitatively, we show in Itskhoki and Mukhin (2021a) that financial shocks should account for around 80–90 percent of the nominal exchange rate volatility for the model to be quantitatively consistent with the Backus-Smith correlation in the data, given that most countries exhibit significant home bias and have a large nontradable share.

The purchasing power parity (PPP) puzzle. The PPP puzzle emphasizes the fact that the real exchange rate closely tracks the nominal exchange rate at most frequencies, inheriting both its volatility and persistence.[11] From the definition of the real exchange rate, this implies that inflation π_t is small and largely uncorrelated with exchange rate changes e_t. From (12) and (15), we see that the model is consistent with PPP puzzle if monetary inflation shocks π_{Nt} are small in the variation of the nominal exchange rate (16), and home bias is large (tradable share γ is small). In fact, the real and nominal exchange rates follow an equally persistent near-random walk process if financial shocks z_t are the main source of their volatility.

The simple model presented here is special as it assumes that the law of one price holds for a homogenous tradable good. In a more realistic model with home bias in imperfectly substitutable tradable goods and law-of-one-price violations due to sticky local-currency prices, the real exchange rate q_t perfectly traces the nominal exchange rate e_t even when $\gamma \gg 0$, as long as the volatility in the exchange rates is not due to monetary shocks.[12] The reason is that monetary policy can act to effectively stabilize consumer price inflation, while the nominal and real exchange rates are volatile and persistent in response to financial shocks z_t.[13]

Meese-Rogoff disconnect puzzle. Another crucial property of the nominal (and real) exchange rate is that it is largely uncorrelated with a whole range of macroeconomic fundamentals, both nominal and real, and tends to be an order of magnitude more volatile than

11. See Rogoff (1996) and Appendix figure A1a,b.

12. See Itskhoki and Mukhin (2021a), Eichenbaum and others (2021), Blanco and Cravino (2020).

13. Additionally, in the data, the wage-based real exchange rate tracks closely the nominal exchange rate. Given that $w_t = \pi_{Nt} + a_{Nt} + x_t$ and again assuming $w_t^* = 0$, we have $q_t^w = w_t^* + e_t - w_t = -(z_t - a_{Tt})$, which tracks e_t and q_t provided that z_t is the main source of variation.

various macroeconomic aggregates.[14] Figure 1 below and Appendix figure A1 illustrate the order-of-magnitude difference in the volatility of exchange rates and macroeconomic fundamentals under the floating exchange rate regime. Since we have already studied the exchange rate comovement with consumption and inflation above, we now focus on the real GDP given by (14):[15]

$$\frac{\operatorname{cov}(y_t, e_t)}{\operatorname{var}(e_t)} = \frac{\operatorname{cov}\big((1-\gamma)(a_{Nt} + x_t) + \gamma a_{Tt}, e_t\big)}{\operatorname{var}(e_t)}.$$

As long as productivity and monetary shocks (a_{Nt}, a_{Tt}, x_t) account for a small share of variation in the nominal exchange rate (16), which in turn is mostly driven by financial shocks z_t, the correlation between the nominal exchange rate and the real GDP is arbitrarily close to zero, while their relative volatility is arbitrarily large, in line with disconnect properties.

Note that this does not mean that conventional macroeconomic shocks (a_{Nt}, a_{Tt}, x_t) are absent. In contrast, they are essential to ensure the conventional business-cycle dynamics of consumption, output and inflation. However, their relative contribution to the large exchange rate volatility is limited, as asset demand shocks $\hat{\psi}_t$ result in a considerably more volatile source of exchange rate fluctuations z_t, in particular when β and ρ are close to 1 in (16). These shocks feed back into macro dynamics via the expenditure switching effect on consumption in (13), which is proportionally small with the openness of the economy γ. More open economies exhibit both less volatile equilibrium exchange rates and less exchange rate disconnect—consistent with the model with imperfectly substitutable home and foreign tradable, as we show in Itskhoki and Mukhin (2021a).

UIP puzzle. Lastly, we turn to the forward premium puzzle, which emphasizes systematic UIP violations, namely that returns on a currency carry trade $R_t - R_t^* \frac{\mathcal{E}_{t+1}}{\mathcal{E}_t}$ co-move systematically with

14. See e.g., Meese and Rogoff (1983).

15. Similarly, we could focus on aggregate employment $(\ell_t = x_t)$ or nominal expenditure (e.g., $P_{Nt}C_{Nt} = P_{Nt}A_{Nt}X_t$).

the interest-rate differential $R_t - R_t^*$.[16] Combining together the two household Euler equations (3)-(4) and log-linearizing results in:[17]

$$i_t - i_t^* - \mathbb{E}_t e_{t+1} = \hat{\psi}_t, \tag{17}$$

where $i_t - i_t^* = \log R_t - \log R_t^*$ and $e_{t+1} = \Delta \log \mathcal{E}_{t+1}$. The foreign-currency demand shock $\hat{\psi}_t$ results in UIP deviations, which can be either long-run mean zero or nonzero, as is the case for developed versus developing countries.[18] The Fama regression, however, emphasizes that $i_t - i_t^* = \mathbb{E}_t e_{t+1}$ systematically increases with $i_t - i_t^*$, or in other words e_{t+1} tends to be negative (appreciate) when $i_t - i_t^*$ increases, albeit with a vanishingly small predictive power (i.e., $R^2 \approx 0.01$).

Our simple models predicts that the coefficient in the Fama regression of e_{t+1} on $i_t - i_t^*$ is indeterminant and the R^2=0, independently of the presence or absence of $\hat{\psi}_t$. This is because $i_t^* = 0$ and $i_t = \mathbb{E}_t \left\{c_{t+1} + \pi_{t+1}\right\} = \mathbb{E}_t \left\{a_{Nt+1} + \pi_{Nt+1} + x_{t+1}\right\} = 0$ under random walk shocks. This emphasizes the weak identifying power of the Fama regression. If we run a real version of the Fama regression of q_{t+1} on $r_t - r_t^*$, where $r_t = i_t - \mathbb{E}_t \pi_{t+1}$, we identify a negative coefficient as we regress $(1-\gamma)\, e_{t+1}$ on $-\gamma \mathbb{E}_t e_{t+1}$, provided that z_t shocks account for some variation in e_t. The R^2 of this regression would still be close to zero—capturing the robust empirical property of the Fama regression.

2.2 Mussa and Other Puzzles

The Mussa (1986) puzzle concerns the switch from a pegged to a floating nominal exchange rate regime, which empirically is associated with a dramatic increase in the volatility of both nominal and real exchange rates, yet little change in the properties of other macroeconomic variables. We illustrate this in figure 1 and Appendix figure A1, which show a dramatic increase in the volatility of both nominal and real exchange rates immediately after the end of Bretton Woods, while the behavior of consumption, real GDP, and inflation did not experience any discernible discontinuity around this breakpoint.

16. See Fama (1984) and figure 2b below.

17. The nonlinear condition is $\mathbb{E}_t \left\{ \frac{W_t}{W_{t+1}} \left[R_t e^{\hat{\psi}_t} - R_t^* \frac{\mathcal{E}_{t+1}}{\mathcal{E}_t} \right] \right\} = 0$, where we use the fact that optimal expenditure $P_{Tt} C_{Tt} \propto P_{Nt} C_{Nt} = W_t$. The higher order term, thus, depends on $\text{cov}_t\left(W_{t+1}^{-1}, \mathcal{E}_{t+1}\right)$, which is close to zero both in the data and in the model that satisfies the disconnect property, as discussed above.

18. See e.g., Hassan and Mano (2018) and Kalemli-Özcan and Varela (2021).

The simple model above allows us to investigate what features are necessary for a model to match these empirical patterns. The expression for the equilibrium nominal exchange rate in (16) shows how a change in monetary policy x_t can accommodate a fixed nominal exchange rate. Indeed, setting $x_t = z_t - \pi_{Nt} - (a_{Nt} - a_{Tt})$ ensures $e_t = 0$, while we think of a floating exchange rate regime as output-gap stabilization with $x_t = 0$ and $e_t = (a_{Nt} - a_{Tt}) + \pi_{Nt} - z_t$. Such switch in monetary policy has dramatic consequences for macroeconomic quantities.

On the one hand, under fully sticky prices with $\pi_{Nt} = 0$, we have $q_t = (1 - \gamma)e_t$, and indeed a change in the volatility of the nominal exchange rate induces a proportional change in the volatility of the real exchange rare q_t, in line with the empirical patterns. Note that this property of the model is independent of the nature of the shocks driving the exchange rate and is the consequence of stable inflation under both regimes, which empirically indeed remained stable even after a switch to volatile floating exchange rates.

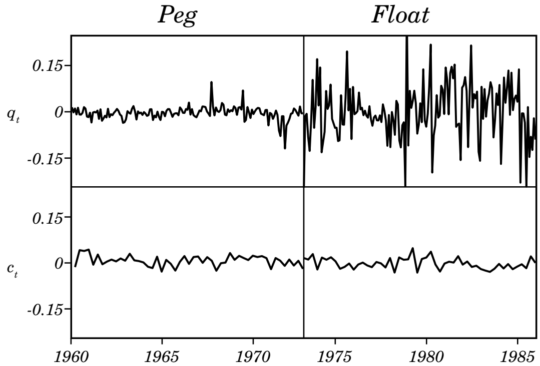
On the other hand, in contrast with the data, such change in monetary policy has equally large consequences for macroeconomic aggregates. We focus, for example, on real GDP and the results for aggregate consumption are similar:

$$y_t = \begin{cases} (1-\gamma)a_{Nt} + \gamma a_{Tt}, & \text{under float} \\ a_{Tt} + z_t, & \text{under peg.} \end{cases}$$

That is, under the float, the real GDP reflects average productivity of the economy given the stabilized output gap, while the financial shock z_t—the key drivers of the exchange rate (see above)—does not affect GDP, as it is absorbed by the exchange rate.[19] In contrast, under the peg, both real GDP and aggregate consumption reflect one-to-one financial shocks z_t, irrespectively of the openness of the economy. This is because monetary policy needs to absorb exchange rate shocks and thus pass on financial shocks into fluctuations of the output gap x_t, which affects employment, consumption, and output independently of the openness of the economy. This is in sharp contrast with the empirical Mussa patterns shown in figure 1.

19. The exact orthogonality of the real GDP with e_t (and thus with z_t) is a knife-edge implication of the Cobb-Douglas utility and other special assumption of our model; more generally, the real GDP is exposed to the exchange rate fluctuations, like aggregate consumption, with an elasticity proportional to the openness of the economy γ. Imperfectly substitutable tradable goods and local currency price stickiness of exports further mute this transmission along with low γ. See Itskhoki and Mukhin (2021a,b).

Figure 1. Real Exchange Rate and Aggregate Consumption during and after Bretton Woods



Source: Itskhoki and Mukhin (2021b).
Note: Monthly real exchange rate changes q_t (G7 countries plus Spain, without Canada against the U.S.) and quarterly aggregate consumption growth rates c_t (average for G7 countries); both series annualized and in log points (that is, 0.15 corresponds to 15 log points, approximately 15%). The breakup of Bretton Woods is dated 1973.1. See also Appendix figure A1.

Itskhoki and Mukhin (2021b) show that in a large class of conventional business-cycle models there exists a robust sufficient statistic $\sigma\,(c_t - c_t^*) - q_t$, where σ is risk aversion, that does not change its statistical properties with a switch in the monetary regime, even if consumption and the real exchange rate change their behavior. Indeed, this is the case in the model presented here with $\sigma = 1$:

$$\left(c_t - c_t^*\right) - q_t \quad = a_{Nt} + x_t - \gamma\left(e_t - \pi_{Nt}\right) - \left(1-\gamma\right)\left(e_t - \pi_{Nt}\right) = a_{Tt} + z_t,$$

where we used (13), (15), and (16). So long as the endowment shock a_{Tt} and the financial shock z_t do not change their properties across monetary regimes, changes in monetary policy x_t and the associated changes in the behavior of exchange rates do not affect this sufficient statistic. In the data, however, $(c_t - c_t^*) - q_t$ dramatically changes its behavior along with q_t following a switch between a peg and a float.

Resolution. The result above suggests a feature of the model that can lead to a resolution of the Mussa puzzle. Indeed, it requires that some shocks change their properties with a change in a monetary regime. In particular, in Itskhoki and Mukhin (2021b), we show that the volatility of financial shocks has to be endogenous to the exchange rate regime and, specifically, increasing in the equilibrium exchange rate volatility:[20]

$$\hat{\psi}_t = \chi\left(\sigma_e^2\right)\cdot\psi_t, \text{ where } \chi(0)=0,\ \chi'(\cdot)>0 \text{ and } \sigma_e^2 = \mathrm{var}_t\left(e_{t+1}\right). \tag{18}$$

In the following section, we describe a microfoundation for such an endogenous change in $\hat{\psi}_t$, which is also essential for the optimal policy analysis.

Under (18), an exchange rate peg with $\sigma_e^2 = 0$ results in $\hat{\psi}_t = 0$ and consequently $z_t = 0$, eliminating financial shocks as a driver of both nominal and real exchange rates. Recall that exchange rate disconnect under the float requires that financial shocks z_t are the key drivers of the floating exchange rates, explaining the dramatic shift in their volatility with the exchange rate regime, as observed in figure 1 and Appendix figure A1.

As the exchange rate changes from $e_t = (a_{Nt} - a_{Tt}) - z_t$ to $e_t = 0$, the output gap needs to change only from $x_t = 0$ to $x_t = a_{Nt} - a_{Tt}$ to accommodate a switch to a peg. To the extent z_t accounts for the bulk of the exchange rate variation under the float and $a_{Nt} - a_{Tt}$ are (relatively) stable, this requires only a minor change in monetary policy. Consequently, the real GDP and aggregate consumption also change only mildly, e.g., from $y_t = (1 - \gamma)\, a_{Nt} + \gamma a_{Tt}$ under the float to $y_t = a_{Tt}$ under the peg. This explains why we do not observe a major breakpoint in the behavior of these macroeconomic aggregates.

Home bias in consumption and nontradables (lowγ) shield macroeconomic aggregates from exchange rate volatility under the float, as we discussed above. More importantly, however, endogenous financial volatility in (18) shields monetary policy and consequently macroeconomic aggregates from financial volatility under the peg. Without this, monetary policy would need to absorb volatile financial shocks to stabilize the exchange rate, and consequently pass on this volatility into inflation, consumption, and output, irrespectively of the openness of the economy.

Other puzzles. Consider three exchange rate puzzles that change their properties with the exchange rate regime. First, consider Balassa-Samuelson that suggests that the real exchange rate should evolve with relative nontradable productivity $a_{Nt} - a_{Tt}$. Indeed, we have:

$$q_t = (1 - \gamma)[(a_{Nt} - a_{Tt}) + x_t - z_t].$$

20. See also Kollmann (2005).

Thus to the extent z_t dominates the volatility of exchange rates under the float, it is difficult to isolate the Balassa-Samuelson forces from the time series properties of the real exchange rate. In contrast, if z_t disappears under the peg without a change in monetary policy $x_t = 0$, then q_t is shaped entirely by the Balassa-Samuelson forces under the peg, for which there is indeed empirical evidence since the introduction of euro.[21]

Second, UIP holds considerably better under the peg than under the float, in line with (17), provided that $\hat{\psi}_t$ has an endogenously reduced volatility under the peg. Furthermore, the negative sign of the Fama regression coefficient persistent under the float, either turns zero or becomes positive under the peg, closer to the theoretical benchmark (see figure 2b). Similarly, the Backus-Smith correlation turns from negative to positive under the peg, which is again in line with the Mussa mechanism. We rewrite the Backus-Smith covariation in the model as follows:

$$\frac{\mathrm{Cov}\left(c_t, q_t\right)}{\mathrm{var}\left(q_t\right)} = -\frac{\gamma}{1-\gamma} + \frac{1}{1-\gamma}\frac{\mathrm{var}\left(a_{Nt} - a_{Tt}\right) + \mathrm{cov}\left(a_{Tt}, a_{Nt} - a_{Tt}\right)}{\mathrm{var}\left(a_{Nt} - a_{Tt}\right) + \mathrm{var}\left(z_t\right)}$$

where we assumed for simplicity that $x_t = 0$ under both the float and the peg and that the financial shock z_t is orthogonal with productivity $(a_{Nt} - a_{Tt})$.[22] If z_t is the dominant shock under the float, then the Backus-Smith covariation is mildly negative, as in the data. If the variance of z_t declines towards zero under the peg, the Backus-Smith covariance increases and turns positive, provided a_{Tt} and a_{Nt} are not strongly negatively correlated. This is again consistent with the data, as we show in figure 2a.[23]

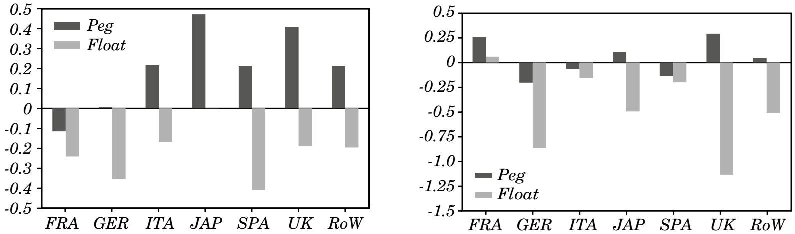
21. See Berka and others (2012, 2018).

22. In our simple model, a monetary policy x_t that fully stabilizes nominal exchange rate also fully stabilizes the real exchange rate, and thus the Backus-Smith moment we focus on is zero or indeterminant under the peg. More generally, the real exchange rate reflects relative inflation under the peg, which is nonzero (see Appendix figure A1), and the Backus-Smith correlation is well-defined in the data, consistent with the description we offer in the text.

23. See also Devereux and Hnatkovska (2020) and Colacito and Croce (2013).

Figure 2. Backus-Smith Correlation and Fama Coefficient before and after the End of Bretton Woods

(a) *Backus-Smith correlations* (b) *Fama coefficient,* β_F



Source: Itskhoki and Mukhin (2021b).
Note: The left panel displays the Backus-Smith correlation, corr $(c_t - c_t^*, q_t)$ in growth rates, using annual data for 1960–71 for Peg and 1973–1989 for Float. The right panel displays Fama regression coefficient β_F, obtained from an OLS regression of depreciation rate $e_{(t+1)}$ on $(i_t - i_t^*)$, using monthly data for 1960.1–1971.7 for Peg and 1973.1–1989.12 for Float. G7 countries (plus Spain, without Canada) against the United States.

3. Exchange Rate Policies

Two key features are essential for the model to be consistent with the combined empirical properties of exchange rates. First, financial shocks $\hat{\psi}_t$ must account for the bulk of exchange rate volatility under a floating regime. A range of models of the international financial market can give rise to such shocks.[24] Second, the evidence on the switch of the floating regime to an exchange rate peg further requires that the volatility of these financial shocks endogenously decreases with a reduction in equilibrium exchange rate volatility, that is $\hat{\psi}_t = \chi(\sigma_e^2)\,\hat{\psi}_t$, where $\chi(\cdot) = 0$ is an increasing function of exchange rate volatility $\sigma_e^2 = \text{var}_t\,(e_{t+1})$. We next describe a micro-founded model for this reduced form, which then allows us to proceed with the analysis of the optimal exchange rate policies.

24. Exogenous UIP shocks are commonly used in the international macro literature (see e.g., Devereux and Engel, 2002; Kollmann, 2005; Farhi and Werning, 2012), and can be viewed to emerge from exogenous asset demand, as in the literature following Kouri (1976, 1983). Models of UIP deviations include models with incomplete information, expectational errors and heterogeneous beliefs (Evans and Lyons, 2002; Gourinchas and Tornell, 2004; Bacchetta and van Wincoop, 2006), financial frictions (Gabaix and Maggiori, 2015; Adrian and others, 2015; Camanho and others ,2018), liquidity premia (Jiang and others, 2021; Bianchi and others, 2021), habits, long-run risk, and rare disasters (Verdelhan, 2010; Colacito and Croce, 2013; Farhi and Gabaix, 2016), and alternative formulations of segmented markets (Jeanne and Rose, 2002; Alvarez and others, 2009).

3.1 A Model of the Financial Market

The general modeling environment is the same as in section 1, with the only difference that household do not have direct access to the foreign-currency (dollar) bond, i.e., $B_t^* \equiv 0$ and the Euler equation (4) no longer applies. The households can only save and borrow using the home-currency bond B_t with interest rate R_t according to the optimality condition (3). In addition, we introduce an explicit model of the financial market which intermediates international capital flows.

Apart from the households, three types of agents trade home- and foreign-currency bonds in the international financial market. Namely, these are the government, noise traders and arbitrageurs. The government holds a portfolio of (F_t, F_t^*) units of home- and foreign-currency bonds, respectively, with the value of the portfolio (government net foreign assets) given by $F_t/R_t + \mathcal{E}_t F_t^*/R_t^*$. Changes in F_t and F_t^* correspond to open market operations of the government.

Noise traders hold a zero capital portfolio (N_t, N_t^*) of the two bonds, such that $N_t/R_t + \mathcal{E}_t N_t^*/R_t^* = 0$, and $N_t^*/R_t^* = \hat{\psi}_t$ is the liquidity demand for foreign currency by the noise traders, that is $\hat{\psi}_t$ is a random variable uncorrelated with macroeconomic fundamentals. A positive $\hat{\psi}_t$ means that noise traders short home-currency bonds to buy foreign-currency bonds, and vice versa.

Finally, the arbitrageurs also hold a zero capital portfolio (D_t, D_t^*) such that $D_t/R_t + \mathcal{E}_t D_t^*/R_t^* = 0$, with a return on one foreign-currency unit holding of such portfolio given by $\tilde{R}_{t+1}^* = R_t^* - R_t \frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}$ in dollars. In other words, the income from this carry trade is given by $\pi_{t+1}^{D*} = D_t^* - \frac{D_t}{\mathcal{E}_{t+1}} = \tilde{R}_{t+1}^* \cdot \frac{D_t^*}{R_t^*}$ in foreign currency, where we used the zero-capital constraint linking D_t and D_t^*. Arbitrageurs choose their portfolio (D_t, D_t^*) to maximize min-variance preferences over profits, $V_t\left(\pi_{t+1}^{D*}\right) = \mathbb{E}_t\left\{\Theta_{t+1}\pi_{t+1}^{D*}\right\} - \frac{\omega}{2}\mathrm{var}_t\left(\pi_{t+1}^{D*}\right)$, where $\Theta_{t+1} = \beta\frac{C_{Tt}}{C_{Tt+1}}$ is the stochastic discount factor of home households, and the second term in $V_t(\cdot)$ reflects the additional risk penalty of the arbitrageurs with ω being the risk aversion parameter. The optimal portfolio choice satisfies:

$$\frac{D_t^*}{R_t^*} = \frac{\mathbb{E}_t\left\{\Theta_{t+1}\tilde{R}_{t+1}^*\right\}}{\omega\sigma_t^2},$$

where $\sigma_t^2 \equiv \mathrm{var}_t\left(\tilde{R}_{t+1}^*\right) = R_t^2 \cdot \mathrm{var}_t\left(\frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}\right)$ measures the carry-trade risk which is associated with the nominal exchange rate volatility.

The market clearing in the financial market requires that the home-currency bond positions of all four types of agents balance out:

$$B_t + N_t + D_t + F_t = 0.$$

The foreign-currency bond is in perfect elastic international supply at an exogenous interest rate R_t^*.

The government budget constraint from operations in the financial market is given by:

$$\frac{F_t}{R_t} + \frac{\mathcal{E}_t F_t^*}{R_t^*} = F_{t-1} + \mathcal{E}_t F_{t-1}^* + \tau \mathcal{E}_t \pi_t^* - T_t \;\text{ with }\; \pi_t^* = \tilde{R}_t^* \cdot \frac{N_{t-1}^* + D_{t-1}^*}{R_{t-1}^*},$$

where T_t is the lump-sum transfer to the home households and π_t^* is the combined income from the financial transactions of noise traders and arbitrageurs (in dollars). Note that parameter $\tau \in [0,1]$ can be viewed as either the home country's ownership share of the financial sector or a tax on financial transactions imposed by the home government.[25]

Equilibrium. Define the net foreign asset (NFA) position of the home country, B_t^* in foreign currency, which has the home-currency value:

$$\frac{\mathcal{E}_t B_t^*}{R_t^*} = \frac{B_t + F_t}{R_t} + \frac{\mathcal{E}_t F_t^*}{R_t^*},$$

that is the value of the combined position of the home households and the government. Using B_t^*, we prove in Appendix B the following lemma that characterize the open economy equilibrium conditions.

Lemma 1. *The NFA of the home country equals the combined foreign-currency bond position in the financial market, $B_t^* = F_t^* + N_t^* + D_t^*$, and the combined home-country budget constraint in foreign-currency terms is given by:*

$$\frac{B_t^*}{R_t^*} - B_t^* = \left(Y_{Tt} - C_{Tt}\right) - \left(1 - \tau\right) \tilde{R}_t^* \frac{B_{t-1}^* - F_{t-1}^*}{R_{t-1}^*}. \tag{19}$$

25. Note that the arbitrageur's problem omits τ without loss of generality, as a change in income share τ is isomorphic to a re-parameterization of the risk aversion ω, and we take both ω and τ as fixed parameters in our analysis.

The international risk-sharing condition is given by:

$$\beta R_t^* \mathbb{E}_t \frac{C_{Tt}}{C_{T,t+1}} = 1 + \omega\sigma_t^2 \frac{B_t^* - N_t^* - F_t^*}{R_t^*}, \text{ where } \sigma_t^2 = R_t^2 \cdot \mathrm{var}_t\left(\frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}\right). \quad (20)$$

The international risk-sharing wedge is $\hat{\Psi}_t \equiv \omega\sigma_t^2 \frac{B_t^* - N_t^* - F_t^*}{R_t^*}$.

Conditions (19) and (20) are the segmented markets counterparts to the equilibrium conditions (6) and (4) in the baseline model in section 1. The last term in the budget constraint (19) reflects the international transfer of financial-sector income from the home country to the rest of the world. When $\tau = 1$, that is either all income is taxed away or the financial sector is owned by the domestic residents, there is no international transfer and the budget constraint is simply $B_t^*/R_t^* - B_t^* = Y_{Tt} - C_{Tt}$, exactly as before in (6).

The international risk-sharing condition (20) specializes (4) to the case of a segmented market equilibrium, which provides a particular structural interpretation $\hat{\Psi}_t$ to the reduced-form risk-sharing wedge $\hat{\psi}_t$ in (4). When $\hat{\psi}_t = 0$, the international risk-sharing condition reduces to the conventional Euler equation for the foreign-currency bond, $\beta R_t^* \mathbb{E}_t \frac{C_{Tt}}{C_{T,t+1}} = 1$, a property of the constrained optimal risk sharing in this economy. Combining international risk sharing (20) with the home household Euler equation (3), we obtain the modified UIP condition that holds in this economy:

$$\mathbb{E}_t\left\{\frac{\beta C_{Tt}}{C_{T,t+1}}\left[R_t^* - R_t \frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}\right]\right\} = \omega\sigma_t^2 \frac{B_t^* - N_t^* - F_t^*}{R_t^*} = \hat{\Psi}_t. \quad (21)$$

Note that $\hat{\Psi}_t$ is the UIP wedge. When $\hat{\Psi}_t = 0$, whether due to $\omega\sigma_t^2 = 0$ or to $D_t^* = B_t^* - N_t^* - F_t^* = 0$, the UIP holds from the perspective of the home households. Thus, in the limit of risk neutral arbitrageurs $\omega \to 0$, the international financial market converges to a frictionless two-bond market where UIP holds.

To summarize, condition (7) still characterizes the equilibrium allocation $\{C_{Nt}, L_t, Y_{Nt}\}$ in the nontradable sector given sticky prices $P_{Nt} \equiv 1$ and where we think of W_t as directly controlled by monetary policy R_t.[26] Given (7) and the expenditure switching condition (8), the

26. Recall that the choice of domestic policy rate R_t allows to choose the path of nominal wages W_t, as they are linked by the household Euler equation (3), which in light of (5) can be written as $\beta R_t \mathbb{E}_t\{W_t / W_{t+1}\} = 1$; as usual, one needs to ensure the uniqueness of the implemented equilibrium path $\{W_t\}$.

dynamic equilibrium system (19)–(20) characterizes the equilibrium path of $\{C_{Tt}, B_t^*, \mathcal{E}_t\}$ and the implied $\{\sigma_t^2\}$ in the tradable sector. The equilibrium path is shaped by the endowment process $Y_{Tt} = A_{Tt}$, the initial condition B_{-1}^* ,the path of policies $\{R_t, F_t, F_t^*\}$ and exogenous shocks $\{A_{Nt}, A_{Tt}, R_t^*, N_t^*\}$, where recall that $N_t^* = \hat{\psi}_t$ is the noise trader liquidity shock for foreign versus home currency.[27]

3.2 Optimal Policy

We start with the analysis of optimal policies in the case with $\tau = 1$, namely when all income in the financial sector remains in the home country and there is no international transfer associated with noise traders and/or arbitrageurs. The planner's problem in this case delivers the constrained optimum as there is no incentive to manipulate risk sharing or monetary policy to achieve a monetary transfer from the rest of the world. We consider the case with $\tau < 1$ in section 3.2.4.

We use the equilibrium characterization to simplify the policy problem. In particular, we substitute the solution for equilibrium allocation in the nontradable sector (7), namely $C_{Nt} = W_t$ and $L_t = W_t / A_{Nt}$ given fully sticky prices $P_{Nt} = 1$, directly into the household utility function (1). This results in the following welfare objective:

$$\mathbb{W}_0 = max \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left[\gamma \log C_{Tt} + \left(1-\gamma\right)\left(\log W_t - \frac{W_t}{A_{Nt}} \right)\right]. \tag{22}$$

We treat the nominal wage W_t as the instrument of monetary policy, since any path of W_t can be implemented with a suitable interest-rate rule R_t, as we discussed above.

Given W_t and FX interventions F_t^*, tradable consumption must satisfy the country budget constraint (19), the international risk-sharing condition (20), and the expenditure switching condition (8), which we reproduce here as:

27. From $\{B_t^*, F_t^*, N_t^*\}$ we can recover the equilibrium position of intermediaries $D_t^* = B_t^* - F_t^* - N_t^*$ (by market clearing in Lemma 1), and the household home-currency bond position is $B_t / R_t = \mathcal{E}_t (B_t^* - F_t^*)/R_t^* - F_t / R_t$. Note that the home-currency position of the government F_t simply crowds out B_t one-for-one without changing the equilibrium path, a form of Ricardian equivalence in this economy.

$$\frac{B_t^*}{R_t^*} - B_{t-1}^* = Y_{Tt} - C_{Tt} \tag{23}$$

$$\beta R_t^* \mathbb{E}_t \frac{C_{Tt}}{C_{Tt+1}} = 1 + \omega\sigma_t^2 \frac{B_t^* - N_t^* - F_t^*}{R_t^*} \text{ with } \sigma_t^2 = R_t^2 \cdot \text{var}_t\left(\frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}\right), \tag{24}$$

$$\mathcal{E}_t = \frac{\gamma}{1-\gamma}\frac{W_t}{C_{Tt}} \tag{25}$$

where we used $\tau = 1$ in (19) and $C_{Nt} = W_t$ in (8).[28] The unconventional nature of this policy problem is that the equilibrium volatility of the nominal exchange rate σ_t^2 endogenously magnifies the intermediation friction in international risk sharing.

3.2.1 Full Optimal Policies

The planner chooses the path of monetary policy and FX interventions $\{W_t, F_t^*\}$, and the implied equilibrium allocation $\{C_{Tt}, B_t^*, \mathcal{E}_t, \sigma_t^2\}$, to maximize (22) subject to (23)–(25) and given the path of shocks $\{A_{Nt}, A_{Tt}, R_t^*, N_t^*\}$ with $Y_{Tt} = A_{Tt}$.

We note that the policy instrument F_t^* enters only in the international risk-sharing constraint (24), and thus it would be chosen to relax this constraint (that is, ensure a zero Lagrange multiplier). The optimal choice of B_t^* when (24) is not binding requires:

$$\beta R_t^* \mathbb{E}_t \frac{C_{Tt}}{C_{Tt+1}} = 1 \tag{26}$$

that is international risk sharing without a wedge (i.e., $\hat{\Psi}_t = 0$ in Lemma 1). Combining this undistorted risk-sharing condition with the budget constraint (23) determines the unique optimal path of $\{C_{Tt}, B_t^*\}$.

By consequence, this requires setting $F_t^* = B_t^* = N_t^*$ to ensure zero wedge $\hat{\Psi}_t = 0$ independently of the equilibrium volatility of the nominal exchange rate σ_t^2. This characterizes the optimal FX interventions, which lean against the wind—in fact, fully eliminate the wind—by fully accommodating the NFA demand of the households B_t^* and the

28. Another side equation which defines R_t in (24) is the home-currency Euler equation (3), which we write as $\beta R_t \mathbb{E}_t\{W_t / W_{t+1}\} = 1$.

liquidity demand of the noise traders N_t^*. As a result, the arbitrageurs have no job left, and $D_t^* = 0$, the equilibrium risk premium is eliminated, and international intermediation is frictionless. Since imperfect intermediation under segmented markets is the only source of UIP deviations in this economy, the UIP holds under the optimal policy.[29]

Next, consider the optimal monetary policy, namely the choice of $\{W_t\}$. Note that with the undistorted risk sharing, the nominal exchange rate $\mathcal{E}_t$ no longer constrains the optimization over W_t, and the expenditure switching condition (25) acts merely as a side equation. The choice of W_t then becomes static:

$$\tilde{W}_t = \arg\max_{W_t} \{\log W_t - W_t/A_{Nt}\} = A_{Nt}. \tag{27}$$

Setting $W_t = A_{Nt}$ eliminates the state-by-state output gap, that is $X_t = W_t/A_{Nt} = 1$. The equilibrium nominal exchange rate obtains from (25) and equals $\varepsilon_t = \frac{\gamma}{1-\gamma}\frac{A_{Nt}}{C_{Tt}}$.

We summarize this discussion in:

Proposition 1. *The constrained optimum allocation denoted with* $\{\tilde{C}_{Tt}, \tilde{W}_t, \tilde{B}_t^*, \tilde{F}_t^*, \tilde{\varepsilon}_t^*\}$ *maximizes welfare (22) subject to the budget constraint (23) alone, and it is implemented with monetary policy* $\tilde{W}_t = A_{Nt}$, *which closes the state-by-state output gap, and FX interventions* $\tilde{F}_t^* = B_t^* - N_t^*$, *which eliminates the risk-sharing (UIP) wedge in (24). The optimum consumption path* $\{\tilde{C}_{Tt}\}$ *is the unique path that satisfies the dynamic system (23) and (26). The nominal exchange rate is given by* $\tilde{\mathcal{E}}_t = \frac{\gamma}{1-\gamma}\frac{A_{Nt}}{\tilde{C}_{Tt}}$. *The optimal policy is time consistent.*

Intuitively, there are two distortions—output gap due to sticky prices and imperfect risk sharing due to the intermediation friction (under limits to arbitrage)—and two policy instruments (monetary policy and FX interventions), which allow the planner to address both

29. By UIP condition we mean here the household indifference condition between the home- and foreign-currency bonds, that is $\mathbb{E}_t\left\{\frac{\beta C_{Tt}}{C_{T,t+1}}\left[R_t^* - R_t\frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}\right]\right\} = 0$, which features a representative household's UIP risk premium. More generally, the planner wants to illuminate the intermediation wedge, leaving intact the fundamental sources of the risk premium.

distortions and deliver the constrained optimum.[30] The property of the constrained optimum is zero wedges in production (output gap) and in international risk sharing, $X_t = 1$ and $\hat{\Psi}_t = 0$. The maximum utility is given by $\tilde{\mathbb{W}}_0 = \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left[\gamma \log \tilde{C}_{Tt} + (1-\gamma)(\log A_t - 1) \right]$, and we use it as the benchmark for the remaining analysis:

$$\mathbb{W}_0 - \tilde{\mathbb{W}}_0 = \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left[\gamma \log \frac{C_{Tt}}{\tilde{C}_{Tt}} + (1-\gamma)\left(\log \frac{W_t}{A_t} - \frac{W_t - A_t}{A_t} \right) \right] \leq 0.$$

where the first term is the loss from risk-sharing distortions and the second term is the loss from the output gap.

Importantly, the optimal policy is time consistent, as both instruments remove the respective distortions contemporaneously and require no intertemporal promises. As a result, the implementation of the constrained optimum allocation does not require commitment on the part of the monetary authority.

There is no closed form characterization of $\tilde{C}_{Tt}$ in the presence of uninsured country risk in Y_{Tt}, but when $Y_{Tt} = A_{Tt}$ follows a random walk, $\tilde{C}_{Tt}$ follows a near-random walk with changes in $\tilde{C}_{Tt}$ approximately equal to changes in A_{Tt}. What are the implications of this for the nominal and real exchange rate? The nominal exchange rate $\tilde{\mathcal{E}}_t = \frac{\gamma}{1-\gamma} \frac{A_{Nt}}{\tilde{C}_{Tt}}$, as well as the real exchange rate $\tilde{Q}_t = \tilde{\mathcal{E}}_t^{1-\gamma}$, appreciates with the relative productivity in the tradable sector, that is, when tradable endowment A_{Tt} increases sharper than nontradable productivity A_{Nt}. Indeed, this is the Balassa-Samuelson force, which shapes the path of the real exchange rate in proportion with the relative tradable-nontradable productivity. Under sticky prices, implementing this path for the real exchange rate requires the nominal exchange rate to follow the same relative productivities.

Implementing the constrained optimum in an economy with sticky prices and frictional financial market requires an active use of both monetary policy and FX interventions but does not require the use

30. Note that the constrained optimum is not first best as international financial market is incomplete and only allows to share risk in expectation given the foreign interest rate R_t^*. This is equivalent to a single foreign-currency bond economy. Interestingly, the presence of the home-currency bond is irrelevant for the optimal allocation, as R_t is merely a side variable and does not affect the equilibrium allocation in this case, and the planner has no incentive to use any additional instrument (e.g., capital controls; see below).

of capital controls. The goal of FX interventions is not to eliminate exchange rate volatility, but rather to eliminate the risk-sharing wedge—the UIP deviation $\hat{\Psi}_t$ due to the intermediation friction. No UIP deviations are, in fact, consistent with a volatile nominal exchange rate, which itself is generally a consequence of the optimal monetary policy stabilizing output gap.[31] In segmented financial markets, FX interventions provide the government with an important additional tool, which allows to fix distortions associated with frictional intermediation. The use of FX interventions does not interfere with monetary policy, which is focused on domestic output-gap stabilization, as in the closed economy, and does not generally require the use of capital controls. In this sense, such economy does not feature the trilemma trade-off present in conventional monetary models with a frictionless financial market.[32]

3.2.2 Divine Coincidence: Fixed Exchange Rate

In the constrained optimum allocation, FX interventions $\tilde{F}_{Tt} = B_t^* - N_t^*$ eliminate the risk-sharing wedge ($\hat{\Psi}_t = 0$), but do not result in a stable exchange rate ($\mathcal{E}_t \neq const$ in general). Indeed, the nominal exchange rate traces the frictionless real exchange rate, which in turn reflects the relative movements in nontradable productivity (relative to tradable endowment). We now explore the special case when a fixed exchange rate implements the constrained optimum.

Note also that the constrained optimum implementation requires the use of both instruments—monetary policy W_t and FX interventions F_t^*—and, in general, it cannot be implemented with monetary policy alone. There exists, however, an important special, yet robust, case when monetary policy alone can simultaneously implement both goals—output-gap stabilization and elimination of the international risk-sharing wedge—without any need to use FX interventions. This case relies on the full stabilization of the nominal exchange rate—the fixed exchange rate—which can be achieved by means of monetary

31. As shown above, the nominal exchange rate implementing the first best follows the relative nontradable productivity. Arguably, the volatility of relative productivities is not as large as the observed volatility of floating exchange rates, e.g., dollar/euro (10% annualized standard deviation). Thus, it is likely that optimal FX interventions partially stabilize the exchange rate relative to laissez-faire, as we further discuss below.

32. Note that this does not mean however that any path of the exchange rate can be implemented without compromising the ability of monetary policy to stabilize inflation and output gap, and in this sense the trilemma is still present.

policy and thus eliminates the need to use FX interventions. We refer to this special case as the *divine coincidence* in an open economy.

Indeed, examining the general policy problem (22), the limiting case with a commitment to fixed exchange rate $\mathcal{E}_t = const$ implies $\sigma_t^2 = 0$, and thus eliminates the risk-sharing wedge (ensures $\hat{\Psi}_t = 0$), irrespective of the use of the other instrument F_t^*. Furthermore, since $\mathcal{E}_t = \frac{\gamma}{1-\gamma}\frac{W_t}{C_{Tt}}$, monetary policy can always ensure a fixed exchange rate by setting $W_t / C_{Tt} = const$.

The only remaining question is when such monetary policy can also be optimal from the point of view of the output-gap stabilization, that is, ensure that $X_t = W_t / A_{Nt} = 1$. While being a knife-edge case, it is an important one and can be formulated as follows: if the first-best real exchange rate—i.e., the real exchange rate corresponding to the first-best allocation with zero output gap—is constant, then fixed nominal exchange rate is the optimal policy stabilizing simultaneously output gap and international risk sharing. Indeed, recall that the real and nominal exchange rates perfectly comove under sticky prices, $\mathcal{Q}_t = \mathcal{E}_t^{1-\gamma}$, so that if the first-best real exchange rate $\tilde{Q}_t = (\frac{\gamma}{1-\gamma}\frac{A_{Nt}}{\tilde{C}_{Tt}})^{1-\gamma} = const$, then it can always be implemented with $\mathcal{E}_t = const$ independently of the degree of price stickiness. Furthermore, this is an "if and only if" statement, and the fixed exchange rate is necessarily suboptimal whenever $\tilde{Q}_t \neq const$ and prices are (at least partially) sticky.

Proposition 2. *The fixed nominal exchange rate implements the constrained optimum allocation if and only if the first-best real exchange rate is stable,* $\tilde{Q}_t = const$. *In this case, monetary policy alone can achieve both goals of output-gap stabilization,* $X_t = 1$, *and elimination of the international risk-sharing wedge,* $\hat{\Psi}_t = 0$, *without the use of FX interventions or capital controls.*

When can we expect the first-best real exchange rate to be stable? In our setup, this is the case when Balassa-Samuelson forces exactly offset each other and, in particular, the nontradable productivity and tradable endowment comove in lockstep. Formally, this would require a near-random walk perfectly correlated processes in both $Y_{Tt} = A_{Tt}$ and A_{Nt}, so that C_{Tt} tracks Y_{Tt} and thus $A_{Nt} / C_{Tt} = const$.[33] More generally, the real exchange rate may also vary because of the differential

33. In a linearized environment, this is exactly the case, as $c_{Tt} = y_{Tt}$ under a random walk endowment, but in a full nonlinear problem, the path of C_{Tt} differs from that of Y_{Tt} due to precautionary savings from uninsured idiosyncratic risk.

evolution of home and foreign tradable productivity under home bias in tradable consumption. The divine coincidence principle generalizes to those environments and still suggests that if one can argue that the first-best real exchange rate is stable, then a fixed nominal exchange rate regime implements the constrained optimum and achieves both policy objectives without the need to use other instruments such as exchange rate interventions or capital controls. In other words, divine coincidence is exactly the case where inflation (output-gap) stabilization does not come into conflict with a fixed exchange rate and thus the trilemma, if present, is not binding.

Implementation. We focused above on the direct implementation of the peg using W_t. Two remarks are in order. First, the same allocation can be implemented using an interest rate R_t rule, as pointed out above. Second, and more importantly, either W_t or R_t implementation can either target output gap or nominal exchange rate directly. Indeed, divine coincidence implies that fixed exchange rate equilibrium corresponds to the zero output-gap equilibrium. However, the implementation of the policy does matter, as targeting output gap may be consistent with multiple exchange rate equilibria, one with $\sigma_t^2 = 0$ and another with $\sigma_t^2 > 0$, and only the former one ensures undistorted international risk sharing.[34] Therefore, in terms of implementation, a monetary policy that explicitly targets the nominal exchange rate can be superior to that stabilizing the output gap, even under divine coincidence. In this sense, the model captures the idea of using a nominal peg to anchor expectations, although the focus is on the financial-market expectations rather than inflation expectations of households and firms.[35]

3.2.3 Single Instrument without Divine Coincidence

Proposition 1 characterized the optimal joint use of monetary policy and FX interventions, which allows to implement the optimal

34. Formally, compare the case with $W_t = A_{Nt}$ and $W_t = \kappa C_{Tt}$ for some appropriately chosen $\kappa > 0$, which under divine coincidence are both consistent with the optimal allocation. While the latter implementation ensures $\mathcal{E}_t = const$ from (25) and thus $\sigma_t^2 = 0$, the former may be consistent with multiple equilibria that solve $\beta R_t^* \mathbb{E}_t \frac{C_{Tt}}{C_{Tt+1}} = 1 + \omega\sigma_t^2 \frac{B_t^* - N_t^*}{R_t^*}$ where $\sigma_t^2 = R_t^2 \cdot \mathrm{var}_t\left(\frac{C_{Tt+1}/C_{Tt}}{A_{Nt+1}/A_{Nt}}\right)$, in addition to the budget constraint (23). The multiplicity of solutions for (C_{Tt}, σ_t^2) translates into the multiplicity of solutions for $\mathcal{E}_t$, with $\sigma_t^2 = 0$ solution welfare dominating other possible solutions.

35. *Cf*. Marcet and Nicolini (2003).

allocation by eliminating both the output gap and the international risk-sharing wedge state by state. Proposition 2 shows how monetary policy can fully stabilize the nominal exchange rate, which immediately eliminates the risk-sharing wedge without the use of FX interventions, and further characterizes circumstances when it is also optimal from the point of output-gap stabilization. As a corollary, when prices are flexible and thus the output gap is absent irrespective of monetary policy, the optimal risk sharing can be always achieved by monetary policy that stabilizes the nominal exchange rate, without the use of FX interventions. In other words, equilibrium nominal exchange rate volatility can be desirable only under sticky prices, when it needs to accommodate the real exchange rate variation that cannot be achieved via adjustment of prices.

We now consider the reverse case of whether the output gap can be stabilized by FX interventions alone, when monetary policy is constrained, e.g., by the zero lower bound $R_t \geq \underline{R}$ or fixed exchange rate $\mathcal{E}_t = \bar{\mathcal{E}}$.[36] In contrast to the previous case, it is *not* possible to implement the first-best allocation with FX interventions. In particular, fixed exchange rate implies $\sigma_t^2 = 0$ in (20) and, while it immediately eliminates the risk-sharing wedge, it also makes FX interventions F_t^* irrelevant for the equilibrium allocation. F_t^* can still affect allocation $\{C_{Tt}, C_{Nt}\}$ under the zero-lower-bound constraint if $\sigma_t^2 > 0$. However, under separable utility, F_t^* is optimally used to only eliminate the risk-sharing wedge in tradables without targeting the allocation of nontradables and the output gap.[37]

This analysis in particular suggests that FX interventions cannot substitute for monetary policy. We next explore the optimal use of monetary policy in the presence of both frictions when FX

36. Recall that, under sticky prices, $P_{Nt} = 1$, we have $C_{Nt} = W_t$, and the loss from the output gap can be written as $\log\frac{C_{Nt}}{A_t} - \frac{C_{Nt}-A_t}{A_t} \leq 0$. Furthermore, C_{Nt} must satisfy $\beta R_t \mathbb{E}_t\left\{C_{Nt} / C_{N,t+1}\right\} = 1$ and $\mathcal{E}_t = \frac{\gamma}{1-\gamma}\frac{C_{Nt}}{C_{Tt}}$, with the former possibly constrained by the zero lower bound and the latter by the fixed exchange rate.

37. With nonseparable utility in (C_{Tt}, C_{Nt}), FX interventions can depart from the optimal risk sharing $\beta R_t^* \mathbb{E}_t\left\{u_{Tt}/u_{Tt+1}\right\} = 1$ in order to relax the constraint imposed by $\beta R_t \mathbb{E}_t\left\{u_{Nt} / u_{Nt+1}\right\} = 1$ when R_t cannot adjust (where u_{Tt} and u_{Nt} correspond to marginal utility of tradable and nontradable consumption, respectively). As in the general theory of second best, the constrained optimal policy introduces a wedge into international risk sharing if it allows to reduce the domestic output gap. Unlike capital controls or other taxes, however, which can directly distort $\beta R_t \mathbb{E}_t\left\{\frac{u_{Tt}}{u_{Tt+1}}\frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}\right\} = 1$, FX interventions are less capable and operate exclusively via their indirect effect on C_{Tt} in (20). Cf. Farhi and Werning (2012), Correia and others (2013), Farhi and others (2014).

interventions F_t^* are not available. In this case, the optimal monetary policy closes the output gap on average and trades off the state-by-state variation in output gap ex post with a reduction in the risk-sharing wedge ex ante by partially stabilizing the future nominal exchange rate. Formally, the optimal monetary policy ensures $\mathbb{E}_t X_{t+1} = 1$, where $X_{t+1} = W_{t+1}/A_{Nt+1}$ is the output gap, but varies $X_{t+1} \neq 1$ state by state to reduce σ_t^2, in particular in periods following large risk-sharing wedges $\hat{\Psi}_t = \omega\sigma_t^2 \frac{N_t^* - B_t^*}{R_t^*} \neq 0$.[38] The policy reduces $C_{Nt+1} = W_{t+1}$ below A_{Nt+1} when C_{Tt+1} is low, and vice versa, which reduces the volatility of $\mathcal{E}_{t+1} = \frac{\gamma}{1-\gamma}\frac{W_{t+1}}{C_{Tt+1}}$ by making tradable and nontradable consumption more correlated. This is the optimal trade-off between the two frictions, namely giving up on fully stabilizing the output gap at $t + 1$ to reduce the international risk-sharing wedge at t to smooth tradable consumption.

We summarize these results in the following proposition and provide a formal proof in Appendix B:

Proposition 3. *(i) Monetary policy can eliminate the risk-sharing wedge, while FX interventions cannot close the output gap when monetary policy is constrained and can only ensure constrained optimal international risk sharing. (ii) Optimal monetary policy in the absence of FX interventions eliminates the output gap on average and uses the state-by-state variation in output gap to partially reduce the volatility of the nominal exchange rate and the ex-ante risk-sharing (UIP) wedge.*

This proposition emphasizes that FX interventions are a direct instrument to offset international risk-sharing wedges emerging as a result of imperfect intermediation. This result generalizes beyond segmented market models and applies in noncompetitive environments with rents and markups and in models with financial constraints.[39] As the same time, FX interventions are ineffective to address other frictions such as output gap or, in richer models, inefficiencies arising from overborrowing due to pecuniary externalities.[40]

The proposition also suggests that pure floats are generally suboptimal when monetary policy focuses exclusively on output-gap and inflation stabilization and FX interventions are not used.

38. In contrast, $\mathbb{E}_t X_{t+1} = 1$ state by state in periods following $\hat{\Psi}_t = 0$, i.e., when risk-sharing UIP deviations are small due to a combination of small risk aversion ω, small exchange rate volatility σ_t^2, and/or small equilibrium financial flows $N_t^* - B_t^*$.

39. For example, Gabaix and Maggiori (2015), Adrian and others (2015), Jiang and others (2021), Bianchi and others (2021).

40. For example, Basu and others (2020).

Instead, partial and crawling pegs whereby either FX interventions or monetary policy are used to partially stabilize or eliminate short-run exchange rate volatility are generally superior to pure floats, as well as to outright pegs. Full pegs are optimal under divine coincidence and pure floats are optimal when wedges arising from intermediation frictions are negligible. The latter happens when either risk-bearing capacity is large (small ω) or financial flows $N_t^* - B_t^*$ are small relative to the absorption capacity of the financial market (a deep financial market). A sign of a deep financial market are small UIP deviations despite large ex-post exchange rate volatility. In contrast, when UIP deviations are large, this may indicate frictional intermediation and call for policy intervention to smooth out UIP deviations. In other words, large ex-ante UIP deviations is a necessary condition for a welfare improving exchange rate intervention.[41]

Discretionary policy. An important property of the optimal policies in Proposition 1 was time consistency and no need for commitment to implement them. As described above, the optimal monetary policy in the absence of FX interventions trades off output-gap stabilization at $t + 1$ for a reduction in the risk-sharing wedge at t. This requires commitment on the part of the monetary authority, as the only time-consistent discretionary outcome is the state-by-state output-gap stabilization, $X_{t+1} = 1$, which leaves a laissez-faire international risk-sharing wedge $\hat{\Psi}_t$. This is suboptimal, as shown in Proposition 3.

3.2.4 International Transfers. Capital Controls

We now consider the case with international transfers when $\tau < 1$ in the country budget constraint (19), which we rewrite as:

$$\frac{B_t^*}{R_t^*} - B_{t-1}^* = \left(Y_{Tt} - C_{Tt}\right) - \tilde{\tau}\tilde{R}_t \frac{B_{t-1}^* - F_{t-1}^*}{R_{t-1}^*}, \tag{28}$$

where we denoted $\tilde{\tau} \equiv 1 - \tau > 0$ and carry trade return $\tilde{R}_t \equiv R_{t-1}^* - R_{t-1}\frac{\mathcal{E}_{t-1}}{\mathcal{E}_t}$. Thus, the planner maximizes the objective (22) subject to (28), (24)–(25) and the Euler equation (3), which determines R_t. For convenience, we

41. If UIP deviations reflect default or counterparty risk rather than intermediation friction or rents, then FX interventions are not justified as a policy response. See Amador and others (2019).

combine (3) with (24) to write the constraint as the UIP condition (21) on R^*_{t+1}, which we reproduce here as follows:

$$\mathbb{E}_t\left\{\Theta_{t+1}\tilde{R}^*_t\right\} = \omega\sigma_t^2\,\frac{B^*_t - N^*_t - F^*_t}{R^*_t} = \hat{\Psi}_t, \tag{29}$$

where recall that $\Theta_{t+1} = \beta\frac{C_{Tt}}{C_{Tt+1}}$ is the home household's stochastic discount factor (SDF) for returns in foreign currency and $\hat{\Psi}_t$ is the UIP wedge.

The last term in the budget constraint (28) corresponds to the international wealth transfer, which obtains when the noise traders and arbitrageurs jointly make losses on their financial positions, as their losses are the gains of the combined home households and government sector. Under these circumstances, while it is still feasible, it is no longer optimal for the government to fully eliminate the risk-sharing wedge $\hat{\Psi}_t$ in (29). First, consider the optimal policies from Proposition 1, namely $W_t = A_t$ and $F^*_t = B^*_t - N^*_t$, which still eliminate both the output gap and the risk-sharing wedge. In this case, the country budget constraint becomes:

$$\frac{B^*_t}{R^*_t} - B^*_t = \left(Y_{Tt} - C_{Tt}\right) - \tilde{\tau}\tilde{R}^*_t\psi_{t-1}.$$

Where $\psi_{t-1} = \frac{N^*_{t-1}}{R^*_{t-1}}$ is the exogenous noise trader liquidity demand for dollar relative to home currency. As a result, this allocation is associated with mean-zero idiosyncratic international transfers (evaluated using the home household SDF):

$$\mathbb{E}_{t-1}\left\{\Theta_t\tilde{\tau}\tilde{R}^*_t\psi_{t-1}\right\} = \tilde{\tau}\psi_{t-1}\mathbb{E}_{t-1}\left\{\Theta_t\tilde{R}^*_t\right\} = 0$$

and they contribute to the national income volatility of the home country thus reducing welfare. Can the government improve upon this allocation? In particular, is it feasible to eliminate income risk or even create systematic transfers from the rest of the world.

One can show that departures from $W_t = A_t$, if UIP still holds in expectation, generate at most third-order benefits, while creating second-order losses from departures from output gap. Thus, we focus here for concreteness on monetary policy that stabilizes output gap, $W_t = A_t$, and explore the use of FX interventions F^*_t in the presence of international transfers. We rewrite the budget constraint (28) as:

$$\frac{B_t^*}{R_t^*} - B_{t-1}^* = \left(Y_{Tt} - C_{Tt}\right) - \tilde{\tau}\tilde{R}_t^* \left[\psi_{t-1} + \frac{\mathbb{E}_{t-1}\Theta_t \tilde{R}_t^*}{\omega\sigma_{t-1}^2} \right],$$

and the government has a direct control over the size of the UIP deviation, $\mathbb{E}_t\Theta_{t+1}\tilde{R}_{t+1}^* = \hat{\Psi}_t$ by means of FX interventions F_t^* in (29). Therefore, the tradeoff faced by the policymaker is whether to engineer ex-ante UIP deviations, which distort risk sharing, yet can generate additional national income under certain circumstances.

The expected discounted income (using home SDF) from FX interventions that allow for UIP deviations ($\hat{\Psi}_t \neq 0$) is given by:

$$-\tilde{\tau}\mathbb{E}_t\Theta_{t+1}\tilde{R}_{t+1}^* \left[\psi_t + \frac{\mathbb{E}_t\Theta_{t+1}\tilde{R}_{t+1}^*}{\omega\sigma_t^2} \right] = -\tilde{\tau} \left[\psi_t\hat{\Psi}_t + \frac{\hat{\Psi}_t^2}{\omega\sigma_t^2} \right].$$

Therefore, the expected income is (weakly) negative in the absence of noise trader demand (when $\psi = 0$), and thus $\hat{\Psi}_t = 0$ is optimal in this case as it guarantees both efficient risk sharing and no expected income losses. A corollary of this result is that, if noise traders are domestic and arbitrageurs are foreign, the government can generate no expected income and should ensure $\hat{\Psi}_t = 0$ by setting $F_t^* = B_t^* - N_t^*$ as in Proposition 1.[42]

In the presence of international noise trader demand, the policymaker can generate expected incomes by partially "leaning against the wind" of their currency demand and choosing F_t^* such that:

$$\psi_t\hat{\Psi}_t \propto N_t^* \cdot \left(B_t^* - N_t^* - F_t^*\right) < 0.$$

The income gains of the government are limited, however, by the arbitrageurs, who take positions in the same direction as the government and inversely proportionally to $\omega\sigma_t^2$. As a result, in the limit of $\omega\sigma_t^2 \to 0$, the government cannot sustain any expected income gains, even in the presence of noise traders, and should not attempt to choose $\hat{\Psi}_t \neq 0$, which would be futile anyways. Finally, for any $\omega\sigma_t^2 > 0$, UIP deviations $\hat{\Psi}_t$ in response to $\psi_t \neq 0$ generate income gains that are first order in $\hat{\Psi}_t$ and welfare losses from the resulting risk-sharing wedge that are second order in $\hat{\Psi}_t$, around $\hat{\Psi}_t = 0$. Therefore, nonzero

42. *Cf.* Amador and others (2019), Fanelli and Straub (2021).

UIP deviation $\hat{\Psi}_t$ are necessarily desirable in this case, if sufficiently small.[43] We summarize this discussion in the following proposition:

Proposition 4. (i) *Expected income from FX interventions is weakly negative in the absence of foreign noise trader demand, and thus* $F_t^* = B_t^* - N_t^*$ *to ensure* $\hat{\Psi}_t = 0$ *is optimal in this case.* (ii) *In the presence of foreign noise trade demand, there exist FX interventions* F_t^* *that partially lean against* N_t^* *and generate expected incomes that exceed welfare losses from the induced risk-sharing (UIP) wedge* $\hat{\Psi}_t \neq 0$.

Volatility of the central bank's balance sheet. The policy of FX interventions, whether it results in UIP deviations or not, leads to ex-post income and losses borne by the central bank, even when expected incomes and losses might be zero. In particular, the ex-post income of the central bank is given by $\tilde{R}^*_{t+1}\frac{F_t^*}{R_t^*} = \left[1 - \frac{R_t}{R_t^*}\frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}\right]F_t^*$ and its variance is given by $\sigma_t^2 \cdot (F_t^*/R_t^*)^2$. Thus, two possible constraints on the central bank's balance sheet may be non-negative foreign reserves $F_t^* \geq 0$ or a value at risk constraint $|F_t^*| \leq \alpha R_t^*/\sigma_t$. Both constraints may limit the ability of the central bank to implement the optimal policies and, in particular, the policy $F_t^* = B_t^* - N_t^*$ from Proposition 1 may be infeasible.

Furthermore, the region of feasibility may not be connected, as there is feedback between policy F_t^* and equilibrium exchange rate volatility σ_t^2. More specifically, limited interventions F_t^* may result in large equilibrium exchange rate volatility σ_t^2, while large interventions, vice versa, limit significantly the equilibrium σ_t^2, thus possibly making the intermediate levels of interventions infeasible.

Finally, in cases when sufficiently large interventions are infeasible and the lowest achievable σ_t^2 with FX interventions is large, a fully fixed exchange rate by means of monetary policy may be superior relative to the output-gap-stabilizing monetary policy and the best feasible level of FX interventions. This can be the case, in particular, even when the divine coincidence of Propositions 2 is not satisfied. Thus, this offers a justification for some exchange rate pegs that are adopted despite the resulting output gaps and suboptimal real exchange rate under the peg.

Capital controls. So far, we have left out capital controls from our considerations. Indeed, Propositions 1 and 2 show that optimal

43. The maximum expected income equals $\frac{1}{4}\tilde{\tau}\omega\sigma_t^2\psi_t^2$, and it is achieved when $\hat{\Psi}_t = -\frac{1}{2}\,\omega\sigma_t^2\psi_t$, or equivalently $F_t^* = B_t^* - \frac{1}{2}N_t^* = B_t^* - \frac{1}{2}R_t^*\psi_t$. The optimal intervention additionally takes into account the welfare loss from the risk-sharing wedge which is increasing in $\hat{\Psi}_t^2$.

allocations can be attained without any use of capital controls, as long as there are no international transfers ($\tau = 1$ in (19)) and both monetary policy and FX interventions are available and unconstrained. As soon as we consider the full policy problem, which features a general budget constraint (28) with a possibility of transfers, capital controls become useful. The only constraint that cannot be relaxed is the budget constraint; (24) and (29) can be relaxed provided that there are enough policy instruments. Indeed, FX interventions relax the risk-sharing constraint (24), while capital controls on households (or other intertemporal taxes) relax the UIP condition (29). This effectively makes R_t a free choice variable allowing the government to manipulate UIP deviations with both F_t^* and capital controls, thus further maximizing the rents that can be extracted from noise traders.[44] In general, these rents are limited by the intermediation of arbitrageurs, unless separate capital controls can be levied on the arbitrageurs as well.

4. Conclusion

This paper outlines a simple model of exchange rate determination, which is broadly consistent with the major exchange rate puzzles and uses it to study the optimal exchange rate policy. We emphasize the transmission of monetary and financial shocks via goods and financial markets, which is crucial to explain the PPP, UIP and Mussa puzzles. Sticky prices and financial intermediation frictions imply that there are two wedges in the economy—the output gap and deviations from the optimal risk sharing—and closing them with one policy instrument is only feasible when the optimal real exchange rate is stable. This open economy divine coincidence calls for a fixed nominal exchange rate. More generally, two instruments are required to implement the optimal allocation: while interest-rate policy targets the output gap, FX interventions are used to eliminate UIP deviations, eliminating financial noise but allowing for fundamental exchange rate volatility. When only the monetary instrument is available, the second-best policy balances the two objectives and partially stabilizes the nominal exchange rate, resulting in a partial crawling peg.

While we focus on exchange rate policies, the normative implications are not limited to an open economy environment. It is

44. For further analysis see Itskhoki and Mukhin (2022).

intriguing to study, both theoretically and empirically, the transmission mechanism of monetary policy via financial markets in a closed economy. The ability of a peg to stabilize the risk premium on the carry trade raises the question of whether monetary policy can and should partially stabilize the volatility in the *equity* risk premium by targeting a stock market index. How such policy affects the economy and whether it is desirable are important questions for future research.

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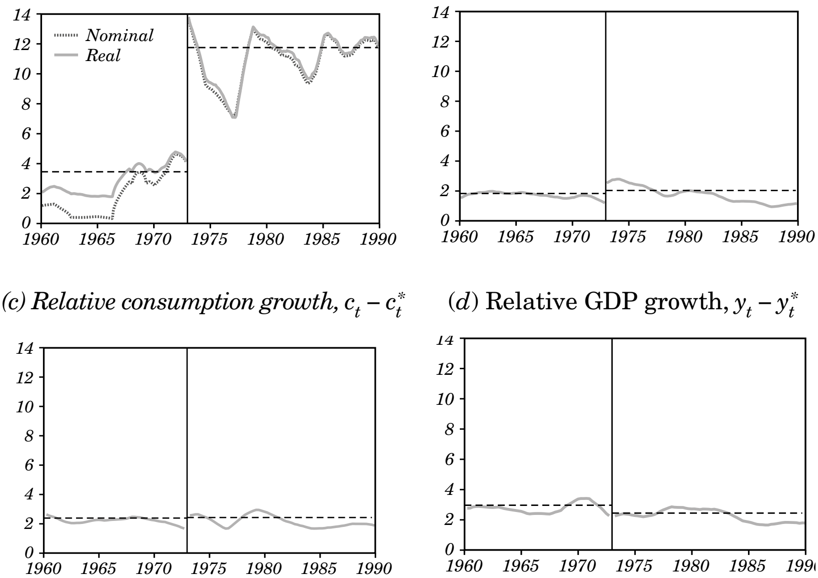
Appendices

A. Additional Figures

Figure A1. Macroeconomic Volatility Over Time

(*a*) *Exchange rates changes, e_t and q_t* (*b*) *Relative inflation, $\pi_t - \pi_t^*$*

(*c*) *Relative consumption growth, $c_t - c_t^*$* (*d*) Relative GDP growth, $y_t - y_t^*$



Source: Itskhoki and Mukhin (2021b).

Note: Annualized standard deviations (in log points) for G7 countries (plus Spain, without Canada) relative to the U.S., estimated as moving averages with a window over 18 months (for exchange rates and inflation) or ten quarters (for consumption and real GDP growth) before and after, treating 1973.1 as the end point for the two regimes; the dashed lines correspond to the average standard deviations under the two regimes. Note that under a full peg ($e_t = 0$), by definition $q_t = \pi_t - \pi_t^*$; under a float, the empirical correlation between e_t and q_t is close to 1 and between q_t (or e_t) and $\pi_t - \pi_t^*$ is close to 0. See figure 1 for raw data series for q_t and c_t.

B. Derivations and Proofs

Proof of Lemma 1. First, we use market clearing for home-currency bond, $B_t + N_t + D_t + F_t = 0$, and the zero capital (carry trade) portfolios of noise traders and arbitrageurs, $D_t/R_t + \mathcal{E}_t D_t^*/R_t^* = 0$, and $N_t/R_t + \mathcal{E}_t N_t^*/R_t^* = 0$, to obtain:

$$\frac{B_t + F_t}{R_t} - \frac{\mathcal{E}_t\left(N_t^* + D_t^*\right)}{R_t^*} = 0.$$

Then using the definition of the country's NFA position,

$$\frac{\mathcal{E}_t B_t^*}{R_t^*} = \frac{B_t + F_t}{R_t} + \frac{\mathcal{E}_t F_t^*}{R_t^*},$$

to express out $B_t + F_t$ and dividing through by $\mathcal{E}_t/R_t^*$ results in $B_t^* = F_t^* + N_t^* + D_t^*$, as stated in the lemma.

Second, substitute firm profits $\Pi_t = P_{Tt}Y_{Tt} + P_{Nt}Y_{Nt} - W_t L_t$ and household consumption expenditure $P_t C_t = P_{Nt}C_{Nt} + P_{Tt}C_{Tt}$ into the household budget constraint (2) and use market clearing $C_{Nt} = Y_{Nt}$ to obtain:

$$\frac{B_t}{R_t} - B_{t-1} = NX_t + T_t,$$

where $NX_t = P_{Tt}Y_{Tt} + P_{Tt}C_{Tt} = \mathcal{E}_t(Y_{Tt} - C_{Tt})$ using the law of one price with $P_{Tt}^* = 1$. Next combine this with the government budget constraint (in the text) to obtain:

$$\frac{B_t + F_t}{R_t} + \frac{\mathcal{E}_t F_t^*}{R_t^*} - B_{t-1} - F_{t-1} - \mathcal{E}_t F_{t-1}^* = NX_t + \tau\mathcal{E}_t\pi_t^*.$$

Using the definition of NFA B_t^* above and the market clearing $B_t + D_t + N_t + F_t = 0$, as well as the result above that $B_t^* = D_t^* + N_t^* + F_t^*$, we rewrite:

$$\frac{\mathcal{E}_t B_t^*}{R_t^*} - \mathcal{E}_t B_{t-1}^* + \mathcal{E}_t\left(D_{t-1}^* + N_{t-1}^*\right) + \left(D_{t-1} + N_{t-1}\right) = NX_t + \tau\mathcal{E}_t\pi_t^*.$$

Finally, recall that $\pi_t^* = \tilde{R}_t^* \frac{D_{t-1}^* + N_{t-1}^*}{R_{t-1}^*} = \left[1 - \frac{R_{t-1}}{R_{t-1}^*}\frac{\mathcal{E}_{t-1}}{\mathcal{E}_t}\right]\left(D_{t-1}^* + N_{t-1}^*\right)$. Subtract $\mathcal{E}_t\pi_t^*$ on both sides:

$$\frac{\mathcal{E}_t B_t^*}{R_t^*} - \mathcal{E}_t B_{t-1}^* = NX_t - \left(1-\tau\right)\tilde{R}_t^* \frac{\mathcal{E}_t\left(D_{t-1}^* + N_{t-1}^*\right)}{R_{t-1}^*},$$

where we used the fact that zero-capital portfolios of noise traders and arbitrageurs imply: $\left(D_{t-1} + N_{t-1}\right) + \frac{R_{t-1}}{R_{t-1}^*}\mathcal{E}_{t-1}\left(D_{t-1}^* + N_{t-1}^*\right) = 0$.

Divide through by ε_t, use the fact that $NX_t/\mathcal{E}_t = Y_{Tt} - C_{Tt}$, and the fact above that $D_{t-1}^* + N_{t-1}^* = B_{t-1}^* - F_{t-1}^*$ to rewrite:

$$\frac{B_t^*}{R_t^*} - B_{t-1}^* = \left(Y_{Tt} - C_{Tt}\right) - \left(1-\tau\right)\tilde{R}_t^* \frac{\mathcal{E}_t\left(B_{t-1}^* - F_{t-1}^*\right)}{R_{t-1}^*},$$

resulting in (19) in the lemma.

Finally, (20) in the lemma follows directly from the optimal portfolio of the arbitrageurs (in the text), which we rewrite expanding the expressions for Θ_{t+1} and $\tilde{R}_{t+1}^*$ as:

$$\omega\sigma_t^2 \frac{D_t^*}{R_t^*} = \mathbb{E}_t\left\{\Theta_{t+1}\tilde{R}_{t+1}^*\right\} = \mathbb{E}_t\left\{\beta\frac{C_{Tt}}{C_{T,t+1}}\cdot\left[R_t^* - R_t\frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}\right]\right\}.$$

Subtracting the household Euler equation (3), after noting that optimal household expenditure $\gamma P_{Nt}C_{Nt} = (1-\gamma)\,\mathcal{E}_t C_{Tt}$ and substituting for $D_t^* = B_{t-1}^* - N_{t-1}^* - F_{t-1}^*$ finishes the proof.

Proof of Proposition 3. Consider the policy problem when F_t^* is constrained, for concreteness $F_t^* = 0$:

$$\max_{\left\{C_{Tt}, C_{Nt}, B_t^*, \mathcal{E}_t, R_t, \sigma_t^2\right\}} \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left[\gamma \log C_{Tt} + \left(1-\gamma\right)\left(\log C_{Nt} - \frac{C_{Nt}}{A_t}\right)\right],$$

subject to

$$\frac{B_t^*}{R_t^*} - B_{t-1}^* = Y_{Tt} - C_{Tt},$$

$$\beta R_t^* \mathbb{E}_t \frac{C_{Tt}}{C_{T,t+1}} = 1 + \omega\sigma_t^2 \frac{B_t^* - N_t^*}{R_t^*},$$

$$\beta R_t^* \mathbb{E}_t \frac{C_{Nt}}{C_{N,t+1}} = 1,$$

$$\mathcal{E}_t = \frac{\gamma}{1-\gamma}\frac{C_{Nt}}{C_{Tt}},$$

$$\sigma_t^2 = R_t^2 . \text{var}_t\left(\frac{\mathcal{E}_t}{\mathcal{E}_{t+1}}\right).$$

Denote $\Gamma_t \equiv 1/C_{Nt}$ and express out R_t and E_t using the third and fourth constraints:

$$\max_{\{C_{Tt},\Gamma_t,B_t^*,\sigma_t^2\}} \mathbb{E}_0 \sum_{t-0}^{\infty} \beta^t \left[\gamma \log C_{Tt} - (\gamma - 1)\left(\log \Gamma_t + \frac{1}{A_t \Gamma_t}\right)\right]$$

subject to $\dfrac{B_t^*}{R_t^*} - B_{t-1}^* = Y_{Tt} - C_{Tt},$

$$\beta R_t^* \mathbb{E}_t \frac{C_{Tt}}{C_{Tt+1}} = 1 - \omega\sigma_t^2 \frac{N_t^* - B_t^*}{R_t^*},$$

$$\beta^2 C_{Tt}^2 \left(\mathbb{E}_t \Gamma_{t+1}\right)^2 \sigma_t^2 = \mathbb{E}_t \left(\Gamma_{t+1} C_{T,t+1}\right)^2 - \left(\mathbb{E}_t C_{T,t+1}\right)^2 \left(\mathbb{E}_t \Gamma_{t+1}\right)^2.$$

Use Lagrange multipliers (λ_t, μ_t, δ_t) for the three remaining constraints:

$$\begin{aligned}\mathcal{L} = \sum_{t-0}^{\infty} \beta^t \Bigg\{ &\left[\gamma \log C_{Tt} - (1-\gamma)\left(\log \Gamma_t + \frac{1}{A_t \Gamma_t}\right)\right] \\ &+ \lambda_t \left[B_{t-1}^* + Y_{Tt} - C_{Tt} - \frac{B_t^*}{R_t^*}\right] + \mu_t \left[1 - \omega\sigma_t^2 \frac{N_t^* - B_t^*}{R_t^*} - \beta R_t^* \mathbb{E}_t \frac{C_{Tt}}{C_{Tt+1}}\right] \\ &+ \delta_t \left[\beta^2 C_{Tt}^2 \left(\mathbb{E}_t \Gamma_{t+1}\right)^2 \sigma_t^2 + \left(\mathbb{E}_t C_{T,t+1}\right)^2 \left(\mathbb{E}_t \Gamma_{t+1}\right)^2 - \mathbb{E}_t \left(\Gamma_{t+1} C_{T,t+1}\right)^2\right]\Bigg\}.\end{aligned}$$

Note that μ_t has the same sign as σ_t^2 $(N_t^* - B_t^*)$ so that $\mu_t\, \sigma_t^2\, (N_t^* - B_t^*) \geq 0$ and $\delta_t \geq 0$, with equalities only if $\sigma_t^2\, (N_t^* - B_t^*) = 0$. Also note that $\mathbb{E}_t$ in the Lagrangian stands for $\Sigma s_{t+1}\, \pi_t\, (s_{t+1})$ where $\pi_{t+1} = \pi_t(s_{t+1})$ is the probability of state s_{t+1} at t+1 conditional on state s_t at t. We take FOCs with respect to σ_t^2 and Γ_{t+1} in state s_{t+1}:

$$-\mu_t \omega \frac{N_t^* - B_t^*}{R_t^*} + \delta_t \beta^2 C_{Tt}^2 \left(\mathbb{E}_t \Gamma_{t+1}\right)^2 = 0$$

$$\beta\pi_{t+1}\left(1-\gamma\right)\frac{1}{\Gamma_{t+1}}\left(\frac{1}{A_{t+1}\Gamma_{t+1}} - 1\right)$$

$$+ 2\delta_t \pi_{t+1}\left[\left(\beta^2 C_{Tt}^2 \sigma_t^2 + \left(\mathbb{E} C_{T,t+1}\right)^2\right)\left(\mathbb{E}_t \Gamma_{t+1}\right) - C_{T,t+1}^2 \Gamma_{t+1}\right] = 0 .$$

Simplify and rewrite:

$$\delta_t \beta^2 C_{Tt}^2 (\mathbb{E}_t \Gamma_{t+1})^2 = \mu_t \omega \frac{N_t^* - B_t^*}{R_t^*},$$

$$\beta\left(1-\gamma\right)\left(\frac{1}{A_{t+1}\Gamma_{t+1}} - 1\right)$$

$$= 2\delta_t \left[\left(\Gamma_{t+1} C_{T,t+1}\right)^2 - \left(\beta^2 C_{Tt}^2 \sigma_t^2 + \left(\mathbb{E}_t C_{T,t+1}\right)^2\right)\left(\mathbb{E}_t \Gamma_{t+1}\right)\Gamma_{t+1}\right].$$

Next take the expectation $\mathbb{E}_t$ of the second condition and use the definition of σ_t^2 to simplify:

$$\beta\left(1-\gamma\right)\mathbb{E}_t\left(\frac{1}{A_{t+1}\Gamma_{t+1}} - 1\right)$$

$$= 2\delta_t [\mathbb{E}_t (\Gamma_{t+1} C_{T,t+1})^2 - \left(\beta^2 C_{Tt}^2 \sigma_t^2 + \left(\mathbb{E}_t C_{T,t+1}\right)^2\right)\left(\mathbb{E}_t \Gamma_{t+1}\right)^2\Big] = 0$$

as the RHS corresponds to the definition of σ_t^2. Thus, average output gap is zero, $\mathbb{E}_t X_{t+1} = 1$.

Now substitute out δ_t:

$$\beta\left(1-\gamma\right)\left(\frac{1}{A_{t+1}\Gamma_{t+1}} - 1\right) = \frac{2\omega\mu_t}{\beta^2}\frac{N_t^* - B_t^*}{R_t^*}\left[\frac{(\Gamma_{t+1} C_{T,t+1})^2}{C_{Tt}^2 (\mathbb{E}_t \Gamma_{t+1})^2} - \frac{\mathbb{E}_t (\Gamma_{t+1} C_{T,t+1})^2}{C_{Tt}^2 (\mathbb{E}_t \Gamma_{t+1})^2}\frac{\Gamma_{t+1}}{\mathbb{E}_t \Gamma_{t+1}}\right],$$

where we used:

$$\mathbb{E}_t \frac{C_{Tt}}{C_{T,t+1}} \frac{\mathcal{E}_t}{\mathcal{E}_{t+1}} = \mathbb{E}_t \frac{W_t}{W_{t+1}} = 1/\left(\beta R_t\right).$$

Rewrite in terms of C_{Nt} and $\mathcal{E}_t$:

$$\beta\left(1-\gamma\right)\left(\frac{C_{N,t+1}}{A_{t+1}} - 1\right) = \frac{2\omega\mu_t}{\beta^2} \frac{\mathcal{E}_t\left(N_t^* - B_t^*\right)}{R_t^*}$$

$$\left[\frac{\left(\frac{1}{\mathcal{E}_{t+1}}\right)^2}{(\mathbb{E}_t \frac{C_{Tt}}{C_{T,t+1}} \frac{\mathcal{E}_t}{\mathcal{E}_{t+1}})^2} - = \frac{\mathbb{E}_t\left(\frac{1}{\mathcal{E}_{t+1}}\right)^2}{(\mathbb{E}_t \frac{C_{Tt}}{C_{T,t+1}} \frac{\mathcal{E}_t}{\mathcal{E}_{t+1}})^2} \frac{\left(\frac{1}{\mathcal{E}_{t+1} C_{T,t+1}}\right)}{\mathbb{E}_t\left(\frac{1}{\mathcal{E}_{t+1} C_{T,t+1}}\right)^2}\right],$$

and further simplify by noting that:

$$\mathbb{E}_t \frac{C_{Tt}}{C_{T,t+1}} \frac{\mathcal{E}_t}{\mathcal{E}_{t+1}} = \mathbb{E}_t \frac{W_t}{W_{t+1}} = 1/\left(\beta R_t\right)$$

$$\beta\gamma C_{N,t+1}\left(X_{t+1} - 1\right) = \frac{2\mathcal{E}_t R_t^2 \mu_t \ \hat{\Psi}_t}{\sigma_t^2}\left[\frac{C_{T,t+1}}{\mathcal{E}_{t+1}} - \frac{\mathbb{E}_t(1/\mathcal{E}_{t+1})^2}{\mathbb{E}_t 1/\left(\mathcal{E}_{t+1} C_{T,t+1}\right)}\right].$$

Therefore, monetary policy uses variation in output gap X_{t+1} around 1 to increase $C_{N,t+1}$ above A_{t+1} when $C_{T,t+1}$ is particularly high, and vice versa, to reduce the volatility of the exchange rate $\mathcal{E}_{t+1}$ α $C_{N,t+1}/ C_{T,t+1}$, thus bringing down σ_t^2 and the period t risk-sharing wedge $\hat{\Psi}_t$, in particular in periods where UIP deviations are large to begin with.

International Risk Spillovers: Implications for Emerging Markets' Monetary Policy Frameworks with an Application to Chile

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Among the factors behind international spillovers, U.S. monetary policy developments retain a major influence. Such developments drive the global financial cycle as strongly demonstrated by Rey (2013), Miranda-Agrippino and Rey (2020), Miranda-Agrippino and Rey (2021). The dramatic U.S. monetary easing during the early months of the Covid-19 pandemic was the single most important factor for the reversal of capital outflows to emerging markets and developing economies.[1] As shown by Kalemli-Özcan (2019), the transmission mechanism for monetary policy spillovers to emerging market economies (EMEs) rests on the effect of U.S. monetary policy on investors' risk sentiments, as those sentiments are more volatile in the case of EMEs. In Kalemli-Özcan (2019), I show that capital flows to emerging markets are particularly "risk-sensitive." This creates a challenge unique to the EME policymakers and their monetary policy frameworks.

Building on and updating my prior work, in this paper I argue that EME policymakers should smooth out this risk sensitivity by not using policy rates but other policy tools instead. A good barometer of this risk sensitivity is the uncovered interest rate parity (UIP) risk premia and, if EME policymakers use policy rate to respond to U.S. monetary policy changes, the UIP risk premia increase further.

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1. See Kalemli-Özcan (2021), Obstfeld (2021).

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This happens because, when U.S. tightens, an emerging market that wants to use monetary policy to limit exchange-rate volatility needs to implement a much larger increase in the domestic policy rate, since U.S. tightening increases the UIP risk premia. Such a large increase in the policy rate can be counterproductive by increasing risk premia further through higher credit costs, spreads, and country risk with dire consequences for the real economy. As a result, the case for flexible exchange rates is stronger under international risk spillovers, since floating exchange rates help to smooth out the UIP risk premia, thus freeing domestic monetary policy's hand to focus on inflation targeting and output stabilization.[2]

Countries may want to limit exchange-rate volatility because of the negative effects of excessive volatility on balance sheets due to extensive debt denominated in foreign currency and/or a high degree of passthrough of currency depreciations to inflation. Calvo and Reinhart (2002) documented a pervasive "fear of floating," where the 'fear' is linked to liability dollarization in EMEs. Burstein and Gopinath (2014) have documented a higher degree of inflation passthrough in EMEs relative to advanced countries. Recent research shows that monetary policy credibility helps to reduce the high degree of passthrough from exchange-rate fluctuations to inflation.[3]

For "fear of floating" linked to foreign-currency debt, countries can limit the extent of foreign-currency debt by using countercyclical prudential policies. Macroprudential and capital-flow management policies can be used countercyclically in a transitory way, to limit unhedged foreign-currency-denominated liabilities not only in the financial sector, as typically done, but also in the nonfinancial corporate sector.[4] The rationale for these policies is to reduce foreign-currency-denominated debt and hence to provide insulation from spillovers that arise from balance-sheet effects of exchange-rate fluctuations with large levels of unhedged foreign-currency-denominated debt. In a monthly panel of over 40 emerging markets since the 2000s, Das and others (2022) find evidence that countercyclical preemptive

2. See Akinci and Queralto (2019) for a model of spillovers from the U.S. to a small open economy with UIP deviations where welfare gains are higher under floating exchange rates. See also Kalemli-Özcan (2019) who shows that free floating EMEs do not experience the negative effects of VIX shocks on GDP growth, whereas EMEs with managed floats do.

3. For example, López-Villavicencio and Mignon (2017), and Carrière-Swallow and others (2021).

4. For example, Basu and others (2020).

macroprudential and capital-flow-management policies reduce foreign-currency debt accumulation and hence lower the UIP premia during risk-off shocks. For the long term, improvements in the quality and transparency of institutions will reduce idiosyncratic country risk and reduce the sensitivity of capital flows in EMEs to global risk premia and foreign investors' risk perceptions. These policies will also provide the credibility needed for implementing desirable countercyclical macroprudential and capital-flow management policies to dampen the international risk spillovers.

In section 1, I document the strong correlation between risk sentiments of foreign investors and capital inflows to EMEs. Section 2 investigates the effects of exogenous shocks to U.S. monetary policy on EMEs' risk premia, UIP and covered interest rate parity (CIP) deviations, and EMEs' domestic monetary policy response. Section 3 presents the case of Chile and Section 4 concludes.

1. Risk Sentiments and Capital Flows

There is a large literature that shows that, in EMEs, net capital flows—capital inflows by foreigners (liabilities) minus capital outflows by domestic residents (assets), equivalently the current account with a reverse sign—are mainly driven by the actions of foreign investors, that is, by the liabilities side or inflows.[5] These capital inflows can be positive or negative during any given quarter, as foreign investors can increase or reduce their financial exposures to a given country. Thus, I focus on these 'gross' capital inflows—what foreigners bring in and what they take out—from the International Monetary Fund's (IMF) Balance of Payments database. Capital inflows are reported both in total and by their components: Foreign Direct Investment (FDI) flows, Portfolio Equity flows, Portfolio Bonds flows, and Other Investment flows. As shown by Avdjiev and others (2022), the largest component of capital flows is debt flows (portfolio bond flows and other investment flows), both for advanced economies (AEs) and EMEs. In addition, global financial intermediaries have an important role in intermediating capital flows between countries (as opposed to direct access to equity markets in lender countries by borrower countries). For these reasons, I focus on total debt flows. From the IMF, Balance of Payments Statistics, and from Avdjiev and others (2022), I have

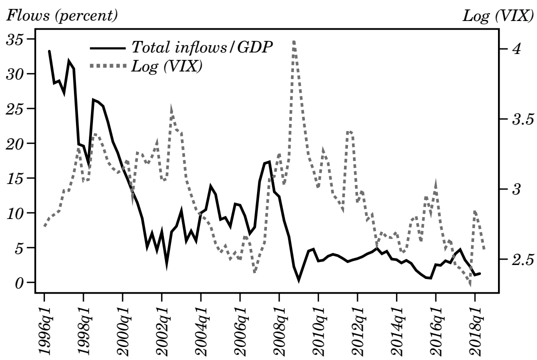
5. Bluedorn and others, 2013; Avdjiev and others, 2022.

quarterly data on capital inflows from 1996 until the end of 2019 for 55 EMEs.

How big of a role do risk sentiments play for capital inflows into EMEs? As shown in chart 1—updated from Kalemli-Özcan (2019) till the end of 2019, plotting the relationship between the VIX and capital inflows into EMEs—,[6] the VIX has an important negative effect on capital inflows to EMEs. Miranda-Agrippino and Rey (2021) show that this relationship is explained by the same global factor that explains the global financial cycle.

The mapping from the changes in U.S. monetary policy to the VIX is not straightforward. As shown by Bekaert and others (2013), Miranda-Agrippino and Rey (2019), and Bruno and Shin (2015), a higher U.S. rate increases the VIX. However, as shown by Rey (2013), there is a feedback effect, and a higher VIX induces an expansionary U.S. policy. As shown above, capital inflows to EMEs move with the VIX, and Kalemli-Özcan (2019) also showed that EME domestic monetary policy responds to such movements. Hence, I will investigate the effect of *exogenous* U.S. monetary policy shocks on risk premia in EMEs, next.

Figure 1. Risk Sentiments and Capital Flows in EMEs



Source: Author's calculations.
Notes: Capital flows are normalized by GDP and plotted as three-quarter moving averages, and these flows are averaged across countries on a given date.

6. The VIX is a forward-looking volatility index of the Chicago Board Options Exchange. It measures U.S. investors' expectation of 30-day volatility and is constructed by using the volatilities implied by a wide range of S&P 500 index options. This chart is updated from Kalemli-Özcan (2019).

2. U.S. Monetary Policy, Risk Premia, and Policy Response

The conventional models imply that domestic credit costs should respond to monetary policy actions, and this response should depend on the expected path of the central bank's policy instrument, which is the short-term interest rate. Gertler and Karadi (2015) argue that, in the presence of financial frictions, the response of credit costs to monetary policy may in part reflect movements in term premia and credit spreads. By using high-frequency identification (surprises in Fed-funds futures occur on FOMC days in a thirty-minute window of the monetary policy announcement), they can rule out the simultaneity of economic news and monetary policy and hence prevent risk premia being 'priced-in' before the announcement.

I use these U.S. monetary policy surprises in a local projections framework—which is shown to estimate the same impulse response functions (IRFs) as the VAR by Plagborg-Møller and Wolf (2021). I implement local projections with instrumental variables (LP-IV) following Jorda (2005), and Stock and Watson (2018). I run the following regressions for EMEs:

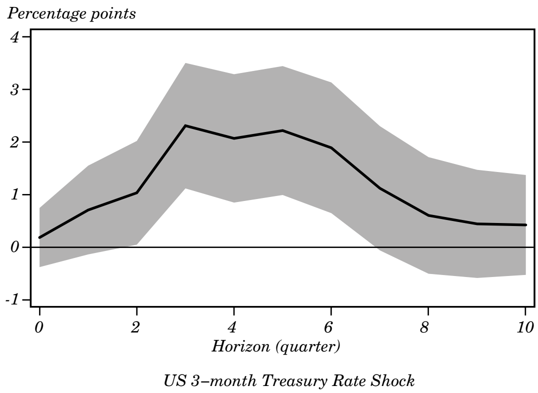
$$(i_{c,t+h} - i_{US,t+h}) = \alpha_c + \beta_h \hat{\imath}_{US,t} + \beta_h^{\omega} W + \varepsilon_{c,t+h}, \quad h = 0,1,2,3...$$

where $(i_{c,t+h} - i_{US,t+h})$ is the 12-month government bond spreads at time $t+h$ in a given country c, vis-a-vis the U.S. α_c is a country-fixed effect, $\hat{\imath}_{US,t}$ is the estimated exogenous U.S. monetary policy shock at time , and β_h is the associated impulse response coefficient.

Chart 2 presents the results, updated from Kalemli-Özcan (2019). In EMEs, spreads increase by 2.2 percentage points after three quarters in response to a 1 percentage point increase in U.S. monetary policy rate.[7] Notice that these are short-term spreads, which means that the risk spillovers do not necessarily come from the term premia in EME. Degasperi and others (2020) show that a similar mechanism is at work for advanced countries working via term premia at the long end of the yield curve. Gourinchas and others (2021) also show increasing government and corporate-bond spreads in EMEs as a response to U.S. monetary policy contractions, where they use the policy surprises directly.

7. Kalemli-Özcan (2019) shows that the pattern is opposite for advanced countries; the spreads decrease by about 0.5pp after one quarter and 1.7pp after six quarters.

Figure 2. Responses of 12-month EME Government Bond Spreads to U.S. Monetary Policy Shocks

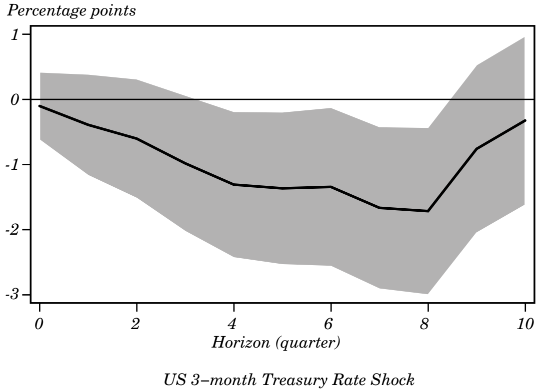


Source: Author's calculations.
Notes: Impulse responses are obtained from panel local projections of 79 EMEs. 95 percent confidence intervals (calculated by using Newey-West standard errors) are shown by the shaded areas. The U.S. policy (3-month treasury rate) is instrumented by Gertler-Karadi shock FF4 (estimated from surprises in 3-month Fed-fund futures).

How does EME domestic monetary policy respond? We cannot be sure if the patterns above are due to rising risk premia of EMEs or a procyclical response of EME domestic monetary policy to contractionary U.S monetary policy.[8] Chart 3 shows that this is not the case. On the contrary: EMEs, on average, run a countercyclical monetary policy as a response to contractionary U.S. monetary policy and lower their policy rates. Consistent with the findings of Kalemli-Özcan (2019), who shows a short rate disconnect—less than full passthrough between monetary policy rates and short-term market interest rates—in EMEs, here also it is clear that EME monetary policy can be ineffective in smoothing the risk premia. Although EME policymakers lower the policy rates as a response to a contractionary U.S. monetary policy shock, on average, EME risk premia still rise.

8. I would like to thank Helene Rey, who raised this point during her discussion of Kalemli-Özcan (2019)

Figure 3. Responses of EME Policy Rates to U.S. Monetary Policy Shocks



Source: Author's calculations.

3. U.S Monetary Policy and UIP Risk Premia

I argue that a good barometer to measure the relation between the U.S. monetary policy and changes in EME risk premia is the fluctuations in UIP premia, that is, dynamic UIP deviations. The standard UIP condition can be stated as follows:

$$E_t[S_t + h]\ (1 + i_{US,t}) = S_t\,(1 + i_{c,t}), \tag{1}$$

where t denotes time and h is the horizon considered. S_t and $E_t[S_{t+h}]$ are the spot exchange rate at time t and the expected (as of time t) exchange rate for h months ahead, respectively. The exchange rate is denominated in units of local currency per U.S. dollar. In turn, $i_{c,t}$ and $i_{US,t}$ are the domestic and U.S. interest rates with the same time horizon for the maturity of the debt as the expected exchange rate. By using equation (1), we express the UIP deviation in logs as,

$$\lambda \equiv i_{c,t} - i_{US,t} - [s^e_{t+h} - s_t], \tag{2}$$

where λ_t denotes the UIP deviation for the domestic currency with respect to the U.S. dollar. Under this specification, a λ_t equal to zero implies that the UIP condition holds and interest-rate differentials and expected exchange-rate movements offset each other fully. Otherwise, if there are positive UIP deviations, there are positive expected excess returns on the domestic currency.

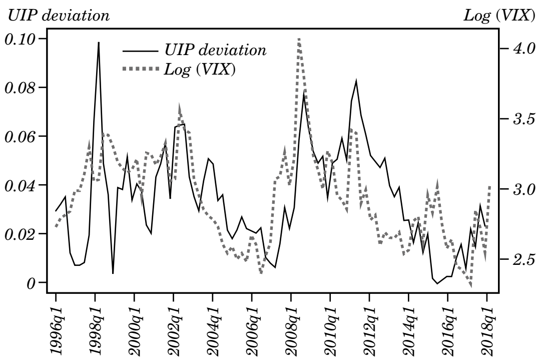
Chart 4 plots the median UIP deviation, λ, for EMEs—reproduced from Kalemli-Özcan and Varela (2019).[9] The correlation is over 60 percent.

The reason why UIP premia are a good barometer is that, in EMEs, UIP deviations move with policy credibility-related country-specific risk, which is captured by interest-rate differentials, while in advanced countries, they move with global risk, captured by exchange-rate fluctuations, as shown in Kalemli-Özcan and Varela (2019) and replicated below:

$$\lambda_t \equiv \underbrace{i_{c,t} - i_{US,t}}_{\text{IR Differential}} + \underbrace{s_t - s^e_{t+h}}_{\text{ER Adjustment}}. \tag{3}$$

Chart 5 plots each part of this decomposition. The sources of the UIP deviations differ greatly as argued.

Figure 4. Risk Sentiments and UIP Deviations



Source: Author's calculations.

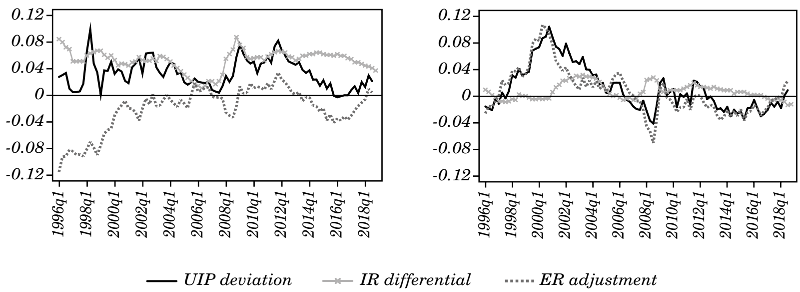
Notes: The figure plots UIP deviations using quarterly observations from 22 EMEs excluding hard pegs. The sample size is lower due to availability of data on expectations of exchange rates, which are obtained from Consensus Forecast. The UIP deviation is calculated as the difference between log interest-rate differentials and the gap between log expected and spot exchange rate. Log interest-rate differentials are the deposit rate differentials vis-a-vis the U.S. The log expected exchange rate is the 12-month ahead expected exchange rate as of month t and the log exchange rate is the spot rate, both nominal and in terms of local currency per U.S. dollar.

9. The dynamic relation between UIP and VIX is first shown by Di Giovanni and others (2022) for Turkey.

These different sources of the UIP deviations underlie the relation between the UIP risk premia and the U.S. policy shocks in EMEs. Chart 6 shows that the UIP deviations in EMEs increase by about 3 percentage points after two quarters in response to a 1 percentage point contractionary U.S. policy-rate shock in EMEs, whereas there is no response in AEs. The response of the UIP deviations implies that EMEs need to provide additional returns to investors to compensate for heightened country risk induced by the contractionary U.S. monetary shock. Hence, global investors expect and earn excess returns from EMEs.

Figure 5. Sources of UIP Deviations

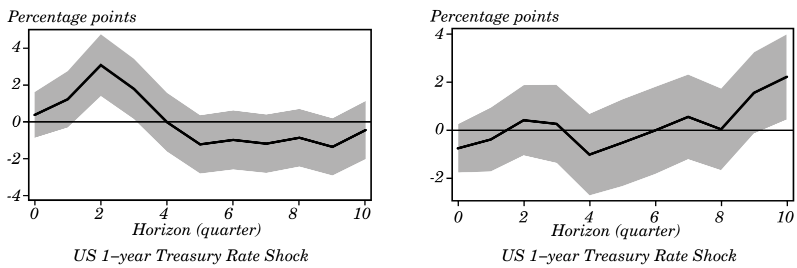
(*a*) *Emerging market economies* (*b*) *Advanced economies*



Source: Author's calculations.
Notes: The figure plots UIP deviations and components using quarterly observations from 22 EMEs and 12 AEs excluding hard pegs.

Figure 6. Responses of UIP Risk Premia to U.S. Monetary Policy Shocks

(*a*) *Emerging market economies* (*b*) *Advanced economies*

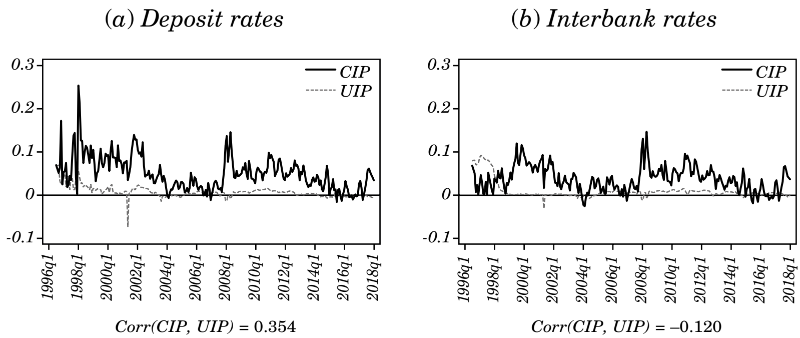


Source: Author's calculations.
Notes: Impulse responses of UIP deviations are obtained from panel local projections. The standard errors are Newey-West and given by the shaded areas. The U.S. 12-month treasury rate is instrumented by using the Gertler-Karadi policy shock (estimated from surprises in 3-month Fed-fund futures).

These movements in UIP premia with the U.S. monetary policy underlie the strong case for flexible exchange rates in EMEs. In terms of equation (2), it is easy to see that, when the U.S. interest rates rise, if also risk premia rise, the domestic monetary policy needs to adjust by raising the policy rates by a large margin if the domestic monetary authority also wants to stabilize the exchange-rate fluctuations. This will not be the case if UIP holds. If UIP holds, there is no role for risk premia in driving the procyclicality in UIP deviations and, although a central bank that wants to stabilize the exchange rates needs to increase the policy rate as a response to U.S. tightening, this increase does not have to be that big. By increasing domestic rates by a large margin, domestic monetary policy not only hurts the domestic economy but also has an impact on country-risk premium through tighter financial conditions, thus increasing the effects of international risk spillovers.

Can UIP deviations be capturing CIP deviations? CIP deviations stem from breaks in the arbitrage condition and can be related to UIP deviations. A UIP wedge can be there even in the absence of such a break in arbitrage if it is driven by risk premia as shown by Akinci and others (2022). This might be the reason why UIP failures are much larger than CIP failures, as they are driven by fluctuations in risk premia and not necessarily a break in arbitrage, as shown in chart 7, replicated from Kalemli-Özcan and Varela (2019).

Figure 7. UIP vs CIP Deviations

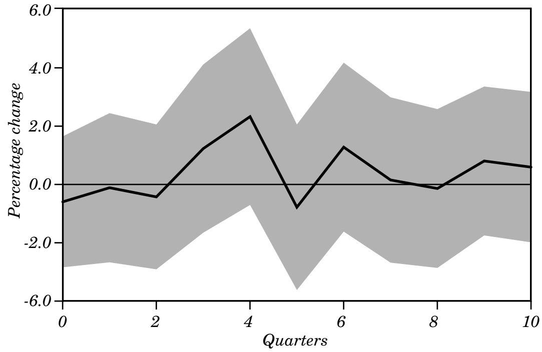


Source: Author's calculations.
Notes: The figure plots the same UIP deviations as before. CIP deviations are calculated by following Du and Schreger (2021) in panel (b) and using deposit rates as in UIP in panel (a).

4. The Case of Chile

In this section, I focus on the case of Chile. Chile has been a successful inflation targeter under a floating exchange-rate regime for some time. This means that we expect Chile's risk premia not to respond systematically to U.S. monetary policy shocks, and hence, our barometer UIP risk premia will also be unresponsive to these shocks. As shown in chart 8, this is exactly what we have found for Chile.

Figure 8. UIP Premia Responses in Chile to U.S. Monetary Policy Shocks



Source: Author's calculations.
Notes: This figure shows the response of the UIP premia to a U.S. monetary policy shock. For data restrictions, we use 12-month deposit rates instead of 12-month treasury rates as the relevant rates to construct the UIP premia. All other controls are the same. The time period spans from 1996.IV to 2016.IV.

5. Conclusion

U.S. monetary policy actions have the potential to spill over to any country as long as international investors' risk perceptions change with changes in U.S. monetary policy. Central bankers are increasingly confronted with the need to better understand and respond to fluctuations related to shifts in risk sentiments, and this can lead to disruptive financial conditions.

I show that UIP deviations are good barometers of such changes in risk sentiments in emerging markets as a response to changes in U.S. monetary policy. Domestic monetary policy can be ineffective under significant UIP deviations in emerging markets; that is, even if domestic policy responds countercyclically to changes in the U.S. policy, it cannot reduce the risk premia. Floating exchange rates can make monetary policy more effective by smoothing out these UIP risk premia. Chile is a case in point. Chile's UIP risk premia do not respond to changes in U.S. policy as its floating exchange rate absorbs such shocks.

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Global Drivers and Macroeconomic Volatility in EMEs: A Dynamic-Factor, General-Equilibrium Perspective

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A common view held by academics as well as policymakers assigns an important role to global factors as drivers of fluctuations in economic activity in emerging market economies (EMEs). This follows naturally from the fact that these economies are often small and open to trade in global goods and capital markets, which makes them vulnerable to shocks in these markets. However, the nature of these global forces as well as their transmission mechanism into EMEs continues to be

Research for this work was conducted while Andrés Fernández was affiliated with the Central Bank of Chile. The information and opinions presented are entirely those of the authors, and no endorsement by the Central Bank of Chile, the IMF or their Board Members or IMF Management is expressed or implied. All results have been reviewed to ensure no confidential data are disclosed.

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debated and is the subject of an active research area in international macroeconomics. While an influential view postulates a financial origin in the form of a global financial cycle (Miranda-Agrippino and Rey, 2020), others have argued in favor of alternative global forces in the form of fluctuations in commodity prices (Fernández and others, 2017, 2018, 2020), changes in sovereign risks (Longstaff and others, 2011; Aguiar and others, 2016), and a common growth factor among EMEs (Claessens and others, 2012).

This paper aims at identifying the global forces that matter the most for EMEs, how they are interrelated, and the way they shape the business cycle in these economies. Our strategy is divided into two steps. First, we estimate a global dynamic factor model by using data from a set of EMEs as well as other variables from advanced economies, and international prices in goods and financial markets. Importantly, given the array of alternative origins of these global forces, our identification assumptions encompass the different views in the literature by allowing for three distinct global factors to coexist: a *financial* factor that captures the comovement of financial variables across countries; a *price* factor that accommodates joint movements in commodity, import prices, and CPIs; and a *growth* factor that captures any further comovement in GDP across EMEs that the aforementioned forces cannot explain and may come, for instance, from common variations in total factor productivity.

While the global dynamic factor model is enough to obtain a proper identification of the three factors and the way they are interrelated, it cannot provide a detailed analysis of the transmission mechanism of shocks to these factors in the EMEs considered. For that purpose, the second step of our analysis zooms in on Chile—one of the countries in our sample of EMEs—and embeds the dynamic factor model as another layer of the Extended Model for Analysis and Simulation (XMAS), which is the large-scale DSGE model used regularly at the Central Bank of Chile for policy analysis and forecasting (García and others, 2019). This allows us to combine the estimated comovement of the global forces pinned down by the dynamic factor model with the rich structure of the DSGE, thereby providing us with an appropriate setup to analyze the transmission mechanism of global disturbances into the Chilean economy. In addition, because the enlarged model inherits the estimated Taylor rule, we can study the way in which changes in global factors trigger monetary policy responses.

Our work highlights three main findings. First, the three estimated global factors display strong comovement, with a preponderance of the financial factor affecting the two other ones. Indeed, a shock to the financial factor—akin to a relaxation of global financial conditions—induces a *risk-on*-type of (delayed) response in the other factors whereby growth in EMEs rises and prices increase. A shock to the price factor, on the other hand, is consistent with a global cost-push shock that triggers a contraction of the growth factor along with price factor hikes and a fall in the financial factor. Shocks to the growth factor have relatively modest effects on the other two factors.

Second, consistent with the conventional wisdom that global forces matter for EMEs, we find that the three identified factors explain an important share of the business cycle in the sample of EMEs considered. Indeed, they account for more than a third of the variance in GDP (39%), of which the financial and price factors explain the majority and the growth factor explains a relatively more modest share. The factors also have the ability to explain an important share of the variance of sovereign risk across the sample EMEs (24%) and even more of their stock-market indices (67%), with the financial factor accounting for the lion's share. Lastly, shocks to the three estimated factors account for a strikingly high share of the variance of the other global variables considered, like GDP and CPIs of EMEs' trading partners (39% and 43%, respectively), import price indices (43%), exchange rates against the U.S. dollar (49%), and world commodity prices (30%). Once again, shocks to the global financial and price factors appear as the main driving force behind this comovement in global variables.

Following a shock to the estimated global financial factor, EMEs' GDP increase, EMBIs fall while stock markets boom, inflation accelerates (with a delay) fueled by swelling import prices along with hikes in the prices of the main commodities exported. In contrast, a shock to the price factor increases the price of imports more than the price of the main commodity exported, which triggers a boost in inflation, a slowdown in economic activity and stock markets, and a rise in sovereign risks. Lastly, a shock to the growth factor that boosts GDP across EMEs implies only modest expansions in inflation and stock-market activity, and even milder drops in EMBIs. Our main results carry on with plausible alternative identification assumptions. Even when we rule out a contemporaneous effect of the financial factor on EMEs' GDP, we still get its already documented preponderant role. This shows, perhaps surprisingly, that global financial forces have the ability to affect economic activity in EMEs regardless of the modeling stance on the contemporary, direct link between them and economic activity.

Our third key result relates to the transmission mechanism of global factors to domestic EMEs' variables. The baseline factor model also allows us to quantitatively assess the relative importance of global factors to both global and domestic variables: while the financial factor explains the most significant part of the variance of global variables, in the case of growth and inflation rates of EMEs, the global price factor entails a comparable role.

The augmented DSGE model for the Chilean economy allows us to study those results more closely. A key finding from the analysis reveals that the relevance of the global financial factor in affecting domestic variables gets dampened, while the opposite happens regarding the global price factor. In order to grasp this contrasting result, we first note that the transmission channel from global factors to domestic variables in the model is not direct but operates through other global variables, such as commodity prices and global demand. Hence, the ultimate role played by factors on the dynamics of domestic variables hinges subsequently on the extent to which shocks to these factors affect global variables, which only then translates into EMEs' performance. Therefore, while a shock to the global financial factor triggers movements in global variables that steer domestic variables in opposing directions, after a global price shock, in contrast, such offsetting effect in domestic variables is no longer present.

The quantitative features of the way in which domestic EMEs' variables correlate with shocks to global forces have relevant policy implications for these economies. In contrast to shocks to the financial factor, monetary policy should react more strongly to price shocks: even though global variables react individually less in this latter case, they all push the economy in the same direction, which ends up calling for a bolder monetary policy response.

The rest of the paper is divided into three sections: Section 1 presents results from the estimated dynamic factor model. Section 2 embeds the dynamic factor structure into the Chilean large-scale DSGE model. Concluding remarks are presented in section 3. Additional material is gathered in the Appendices.

1. A Structural Factor Model

When building the dynamic factor model, we are guided by the literature on global macroeconomic forces shaping the business cycle of EMEs: we postulate a set of common global factors that encompass the various views from the literature. Indeed, regarding the global forces that previous research has documented, the cornerstone pieces involve a global financial cycle (Miranda-Agrippino and Rey, 2020), the price of commodities (Fernández and others, 2017; Fernández and others, 2018; Fernández and others, 2020), sovereign debt spreads (Longstaff and others, 2011; Aguiar and others, 2016), and growth factors (Kose and others, 2012).

Building upon this literature, our modeling strategy writes down our panel dataset as a linear function of three unobserved common factors that, without loss of generality, we associate with financial, price, and growth forces. Crucially, our approach is nonetheless agnostic in terms of how relevant each factor is and the extent to which the three factors are interrelated. By estimating the model, we let the data speak on these issues.

We impose some structure on the contemporary behavior of factors in the estimation stage of a state-space formulation with parameter constraints. More precisely, we impose constraints on the loading matrix of the observation equations. Thus, by limiting the effects of certain factors on, say, commodity prices or financial variables, we are able to associate these factors with certain subsets of the time series data observed. Therefore, our approach allows for the estimation of a set of common factors with an ex-ante association to specific macroeconomic phenomena.

1.1 Data

We estimate our model by using an unbalanced quarterly panel dataset between 2003Q1 to 2018Q4. Similar to Fernández and others (2018) and Bajraj and others (2021), our sample includes mainly commodity-exporting EMEs, namely, Argentina, Brazil, Bulgaria, Chile, Colombia, Ecuador, Malaysia, Mexico, Peru, Russia, South Africa, and Ukraine. For each of these countries, we include a set of variables that characterize EMEs' business cycle (we call them "EME variables"), and another set with EMEs' most relevant external variables (we call them "global variables"). In the first group we include each EME's real GDP,[1] CPI,[2] EMBI Spread,[3] and major stock market

1. IMF data, except Central Reserve Bank of Peru for Peru; and OECD for Russia and South Africa.

2. IMF data, except Bloomberg for Argentina.

3. JP Morgan EMBI Global spreads, from Bloomberg. Following Aguiar and others (2016), we deflate each EME's EMBI with the country's external debt (% of GDP, from the World Bank) and GDP growth (see footnote 10).

indices.[4] In the group of global variables we include each country's import price index;[5] the prices of the top-ten commodities exported by EMEs (crude oil, copper, aluminum, natural gas, coal, iron, gold, coffee, bananas, soybean meal);[6] and real GDP, CPI, and exchange rate (local currency per U.S. dollar) of the EMEs' top-ten trading partners (namely, United States, China, Eurozone, Japan, United Kingdom, India, Korea, Taiwan, Brazil, and Mexico).[7] Additionally, Wu and Xia (2016)'s estimation of the U.S. shadow federal funds rate is included in the set of global variables.

To rule out the presence of integrated series, all the time series for GDP, CPI, stock indices, import price indices, and commodity prices enter the model in first (log) differences, while EMBIs and the shadow federal funds rate enter in first differences. All variables correspond to quarterly averages, and are centered (demeaned) and scaled by the inverse of their standard deviation.

1.2 State-Space Formulation

Let $Y_t = ((Y_{it})^N, (G_{jt})^{10}, CMDTY_t, SFFR_t)'$ denote our vector of observable time series, where $Y_{it} = (GDP_{it}, CPI_{it}, EMBI_{it}, Stock_{it}, ImportPrice_{it})'$ represents the specific variables described above for each EME $i = 1, \ldots, N$ in period $t = 1, \ldots, T$. The vector $G_{jt} = (GDP_{jt}, CPI_{jt}, FX_{jt})$ denotes the observations for each top $j = 1, \ldots, 10$ EMEs' trading

4. In U.S. dollars, as in Miranda-Agrippino and Rey (2020). We use the following indexes from Bloomberg: Merval (ARG), IBOV (BRA), SOFIX (BGR), IPSA (CHL), COLCAP (COL), ECGUBVG (ECU), FBMKLCI (MYS), MEXBOL (MEX), SPBLPGPT (PER), RTSI$ (RUS), PSI20 (ZAF) and PFTS (UKR). U.S. dollar FX are from the BIS.

5. Import price deflator, from Haver Analytics.

6. Commodity prices are from the IMF, expressed in U.S. dollars and deflated with the U.S. CPI (from St. Louis Fed). In order to select the top-ten commodity exports of this group of EMEs, we: (1) rank the commodities exported by each country by their average exports as % of GDP in the period 2003–2018 (data from UN Comtrade); (2) for each commodity, compute the average ranking (across the 12 EMEs); and (3) select the 10 commodities with the highest average ranking. The list is similar if, instead of computing the average, we use each commodity's median ranking across EMEs.

7. The series are from Haver Analytics. For Brazil and Mexico only data on ER are added, given that their GDP and CPI series are included in the group of EME variables. The EMEs' top-ten trading partners correspond to the countries with the highest average trade ranking across the EMEs (for each EME, we rank the trading partners by their average total exports to GDP in the period 2003–2018, and then, for each trading partner, we average these rankings across EMEs).

partners;[8] while the vector $CMDTY_t$ has the stacked observations for the ten commodity prices included; and *SFFR* finally represents the measure of the U.S. shadow rate already mentioned. We model the dynamics of the $(5N+36) \times 1$ vector Y_t as

$$Y_t = \Lambda F_t + u_t, \qquad t = 1,\ldots,T, \tag{1}$$

where F_t is the $q \times 1$ vector of (unobserved) factors and Λ is the $(5N+36) \times q$ matrix of factor loadings.[9]

The factors are meant to capture the common sources of variation in the observed macroeconomic variables across countries. These could be changes in global financial conditions (e.g., changes in global risk appetite or in U.S. monetary policy) which are likely to affect a wide array of variables, shocks that affect commodity prices (e.g., changes in China's investment or growth perspectives), or other changes in global conditions that typically affect EMEs' macroeconomic performance (e.g., changes in global demand, changes in the international prices of capital goods or global inflation). The vector u_t, $u_t \sim N(0,H)$, captures variability at the country-variable level associated with idiosyncratic events or measurement error.

The vector of unobserved factors F_t is assumed to follow an autoregressive process

$$F_t = \Phi F_{t-1} + w_t, \qquad t = 1,\ldots,T, \tag{2}$$

where $w_t \sim N(0,Q)$ and $F_0 \sim N(\mu_0, \Sigma_0)$. The matrices H and Q are assumed to be diagonal, while Φ is left unconstrained. We estimate the model parameters by maximum likelihood and extract the factors by using the Kalman smoother.

8. Mexico and Brazil's GDP and CPI series are excluded from G_{jt}, given that they are already included in Y_{it}. The U.S. FX series is also excluded, given that currency parities are defined with respect to the U.S. dollar.

9. Following Aguiar and others (2016), we include a set of exogenous controls for the exclusive case of spreads, so we in practice estimate

$$Y_t = \Lambda F_t + \Gamma X_t + u_t, \qquad t = 1,\ldots,T,$$

where X_t comprises a vector of zeros, except in the event where the dependent variable is a country spread, in which case we control for the pair (ΔGDP_{it}, Debt-to-GDP_{it}) for country $i = 1,\ldots,N$ in period $t = 1,\ldots,T$ and we constrain Γ so that X_{it} only affects their respective, country-specific spreads.

It should be noted that, without further restrictions, the state-space model defined by equations (1) and (2) does not allow for a structural interpretation of the estimated factors, so we impose a set of constraints on the loading matrix Λ (i.e., we set to 0 some of its entries), and therefore limit the effect of the estimated factors on the observable variables. Among the multiple constraints that could be imposed on the $(5N + 36) \times q$ matrix Λ, we restrict the analysis to those alternatives that appear the most compatible with the set of factors identified by previous research, as laid out above.

1.3 Baseline Specification

We now formally define the set of constraints on the loading matrix and provide their structural interpretation. A guiding principle that we follow is that a specific factor will be pinned down only by the set of observable variables most closely related to it. For example, the common "growth factor" that we estimate will be contemporaneously related only to the time series of GDP, either for country-specific EMEs or those of their main trading partners.

Table 1 presents the full set of restrictions in a schematic format. Column names list the factors that we wish to identify—financial, price, and growth common forces. Then, for each variable listed, we use the black and white circles to specify which factor is allowed to contemporaneously affect each variable. A white circle means that we fix the corresponding entry in Λ to be zero, whereas a black circle means that the corresponding entry is unconstrained. First, we let the 'financial' factor to impact all the variables in the model, hence the black circles in the first column. While the lack of constraints for this factor can be equivalently grasped as a 'global' common force, we will provide further evidence that we can loosely associate it to one of a financial origin. The 'price' factor, in turn, affects merely observable prices, namely commodity prices, import prices, and local CPIs. Lastly, the 'growth' factor is identified based on GDP data, which allows for the identification of a comovement between local EMEs' cycles and the GDP fluctuations of their main trading partners. We will present a variation of these choices later on.

1.3.1 Estimated Global Factors

The estimated factors, along with their historical shocks decomposition are presented in the top panel of figure 1. Since the model is estimated in log-differences, the estimated factors are interpreted in the same way. The bars portray the incidence of each shock in the dynamics of the factors. The bottom panel of the figure presents the estimated factors in levels (net of initial values) and the cumulative effect of the shocks contributions depicted above.

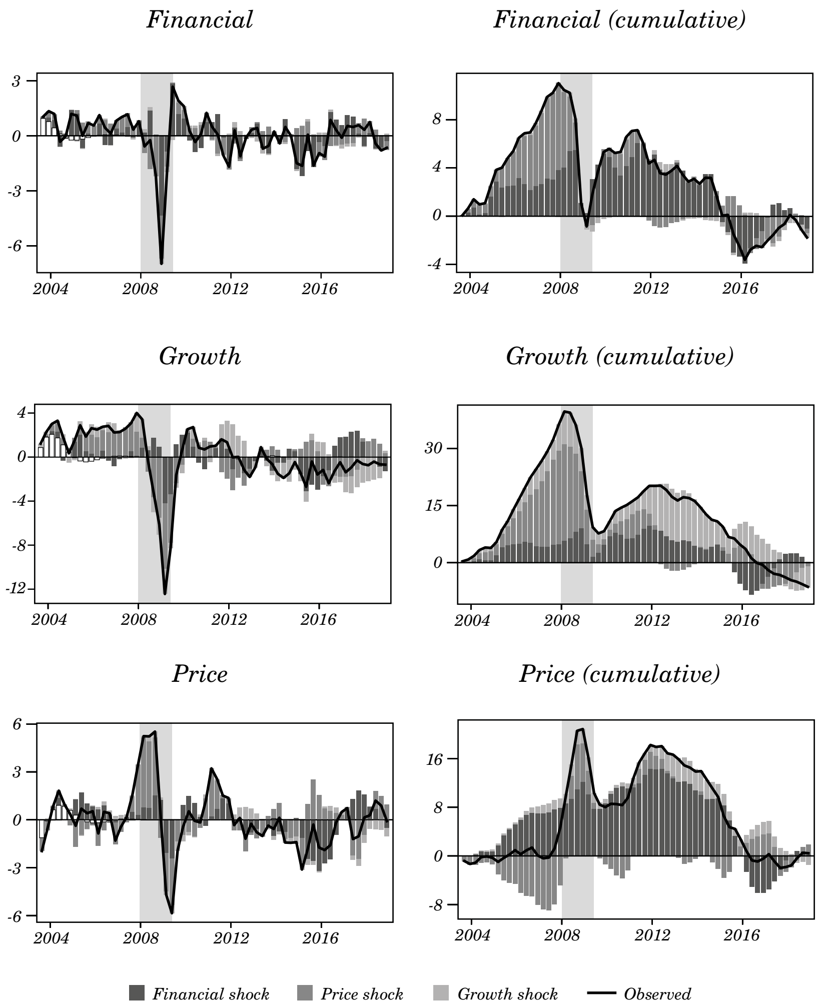
The factors' dynamics are consistent with the U.S. recession indicator as identified by NBER (shaded area), all of them experiencing very significant variability around the Global Financial Crisis (GFC). After increasing consistently in the years 2003–2007, the financial factor leads the fall during the crisis, followed by the growth factor. The price factor, on the other hand, experienced a dramatic increase between 2007 and 2008, and only fell in 2009.

Table 1. Baseline Model
(restrictions on loading matrix)

	Factor		
	Financial	*Price*	*Growth*
EME variables			
GDP EMEs	●	○	●
CPI EMEs	●	●	○
EMBI	●	○	○
Stock market index	●	○	○
Global variables			
Import price index	●	●	○
GDP trade partners	●	○	●
CPI trade partners	●	●	○
Exchange rate	●	○	○
Commodities	●	●	○
Shadow FFR	●	○	○

Source: Authors' calculations.
Notes: White circles refer entries in the Λ matrix that are set to zero, whereas black circles correspond to unconstrained entries.

Figure 1. Historical Decomposition of Factors – Baseline Model



Source: Authors' calculations.
Notes: Top panel: factors as originally estimated in log-differences (centered and scaled such that s.d.=1). Bottom panel: factors in levels obtained by cumulating log-differences. For presentation purposes, initial values are omitted in the cumulative version. Shaded areas denote NBER U.S. recession dates.

The historical shocks decomposition in figure 1 (in particular, the bottom panel) highlights a rich interaction among the estimated factors. Financial shocks not only affect the financial factor but also have significant effects on the price and growth factors. Similarly, price shocks induce important movements in both the financial and the growth factor. The level of interaction among the factors is formally quantified in table 2, which reports the share of each factor's variance explained by the different shocks. Financial shocks are the most relevant, explaining between 35 and 74pp of the factors' 20-quarter-ahead forecast error variance. On the other hand, growth shocks contribute the least, with most of their effect passing through the growth factor, and little effect on the others. Price shocks explain between a quarter and a half of the variance of each factor.

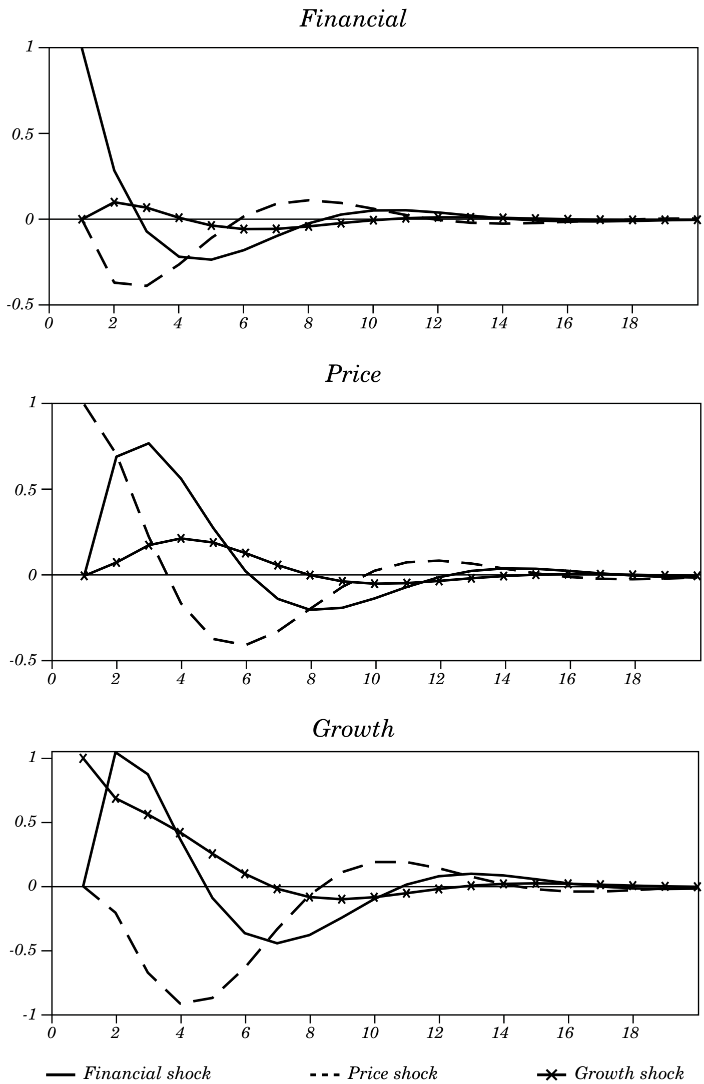
The strong comovement among factors is also reflected in their impulse responses to shocks. Figure 2a shows that, despite their relatively short persistence, shocks to the financial factor induce prominent positive responses (of comparable proportions, between 0.8 and 1 s.d.) in both the price factor and the growth factor. On the other hand, a price shock also has significant effects on financial and growth factors, but in the opposite direction. Finally, shocks to growth tend to be more persistent, but they hardly affect the dynamics of the other factors.

Table 2. Share of Factors' Variance Explained by Global Factor Shocks – Baseline

	Factor		
	Financial	*Price*	*Growth*
Financial Factor	74.3	24.2	1.5
Price Factor	42.1	53.9	4.0
Growth Factor	34.9	37.0	28.0
Average	50.5	38.4	11.2

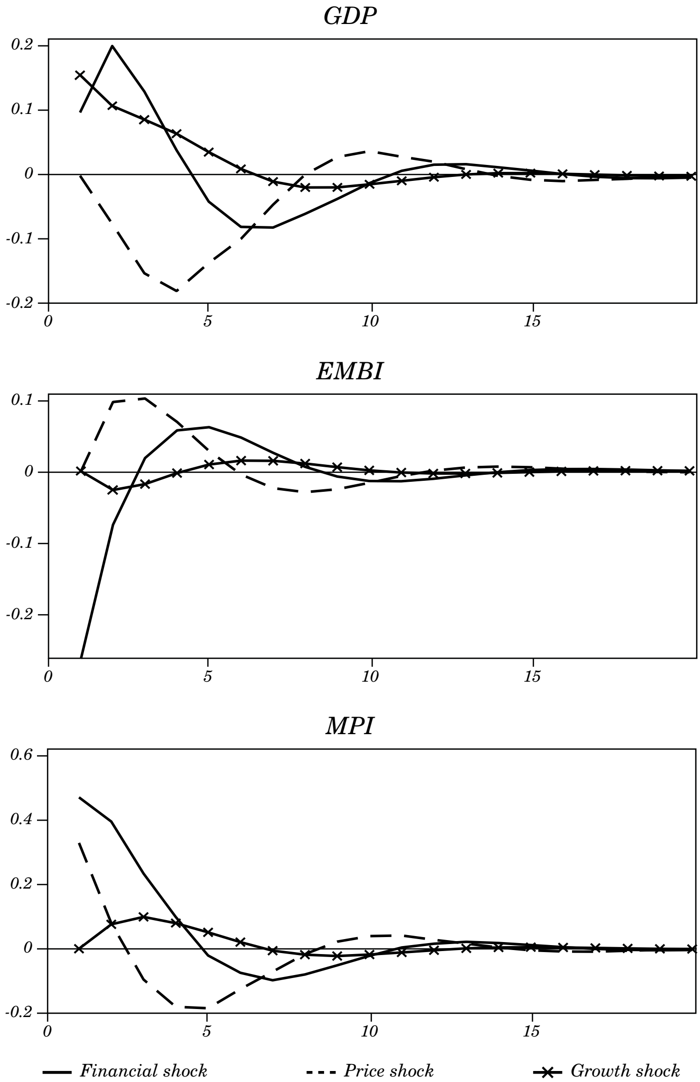
Source: Authors' calculations.
Notes: Percentage. Figures correspond to the share of the 20-period ahead forecast error variance that is attributable to each of the global factors shocks.

Figure 2a. Impulse Response Functions – Baseline Model



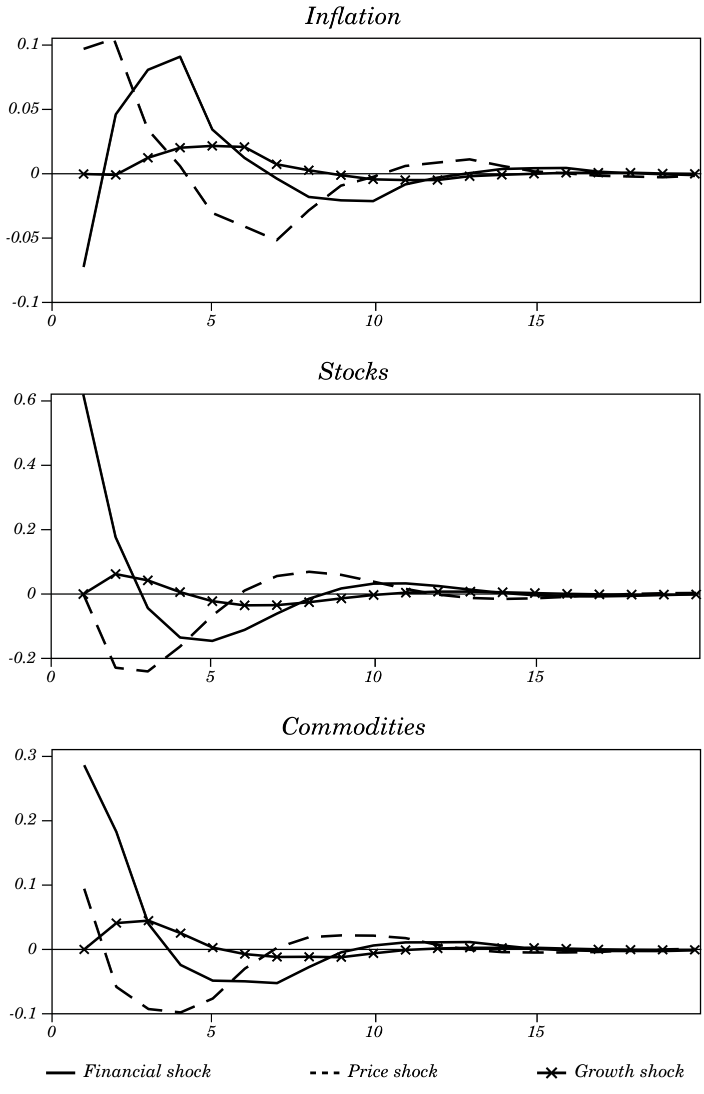
Source: Authors' calculations.
Notes: Impulse response functions to the original 'financial', 'growth' and 'price' shocks.

Figure 2b. Impulse Response Functions – Baseline Model



Source: Authors' calculations.
Notes: Impulse response functions to the original 'financial', 'growth' and 'price' shocks.

Figure 2b. Impulse Response Functions – Baseline Model (continued)



Source: Authors' calculations.
Notes: Impulse response functions to the original 'financial', 'growth' and 'price' shocks.

1.3.2 Relevance of Global Factors

We now explore the relevance of the estimated global factors when explaining the dynamics of the pool of EMEs considered and their main trading partners. Table 3 presents the results of this exploration by means of forecast error variance decomposition analysis. Together, shocks to the three global factors account for more than 38 percent of the variance in GDP of EMEs (sample median), a quarter of the variance of sovereign risks (as measured by the EMBI indices), and more than two-thirds of the variance of the stock-market indices. A more modest role is found when accounting for CPI dynamics, for which the factors explain nine percent.

At the same time, the factors explain a large share of the variance of the EMEs' most relevant external variables (i.e., "global variables")—more specifically, 39 percent of the variance of GDP, 43 percent of that of inflation, and almost 49 percent of the variance in the exchange rate of the EMEs' main trading partners. Shocks to these factors also contribute to an important fraction of the movements in commodity prices, in particular crude oil, copper, and aluminum (the top-three most exported commodities in our sample of EMEs), for which roughly two thirds of the variance is explained.

Table 3 allows us to further appreciate the individual contribution of each one of the factors to the dynamics of the different groups of variables in the model. Not surprisingly, financial shocks are the ones that contribute the most to the variance of the financial variables included in the model (EMEs' stocks and EMBIs, and trading partners' exchange rates). What might be surprising, however, is that financial shocks are also the most relevant ones for commodity prices, as well as for the GDP and inflation of the EMEs' trading partners. On the other hand, shocks to the price factor are the ones that contribute the most to explaining the variance of GDP and inflation in EMEs. We will analyze this in more detail in section 2.3.3, where we use the estimated global shocks in the context of a full DSGE model for the Chilean economy.

How do we interpret these factor shocks? Figure 2b shows that a shock to the global financial factor is associated with a *risk-on* episode when a relaxation of (global) financial conditions induces a strong positive response of EMEs' stock market indices, a reduction of sovereign risk, and a marked increase in the prices of commodities exported by these economies. These episodes also translate into higher growth and inflation in EMEs,[10] as well as into an increase in the price of imports. Price shocks, on the other hand, have very different effects on the dynamics of these emerging commodity-exporting economies: import prices and inflation increase significantly, while economic activity slows down; stocks indices and commodity prices fall, and sovereign risk rises. As such, shocks to the price factor could be interpreted as cost-push shocks or negative (global) supply-side shocks. Finally, growth shocks are mainly associated with increases in EMEs' GDP growth and mild (mostly positive) effects on the rest of their price and financial variables.

Table 3. Share of Variance Explained by Global Factor Shocks
(%, group medians)

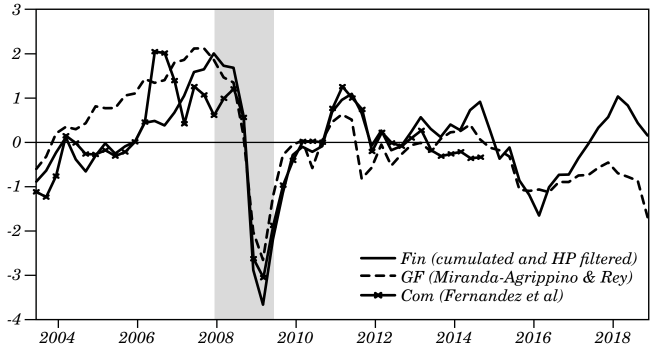
	Factor			
	Financial	*Price*	*Growth*	*Total*
All variables	23.3	12.9	1.0	40.8
EME variables				
GDP EMEs	14.0	15.2	7.2	38.5
CPI EMEs	3.9	5.0	0.3	9.2
EMBI	17.9	5.8	0.4	24.1
Stock market index	49.9	16.2	1.0	67.1
Global variables				
Import price index	28.2	17.4	1.6	43.5
GDP trade partners	22.2	13.0	3.4	39.1
CPI trade partners	24.8	16.9	1.7	43.4
Exchange rate (local currency/USD	36.3	11.8	0.7	48.8
Commodity prices	17.0	9.0	0.7	29.8
Crude oil	49.4	14.0	1.5	64.8
Copper	48.9	14.6	1.0	64.5
Aluminum	50.5	14.1	1.2	65.8

Source: Authors' calculations.
Notes: Baseline Model. Figures correspond to the share of the 20-period ahead forecast error variance that is attributable to each of the global factors shocks. For each column, group medians are reported (which implies that the sum of the columns does not necessarily add up to the total).

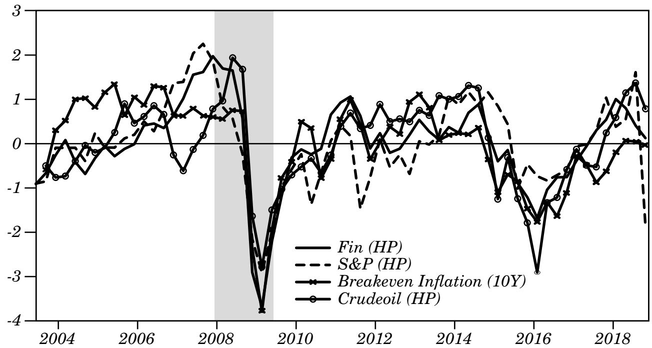
10. Initially, inflation decreases in EMEs as a consequence of a financial shock due to the appreciation of the local currency.

Figure 3: Comparing the 'Financial' Factor

(*a*) *Other factors in the literature*



(*b*) *Other variables*



Source: Authors' calculations.
Notes: Centered and scaled variables (s.d.=1). In both figures, the financial factor is the cyclical component (HP filter) of the cumulative estimated factor. (a) GF is the global financial factor estimated by MirandaAgrippino and Rey (2020); Com is the commodity global factor estimated by Fernández and others (2018). Shaded areas denote NBER U.S. recession dates. (b) U.S. Breakeven inflation (10Y) is expressed in percentage points, obtained from FRED. Cyclical component (HP filter) of the S&P 500 index and Brent oil price, originally obtained from Haver Analytics.

1.3.3 What is Behind the 'Financial' Factor?

Of the three factors, the financial factor has the most prominent role. As shown in table 3, the median share of the variance across all variables explained by it is over 23 percent. Moreover, as mentioned above, we allow it to affect all variables in a contemporary fashion. But this raises the question: why label it *financial*? While the idea of a global financial factor driving business cycles of EMEs seems easy to endorse in a context where such factor is identified by means of purely financial markets data,[11] calling our first factor a *financial* one may appear unwarranted *prima facie*. Part of the answer lies in figure 3 which shows the cyclical component of the cumulative financial factor accompanied by several other time series for comparison.

Figure 3a compares the financial factor to the global financial cycle in Miranda-Agrippino and Rey (2020), which they extract by using 858 asset price series. Similarly, figure 3b displays the cyclical component of the cumulative financial factor together with some of the main financial indicators—the cyclical component of the S&P index and the U.S. 10-Year breakeven inflation rate. We interpret the strong resemblance between our estimated financial factor and these other series as indicative of a financial nature of the factor.

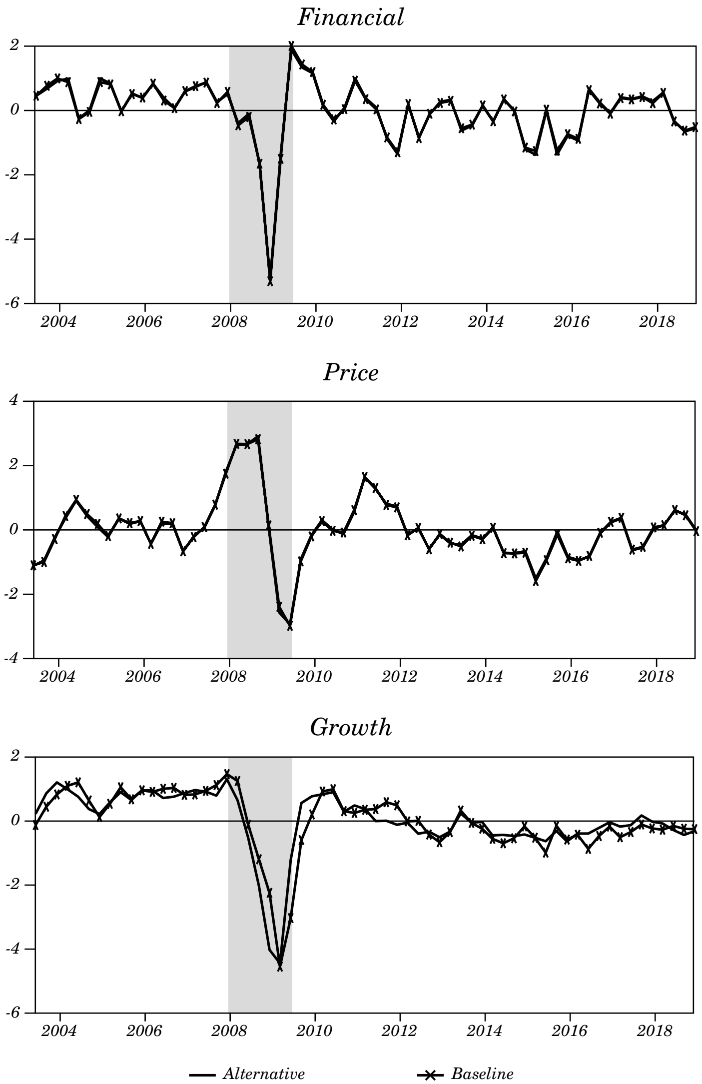
To further explore this idea, we analyze the effect of relaxing the assumption that the financial factor unloads on all the variables of the model. More specifically, we disallow a contemporaneous impact of the financial factor on GDP. This is consistent with a timing assumption often used when identifying financial shocks, whereby shocks in financial markets can affect real economic activity only with a lag.[12] In practice this is implemented by imposing a zero entry in the loading matrix of Equation (1) for all GDP variables, as table 4 describes.

Figure 4 presents the baseline factors and the new ones pinned down by using the alternative identification assumption. The immediate, noticeable remark is that the new financial factor is virtually indistinguishable from the baseline case. The same can be said for the price factor. In other words, the identification of our financial factor does not require the contemporary information provided by GDP: it is already captured by means of the financial variables and prices. A second, more subtle feature is the fact that now the growth factor is more similar to the financial factor. Actually, the correlation between the growth and financial factors increases from 0.33 in the baseline specification to 0.67 in this alternative specification.

Further results—collected in Appendix B—show that results in terms of forecast error variance decomposition qualitatively don't change much, except that we now observe a higher relevance for the growth factor at the expense of the new financial factor. This is not surprising since it is now the only common force inducing activity contemporaneously. Importantly, however, the alternative model has a poorer empirical fit vis-á-vis the baseline scenario related to an overall drop in the variance explained by all three factors, which further validates our baseline specification.

11. For example, Miranda-Agrippino and Rey (2020).
12. For example, Gilchrist and Zakrajsek, 2012.

Figure 4. Alternative Specification: Model without GDP-Financial Factor Channel
(comparison of estimated factors with those of the baseline model)



Source: Authors' calculations.
Notes: The figure shows the factors (in log-diff) estimated with the alternative model specification (with no direct channel between GDP variables and the financial factor) along with those estimated in the baseline model. All factors have been centered and scaled such that s.d. = 1. Shaded areas denote NBER U.S. recession dates.

Table 4. Alternative Specification: Model without GDP-Financial Factor Channel
(Restrictions on Loading Matrix)

	Factor		
	Financial	*Price*	*Growth*
EME variables			
GDP EMEs	○	○	●
CPI EMEs	●	●	○
EMBI	●	○	○
Stock market index	●	○	○
Global variables			
Import price index	●	●	○
GDP trade partners	○	○	●
CPI trade partners	●	●	○
Exchange rate	●	○	○
Commodities	●	●	○
Shadow FFR	●	○	○

Source: Authors' calculations.
Notes: White circles refer entries in the Λ matrix that are set to zero, whereas black circles correspond to unconstrained entries.

Moreover, figures 3a and 3b also display the similarity between the financial factor and the commodity factor of Fernández and others (2018)—which they extract from the cyclical component of country-specific commodity price indices that they construct—and the Brent crude oil price. This could be interpreted as evidence of the *financialization* hypothesis of commodity prices.[13]

Finally, a remark about the growth factor is warranted. It is, perhaps, surprising that the growth factor plays only a minor role in explaining the variance in the data. One possible explanation is

13. Some leading advocates of the financialization hypothesis include Jensen and others (2002), Tang and Xiong (2012), Adams and Glück (2015), and Basak and Pavlova (2016); while Hamilton and Wu (2015), and Chari and Christiano (2017) mark its dismissal.

that part of the commonality in the growth of the economies in our sample is already captured by the financial factor. This explanation is consistent with the results highlighted in the alternative specification above, where after disallowing a contemporaneous impact of the financial factor on GDP, the growth factor adapts by increasing its resemblance to the financial factor. This would suggest caution in the interpretation of the growth factor. Another explanation is that we may be over-restricting the contemporaneous impact of the growth factor and, hence, understating its relevance. However, the restrictions we impose are less severe than they may appear at first sight since they only refer to the contemporaneous impact of the factors on the variables. And, because the transition matrix is left unconstrained, each factor still affects every observable variable with a lag. Nevertheless, this explanation deserves further examination. An alternative approach we may pursue in the future is to impose *sign* restrictions instead of *zero* restrictions on the factor loadings, which could give the model additional flexibility in the identification of the factors while maintaining their structural interpretation.

2.Global Factors and Emerging Economies: Transmission Mechanisms

This section digs deeper into the channels through which global factors affect emerging market economies. To do this, we build on a large-scale DSGE model estimated for Chile—one of the EMEs considered in our pool of economies studied thus far—, augmenting it with a global factors block that comes from the estimated dynamic factor model presented in the previous section.

While the baseline factor model can be used to obtain a reduced form estimate of the global factors aggregate effect on some domestic EMEs' variables, it tells us little about the underlying mechanisms that ultimately determine the empirical results we observe. In contrast, the factor-augmented DSGE model allows us to disentangle the effects that the factors have on EMEs between the different channels that link the domestic and global blocks by taking advantage of the rich structure of the model. As a result, not only does the augmented model show the expected effect that shocks to the factors have on different domestic variables, but can also explain the transmission mechanisms that lead to those aggregate effects, through the lens of the structural model.

2.1 Baseline DSGE Model

The large-scale DSGE model estimated for the Chilean economy is based on García and others (2019). It is regularly used at the Central Bank of Chile for forecasting and policy analysis. The model considers a local economy and an external sector. The local economy interacts with the rest of the world in two dimensions: in the real sector by importing and exporting goods and services, and in the financial sector by trading bonds on international markets.

The following two subsections provide a brief narrative description of the core model's domestic and external blocks. A subsequent section presents how the model is augmented with the dynamic factor block. For further technical details of the DSGE model, readers are referred to García and others (2019).[14]

2.1.1 The Domestic Block

Four types of agents participate in the domestic economy: households, firms, the government, and a central bank. A fraction of households is composed of financially constrained hand-to-mouth agents. They consume private and public goods and services, supply labor to firms, pay taxes on consumption, labor income, and capital income, and receive lump-sum transfers from the government. The fraction of households that are not financially constrained can smooth consumption by saving and borrowing in local and foreign currency. They also invest in capital goods and receive dividends from firms they own (both locally and abroad). Households also face involuntary unemployment spells due to a labor market with search and matching frictions as in Mortensen and Pissarides (1994), which also features endogenous separations and wage rigidities.

Different types of firms are in charge of production. In the non-commodity sector, firms producing domestic goods utilize capital, labor, and oil as inputs, with pricing decisions subject to Calvo-type nominal rigidities. Another set of firms sell differentiated imported goods on the domestic market and are also subject to nominal rigidities. Domestic and imported goods are then combined to form a homogeneous intermediate good used for final consumption or investment goods. The assumption of rigid prices in local currency leads to an incomplete

14. For a description of the DSGE model and how it is regularly used for policy analysis see Central Bank of Chile (2020)

exchange rate passthrough, in line with empirical evidence. Profits generated by firms are delivered in the form of dividends to their owners (unconstrained households).

Finally, the commodity sector is modeled as a representative, capital-intensive exporting firm, with shared ownership between the government and foreign agents.

The government follows a structural balance fiscal rule where the desired spending of each period is defined not by current but by structural or long-term revenues, mimicking the Chilean legislation on fiscal spending. The effective spending path may eventually differ from the rule due to exogenous shocks. Expenditures are split between government consumption, investment in public goods, and transfers to households. These are financed with tax revenues, income from property in the mining sector, and debt issuance. In addition, the government has a program in place to smooth out after-tax gas price volatility, which involves a variable combination of taxes and subsidies for gas consumption.

The central bank conducts monetary policy based on a Taylor-type policy rule. Under this rule, the interest rate responds to deviations of inflation from the 3 percent target and of output growth from long-term growth. When evaluating inflationary pressures, the central bank responds to a weighted average of current and expected inflation, which consider both core and headline measures. Additional exogenous disturbances allow for the effective rate to deviate from what the systematic part of the rule prescribes.

2.1.2 Foreign Block and Linkages with the Domestic Economy

In the foreign block, prices of commodities (copper and oil) and other imported goods (excluding oil) are modeled as exogenous, together with the trading partners' growth and inflation, and a risk-free external rate. The exchange rate is determined through an arbitrage relationship between local and foreign currency interest rates, while the net foreign asset position, as a percentage of GDP, determines the country risk as in Schmitt-Grohé and Uribe (2003). Both the exchange rate and the risk premium dynamics also allow for additional nonsystematic exogenous disturbances.

Below we describe how each variable from the external block is linked with the domestic economy and how movements in those variables affect domestic variables.

• Commodity export prices: A representative firm produces a commodity that is fully exported at an exogenously determined foreign-currency-denominated price. The firm's ownership is shared between the government and foreign investors. Cash flows are shared accordingly, but the government also levies taxes on the foreign investors' profit share. As in Fornero and Kirchner (2018), production uses sector-specific capital, subject to adjustment costs and time-to-build frictions in investment. The labor share of the sector is assumed to be negligible.

A shock to the price of the exportable commodity, by increasing government income, reduces the fiscal financial burden, allowing for an expansion of the spending budget. The shock also triggers an expansion of the sector's investment that, due to the time-to-build technology, is only relevant if the shock is persistent enough to offset the investment lag. Additionally, the currency appreciation that follows the rise of the commodity price reduces marginal costs through cheaper imports. Overall, the shock is both expansionary and deflationary.

• Commodity import prices: Commodity imports, modeled as oil imports, are both directly a part of the final consumption basket and part of the production function of domestic wholesale goods, alongside labor and capital.

A shock in commodity import prices directly affects inflation through higher prices in the gas and energy components of the CPI. However, the impact is partially dampened by a fiscally financed smoothing policy for gas prices that, on the other hand, puts pressure on the fiscal budget. Higher oil prices also affect core CPI (excluding energy and food) through two channels. First, through indexation of non-oil-related prices to past headline inflation. Second, as oil is also an input in the production function of general goods, a higher price raises marginal costs and inflation. The shock is associated with only a modest interest-rate response explained mainly by two reasons. On one hand, monetary policy responds only partially to noncore CPI and short-term inflation. On the other hand, as a higher cost of intermediate imported goods can be understood as a negative supply shock, the pressure to raise rates due to higher inflation is partially dampened by a desire to compensate for the lower output.

• Other import prices: Non-commodity imports are used as an input for the production of final goods, in combination with domestically produced intermediate goods. Thus, a shock to import prices directly raises marginal costs, thus leading to higher inflation and lower output.

• Commercial partners' inflation rate: Higher inflation of commercial partners, all else equal, will make the exportable good more competitive, thus fostering exports. In addition, higher foreign prices, while keeping nominal import prices constant, reduce real import prices ($\frac{P^M}{P}\downarrow = \frac{P^{M*}}{P^*\uparrow} rer$). While the shock does cause a real depreciation, it is not enough to offset the drop in the foreign-currency real import prices, which leads to lower real marginal costs and lower inflation.[15]

• Commercial partners' growth rate: In the model, the demand for non-commodity exports is directly linked with the size of the foreign economy. If commercial partners' GDP is expanding, they will demand more of the local economy exports, thus stimulating domestic GDP. Higher demand will also lead, everything else equal, to more inflation and higher monetary policy rates.

• Foreign financing costs: The relevant interest rate for the decision of holding and acquiring new foreign-currency debt includes both a risk-free rate (proxied by the federal funds rate) and a risk premium. While in the model the former is entirely exogenous and the latter has both exogenous and endogenous components, a shock to either will have the same effect of increasing the financing cost in foreign currency. Thus, alongside an exchange rate depreciation, inflation will rise and output will drop.

2.2 The Factor-Augmented Model

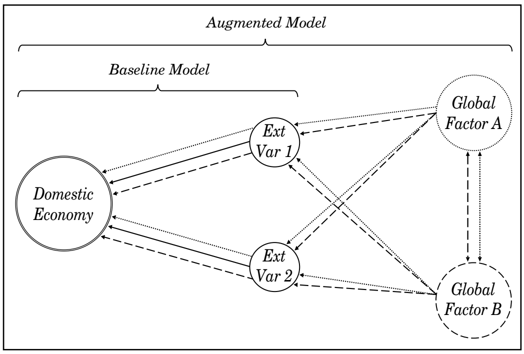
In order to analyze the domestic implications of a shock to the global factors, we augment the baseline DSGE model's external block by allowing for the factors to influence the different variables in the model's external block. To do so, we follow the same structure for the factors as described in the previous section. We only modify the external block; the rest of the model is kept as in the baseline DSGE from García and others (2019). We take the estimated factors F_t and state-transition coefficient matrix Φ from the baseline factor model and re-estimate the matrix of factor loadings Λ and the variance matrix H, allowing for autocorrelation on the exogenous disturbances and

15. The partial adjustment of the exchange rate might be due to the presence of nominal rigidities that inhibit full price adjustments.

keeping the same identification restrictions from table 1.[16] Finally, in order to ensure uniqueness in the steady state, we add, when needed, a small error correction parameter to the dynamic equations.

Figure 5 schematically summarizes the differences between baseline and augmented models. In the former, the model only considers the direct effect of the variables (the solid arrows in the figure). Furthermore, external variables are also assumed to be orthogonal as they are only affected by their own shocks. In contrast, the augmented model allows for indirect effects of the global factors on the domestic economy through their influence on the dynamics of the external variables (the figure's dotted and dashed arrows). In the augmented model, the orthogonality among external variables breaks down, as the systematic effect that the factors have on those variables induces correlation among them.

Figure 5. External Block Structure in the Baseline and Factor-Augmented DSGE Model



Source: Authors' calculations.
Notes: In the diagram, the arrows show the transmission mechanism of a shock originating from a source depicted with the same type of line.

16. We re-estimate the Λ and H matrices due to small differences between the observables in the factor model and the DSGE model. The differences range from the sample size to variables definitions. For the DSGE model, the sample is restricted by the date the Central Bank of Chile started using nominal instead of real rates as the policy instrument, while for the factor model we make use of the longer data availability. Additionally, in order to maintain consistency among countries and as described in section 1.1, for the factor model we construct each country's commercial partners price index by using the top-ten commercial partners. For the DSGE model, we use the official series reported by the statistical department of the Central Bank of Chile, which consider a broader coverage.

In short, the setup provided by the augmented model allows us to combine the comovement in global forces pinned down by the dynamic factor model with the rich propagation mechanisms embedded in the DSGE model. We explore next how shocks to global forces affect domestic variables through the lens of this setup.

2.3 Domestic Implications of Global Factor Shocks

This section describes the model-implied effects that shocks to the factors have on Chile by using the augmented model. We analyze the aggregate impacts while also differentiating between alternative transmission channels. We also emphasize how, for some shocks, different channels reinforce one another, which leads to larger aggregate effects, while for others, the final impact may be dampened due to offsetting effects.

2.3.1 Aggregate and Disaggregate Effects

The augmented DSGE model can be used to predict the expected aggregate effect that a factor shock has on any given variable of interest. By selectively *turning off* different channels, we can further distinguish between the parts of the aggregate effects that are associated with a particular mechanism.

For example, we can ask the model what would the impact of a shock to the financial factor on domestic output be, and call that the *aggregate effect* of the financial factor on GDP. Additionally, by taking advantage of the structural nature of the model, we can further ask what would the impact of a shock to the financial factor on domestic output be in a counterfactual world where all variables from the external block but the oil price remained constant. We would then call the answer to that question the effect of the financial factor on GDP *due to movements in oil prices*.

More formally, lets summarize the augmented model by the following set of equations:

$$E_t\,(D_{t+1}) = D^i\,(Y^t, Z^t) \tag{3}$$

$$Y_t = \Lambda F_t + u_t \tag{4}$$

$$F_t = \Phi F_{t-1} + w_t. \tag{5}$$

The vectors D_t, Y_t, and F_t represent, respectively, the variables from the domestic block, the foreign block, and the global factors at time t. The elements of the factor vector F affect each other with the structure given by Φ and unload on the global variables contained in the vector Y through the loading matrix Λ. The vectors Y^t and Z^t denote all the information available at time t about the past and expected trajectories of external variables Y and other relevant variables Z, and $D^i(Y^t, Z^t)$ denote the policy functions for the expected value of D_{t+i} given the set of information contained in Y^t and Z^t.

We define $Y^j_{t+i} = E_t(Y_t | \in^j_t)$ as the expected response of the vector Y at period $t+i$ given a shock to the factor j at time t. For each of the global variables included in vector Y_t, $Y^{j,k}_{t+i}$ is a vector equal to Y^j_{t+i} with all its elements equal to zero except the one in position k, such that $Y^j_{t+i} = \sum_{k=1}^{N_Y} Y^j_{t+k}$.

We also define $E_t(D_{t+i} | Y_t = Y^{t,j,k}) = D^{i,j,k} (Y^{t,j,k} = Z^t)$ where $Y^{t,j,k}$ denote all information available at time t about the past and expected trajectories of the variable $Y^{j,k}_t$. The policy function $D^{i,j,k}$ $(Y^{t,j,k}, Z^t)$ is then the expected value for D_{t+i}, given shocks to factor j in a counterfactual world where all the external variables, except for the one in position k remain constant. Then, computing $D^{i,j,k}$ $(Y^{t,j,k}, Z^t)$ for every k allows us to decompose the expected response at time $t+i$ of a shock to factor j, through each channel k, of any variable of interest contained in D. In other words, we will be able to decompose the effect that a shock to a factor has in a domestic variable between the shares that can be attributed to each global variable that link the model's domestic and external blocks.

2.3.2 Dynamic Shock Effects

In section 1.3.1 we described how shocks to the global factors affect different global variables. To summarize, the financial shock tends to raise commodity and import prices, as well as commercial partners' inflation rates and GDP growth while easing financial conditions for the EMEs. Shocks to the growth factor induce similar effects, although the responses are more muted and take longer to reach their peaks. On the other hand, shocks to the price factor are associated with increased import prices, a drop in commodity prices, commercial partners' inflation rates and GDP growth, and worsened financial conditions.

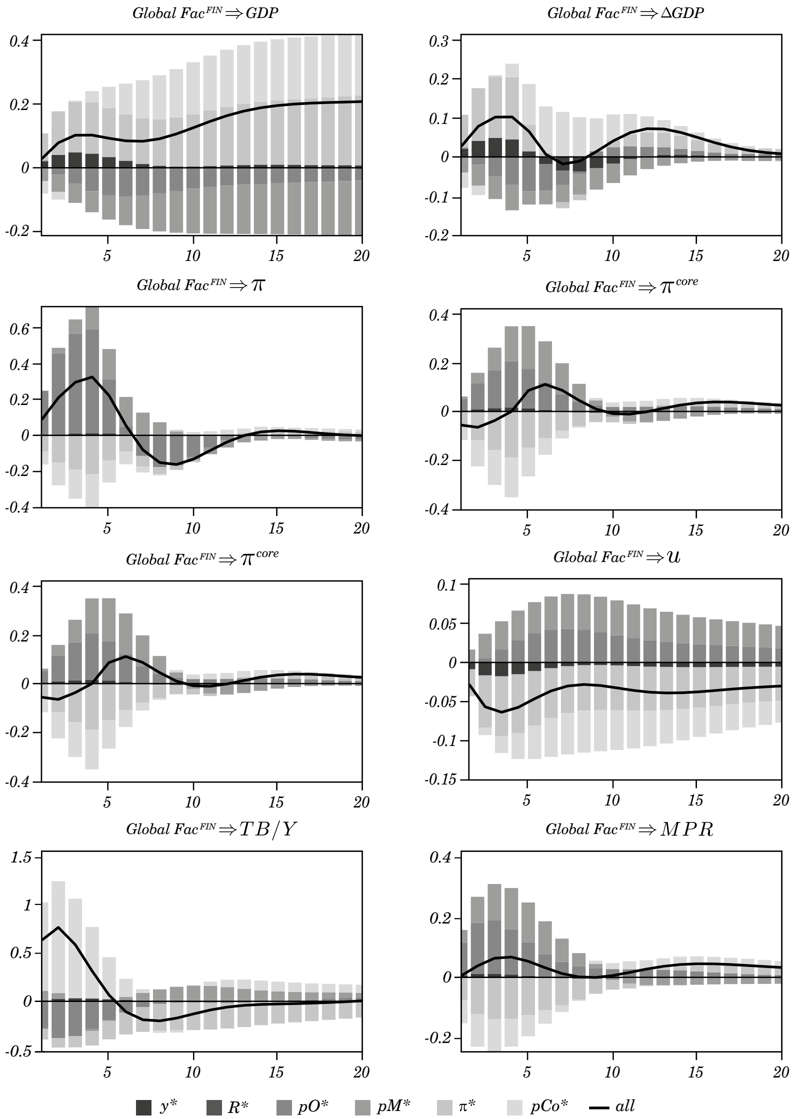
In this section, we use the factor-augmented DSGE model to analyze how the previously described effects end up affecting EMEs' domestic variables. We use the methodology described in section 2.3.1 to decompose the responses in the different channels through which the foreign and domestic blocks are linked. The domestic responses to a financial factor shock are summarized in figure 6.

Regarding commodities, higher export and import prices following a financial shock have opposite effects. On the one hand, a higher price of the exportable commodity price pCo^*, which for Chile corresponds to copper, increases the country's income and the trade balance, and appreciates the exchange rate, thus inducing lower inflation and monetary-policy rates. The higher commodity price also fosters output through incentives to increase the specific investment of the sector (not reported in the figure). On the other hand, a higher commodity import price pO^*, namely oil for Chile, tends to have the opposite effect. Since oil is an input in the production function, an increase in its price acts as a negative supply shock by raising marginal costs and contracting the economy. The higher price also deteriorates the trade balance. Inflation raises through two channels, first through the direct impact on the gas and energy components of headline CPI and, second, through its impact on core inflation (excluding energy and food), by the previously described higher marginal costs and by the indexation of core goods to headline inflation.

Compared to the effects of higher commodity import prices, higher non-oil import prices pM^* have similar implications, though less pronounced, on headline inflation, as it does not affect the noncore basket as much, and more intensive in core inflation, where it affects marginal costs through pricier imported inputs. Higher inflation of commercial partners π^* tends to increase the competitiveness of the domestic economy by fostering exports. Assuming nominal import prices constant, higher foreign inflation makes real import prices drop, and then also marginal costs and inflation. The shock to the financial factor also increases foreign GDP growth y^*, demand for exports, and then domestic GDP. The financial factor also reduces the foreign financing costs, summarized in the model by R^*. This channel, however, shows negligible effects due the estimation sample covering a period where the country's risk premium was low and stable.

Summing up, after a shock to the financial factor, the commodity export price and foreign inflation channels lead to increased output and lower inflation. In contrast, the import price channels in commodity and non-commodity sectors have the opposite effect, leading to lower output and higher inflation. The first set of channels dominates regarding GDP growth, leading to higher output, while the second set of channels dominates in terms of higher overall inflation. Finally, the foreign growth channel positively affects both GDP and inflation, although the effect on the latter is negligible.

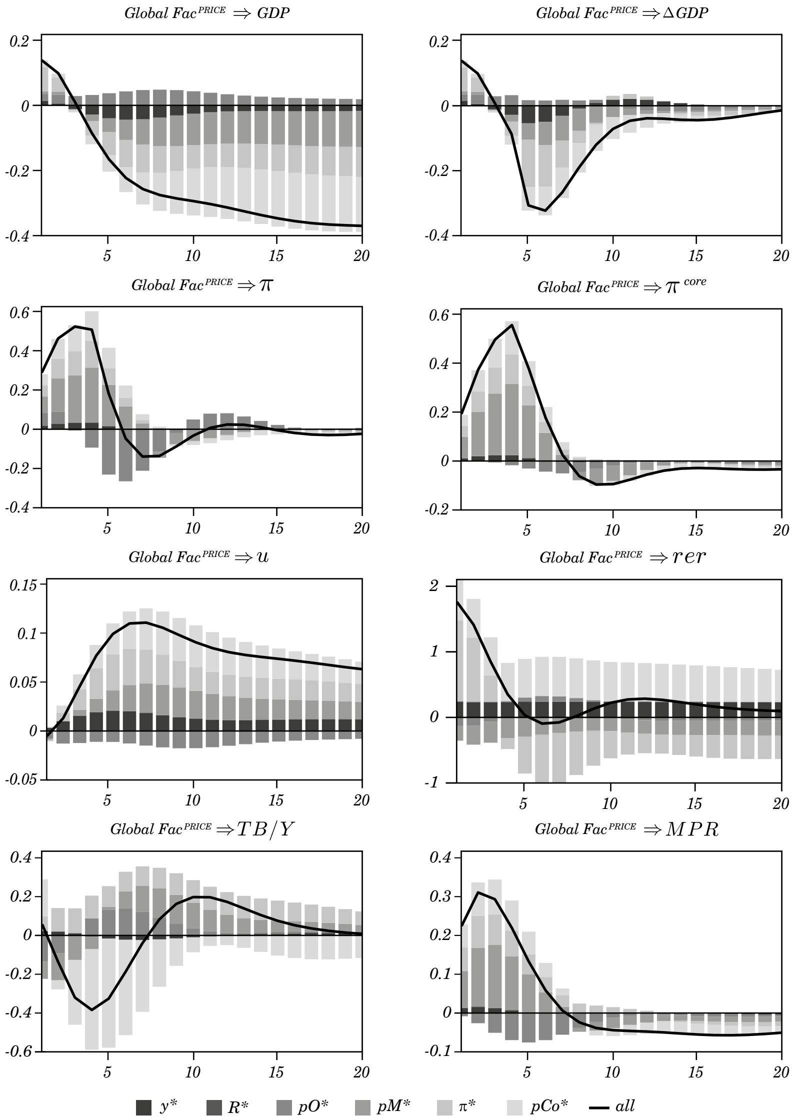
Figure 6. Domestic Effects of a Shock to the Financial Global Factor



Source: Authors' calculations.

Notes: (1) The bars show the response of each variable to one standard deviation shock to the financial global factor shock while keeping only one channel open at the time. (2) The black line is the response of each variable to the shock when all channels are open. It is, by construction, equal to the sum of the bars. (3) GDP refers to the deviation of the level of GDP from the long-run productivity growth path, ΔGDP denote GDP annual growth, π and πcore denote respectively annual headline and core inflation (where food and energy items are removed), u is unemployment, rer is the real exchange rate, TB/Y is the trade balance as a fraction of GDP, and MPR refers to the annualized monetary-policy rate.

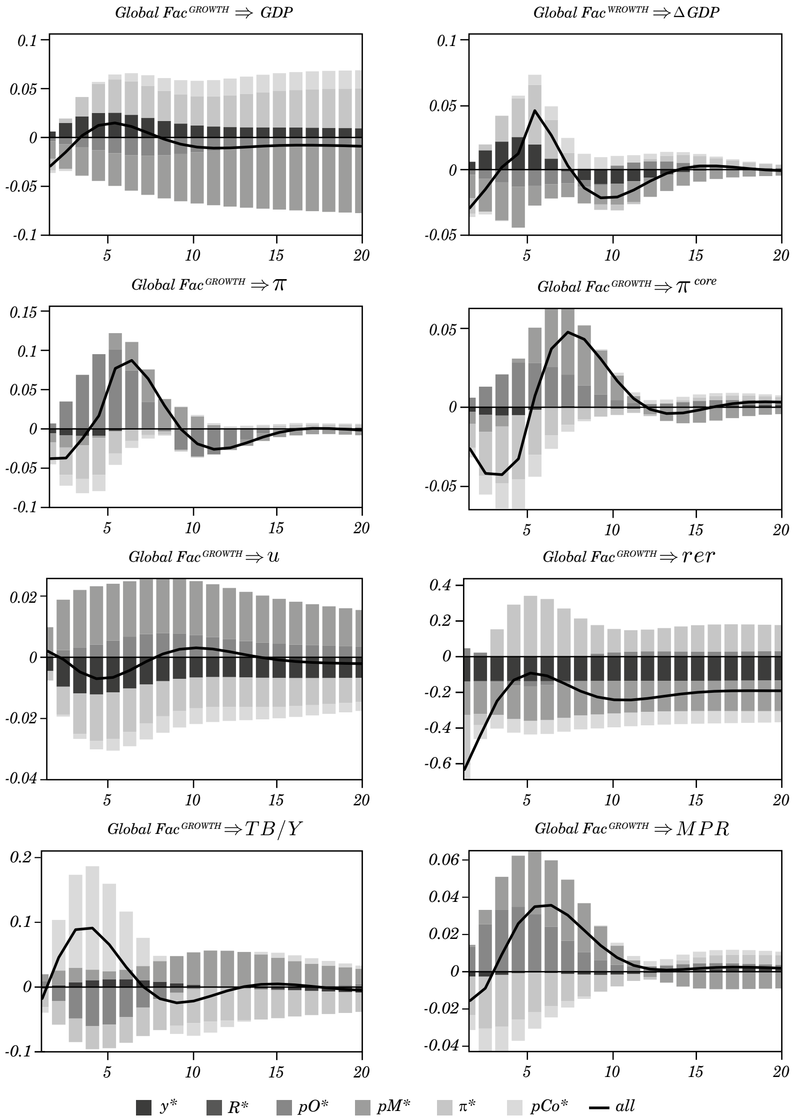
Figure 7. Domestic Effects of a Shock to the Price Global Factor



Source: Authors' calculations.

Notes: (1) The bars show the response of each variable to one standard deviation shock to the price global factor shock while keeping only one channel open at the time. (2) The black line is the response of each variable to the shock when all channels are open. It is, by construction, equal to the sum of the bars. (3) GDP refers to the deviation of the level of GDP from the long-run productivity growth path, ΔGDP denote GDP annual growth, π and πcore denote respectively annual headline and core inflation (where food and energy items are removed), u is unemployment, rer is the real exchange rate, TB/Y is the trade balance as a fraction of GDP, and MPR refers to the annualized monetary-policy rate.

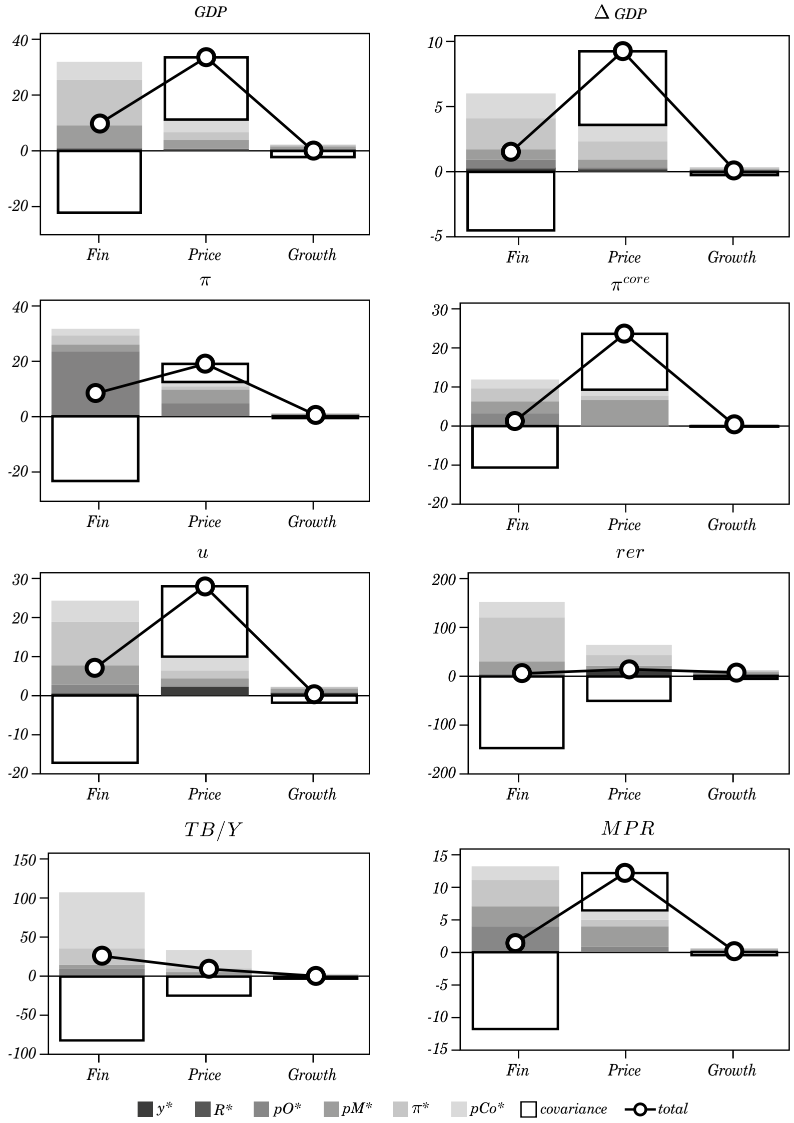
Figure 8. Domestic Effects of a Shock to the Growth Global Factor



Source: Authors' calculations.

Notes: (1) The bars show the response of each variable to one standard deviation shock to the growth global factor shock while keeping only one channel open at the time. (2) The black line is the response of each variable to the shock when all channels are open. It is, by construction, equal to the sum of the bars. (3) GDP refers to the deviation of the level of GDP from the long-run productivity growth path, ΔGDP denote GDP annual growth, π and πcore denote respectively annual headline and core inflation (where food and energy items are removed), u is unemployment, rer is the real exchange rate, TB/Y is the trade balance as a fraction of GDP, and MPR refers to the annualized monetary-policy rate.

Figure 9. Share of Variance Explained by Global Factors per Channel



Source: Authors' calculations.

Notes: (1) The bars correspond to the unconditional share of the forecast error variance that is attributable to each of the global factor's shocks. (2) Each color represents the variance explained by a factor while keeping only the corresponding channel open. The gray bars correspond to the share attributable to covariances, computed as the difference between the explained variance when all channels are open and the sum of the explained variance keeping one channel open at a time. (3) GDP refers to the deviation of the level of GDP from the long-run productivity growth path, ΔGDP denote GDP annual growth, π and πcore denote respectively annual headline and core inflation (where food and energy items are removed), u is unemployment, rer is the real exchange rate, TB/Y is the trade balance as a fraction of GDP, and MPR refers to the annualized monetary-policy rate.

Figure 7 describes the effects of a shock on the global price factor. A key finding of this exercise is that, in contrast to the previous case, all channels point in the same direction, with the exception of the imported commodity . Lower commodity export prices lower exports (and output) and raise inflation. In this case, the currency depreciation channel dominates the deflationary pressures due to lower exports. Lower foreign inflation raises real import prices and marginal costs, acting as a negative supply shock that lowers output and raises inflation. Foreign demand also drops, with a subsequent effect of lower output. The only channel that goes against these drivers is the commodity import price. As with the exported commodity, the factor shock lowers the import price, leading to lower marginal costs, higher output, and less inflation. The deflationary impact is compounded by an additional direct effect in the final consumer basket due to the gas and energy component.

Finally, consistent with the similar effect that shocks to the financial and growth factors have on most foreign variables, figure 8 shows how the domestic effects of a shock to the latter are qualitatively comparable to the responses after a shock to the former, albeit in a smaller scale.

To summarize, shocks to the financial factor lead to higher output and inflation. On the other hand, shocks to the price factors are followed by lower output and higher inflation. Shocks to the growth factor have similar effects to those of the financial factor, although smaller in magnitude. As the aggregate effect on inflation is much more pronounced following shocks to the price factor, so are the associated movements of the monetary-policy rate.

2.3.3 Variance Decomposition and the Role of Covariances

In section 1.3.2 we showed, by using the baseline factor model, that the financial factor has a dominant role in explaining the variance of most global variables as compared with the other global factors. However, as shown in table 3, this fact does not translate into the financial factor explaining an equivalently significant share of EME's GDP and inflation variances, where the price factor has a comparable role.

Table 5. Variance Decomposition in the DSGE Model: the Role of Covariances

	GDP	*ΔGDP*	π	*πcore*	μ	*rer*	*TB/Y*	*MPR*	*Median*
Financial global factor									
Sum of FEVD by channel	32	6	32	12	24	152	108	13	32
Role of covariances	-22	-4	-23	-11	-17	-146	-82	-12	-22
Total explained variances	10	2	8	1	7	6	26	1	7
Price global factor									
Sum of FEVD by channel	11	4	12	9	10	65	34	6	11
Role of covariances	22	6	7	14	18	-50	-24	6	6
Total explained variances	34	9	19	24	28	15	10	12	19
Growth global factor									
Sum of FEVD by channel	2.3	0.4	1.1	0.5	2.2	12.9	3.5	0.6	2.2
Role of covariances	-2.2	-0.2	-0.6	-0.1	-1.8	-4.6	-2.7	-0.4	-1.6
Total explained variances	0.1	0.1	0.5	0.3	0.4	8.3	0.8	0.2	0.4

Source: Authors' calculations.

Notes: (1) The percentages correspond to the unconditional share of the forecast error variance that is attributable to each of the global factor's shocks. (2) Sum of FEVD by channels is computed as the sum of the variance explained by a factor while keeping only one channel open at a time. Total explained variance is the explained variance by the factor when all channels are open. The role of covariances is computed as the difference between both. (3) GDP refers to the deviation of the level of GDP from the long-run productivity growth path, ΔGDP denote GDP annual growth, π and πcore denote respectively annual headline and core inflation (where food and energy items are removed), u is unemployment, *rer* is the real exchange rate, TB/Y is the trade balance as a fraction of GDP, and MPR refers to the annualized monetary-policy rate.

The use of the factor-augmented DSGE model allows us to shed more light onto those results. By decomposing the factor effects by channels, we see that the greater importance of the financial factor in explaining the external variables' variance directly maps to an equivalent role, channel by channel, in explaining the domestic variables' variance. As can be seen by comparing the size of the shaded bars from figure 9, if we consider only the direct effect of the different channels, shocks to the financial factor explain the most variance, followed by shocks to the global price factor, and lastly, shocks to the growth one.

As with the baseline factor model, the analysis carried on with the DSGE also shows that for domestic variables, relative to their role in explaining global variables, the relative importance of the financial factor is dampened, while the impact of the price factor expands. To understand the discrepancy, it is worth paying particular attention to the role of covariances. As we can see by comparing figures 6 and 7, after a shock to the financial factor, different channels push the domestic variables in different directions, dampening the aggregate effect. The opposite happens after a shock to the price factor, where most channels tend to push the domestic variables in the same direction.

The share of domestic variables' variance attributed to a global factor shock can significantly differ depending on whether the shock pushes global variables in similar or opposing directions. Figure 9 and table 5 show the role that the comovements between the different transmission channels have on the aggregate explained variance. On the one hand, the financial factor shows the most significant channel-by-channel effect. However, as their effects tend to cancel each other, the aggregate explained variance is reduced due to this negative covariance effect. On the other hand, for the price factor, while different channels have an individually smaller impact, they tend to always go in the same direction, which leads to an exacerbated effect on the explained variance. This suggests that it is not enough to analyze separately how the factors explain the variance of global variables, given that the extent to which those responses comove can be equally or more important. In this example, while EMEs appear to be relatively well-hedged to deal with shocks to the financial factor, when it comes to shocks to the price factor, *when it rains, it pours*: when one channel affects the economy negatively, they all do.

We showed that the DSGE model manages to capture and explain both the relative dampening and the relative amplification of, respectively, the importance of the financial and price factors in explaining the dynamics of domestic variables. However, the dampening on the financial factor appears to be much more pronounced in the DSGE model than in its empirical counterpart. How can we account for this discrepancy? If the DSGE model were an accurate and comprehensive representation of the true data generating model, it would be expected for the factors to have a similar role in the DSGE and in the reduced form empirical model. However, by comparing tables 3 and 5, it is clear that the relative role of the financial factor in the DSGE model is much smaller. Understanding some key differences between both modeling approaches can be helpful to comprehend the root cause of the disparity regarding the assigned role of the financial factor. The empirical model attempts to maximize, in a reduced form, the covariance between the explanatory variables (the factors) and the dependent variables (from the domestic economy). It is then a helpful tool to get a good answer for the question of "how much" of a role global factors play, at the expense of being silent on 'how' shocks propagate. The structural nature of the DSGE model, on the other hand, is better equipped to answer the question of 'how' factor shocks are transmitted. However, the answer to the "how much" question will only be as accurate as how the different channels through which the factors affect the domestic economy are explicitly modeled. The estimation of the relative importance of a factor could be biased if a relevant transmission mechanism is missing in the model, more so if this missing channel disproportionately affects the transmission of one particular factor. Given that the DSGE model is, with the exception of an endogenous risk premium channel *à la* Schmitt-Grohé and Uribe (2003), absent of financial frictions, it would be reasonable to hypothesize that the model may be underrepresenting the true importance of the financial factor. Adding a global factors block to a model that incorporates financial frictions but is otherwise a similarly featured large-scale DSGE model as the one used in this paper[17] could provide a good test for this hypothesis and is a promising avenue for future research.

17. A suitable model could be the one described in Calani *and others* (2022). The model, also estimated for the Chilean economy, builds on the framework from García *and others* (2019) by introducing, similar to Clerc *and others* (2014), three layers of financial frictions, allowing for households, entrepreneurs, and banks to default on their financial obligations.

3. Conclusions

This paper has analyzed the role of global drivers on the business cycles of EMEs. The distinguishing feature of the analysis lied on the careful identification of multiple external forces by means of a constrained dynamic factor model. In accordance with prominent previous research, we have found empirical support for the overall relevance of a global financial factor—which explained more than a third of GDP fluctuations—followed by external factors akin to price and growth/productivity shocks.

In order to better understand the transmission mechanisms underlying the aggregate effects of shocks to our estimated factors, we focused on Chile—one of the EMEs in our sample—and embedded the previous empirical factor structure as an additional tier of the DSGE model of the Central Bank of Chile, whereby its original foreign variables now hinged on a set of foreign factors. In an apparent puzzling result at first, the aggregate empirical dominance of the financial factor compared to the price factor became the other way around, so we inspected the mechanism and found that, while a shock to the global financial factor triggered movements in global variables that steered domestic variables in opposing directions, after a global price shock, in contrast, such offsetting effects in domestic variables were no longer present. These results enriched our understanding of the consequences for monetary policy, which now should react more strongly in the face of price-factor shocks.

Finally, while we subjected our factor model to many robustness tests, we left aside some relevant issues possibly worth exploring in future work, such as the relation of our financial factor with relevant statistics, for instance, the U.S. break-even inflation. Another relevant research avenue should be the DSGE estimation for different economies, so we could eventually appraise the generality of the *inverse* effects of the financial and price-factor shocks at the local level.

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Appendices

Appendix A
Number of Common Factors

The number of factors that we considered in our baseline model was mostly guided by the set of different drivers underlined by previous research. In this brief digression we tackle the issue of the number of factors from a purely statistical sense in which we specifically apply the cornerstone contributions of Bai and Ng (2002, 2007); Amengual and Watson (2007), and Ahn and Horenstein (2013) to our dataset. The common thread across this literature is the specification of either a dynamic or static approximate factor model that is consequently estimated by principal components. With such estimation results at hand, these papers formulate some penalty criteria that ultimately provides the true number of factors asymptotically. Now, in our case however, since we are posing a state-space model with loading matrix constraints estimated by maximum likelihood, we cannot directly apply the results of the aforementioned tests for our specific formulation, although we can still use such optimal, dynamic factor model results if we actually fit that very same model to our data, and therefore take the optimal number of tests as a guidance for the specification we actually pursue in the paper.

Table A1 shows the number of factors for the aforementioned tests. The main pattern that emerges is the following: from the vantage point of the relatively more short-sample focus of Ahn and Horenstein (2013), we get a single dynamic factor inducing cycles into the features of the emerging economies we considered, while on the contrary, the asymptotic test of Bai and Ng (2002) leads to three factors. In any case, we get a sort of consistency between the number of factors that we include by entirely looking at the literature and those supported by statistical criteria.

Table A1. Statistical Number of Factors

	Statistical Test		
Max. number of factors	*BN*	*AH*	*AW*
2	2	1	1
4	3	1	1
6	3	1	1

Source: Authors' calculations.
Notes: Max. Number of Factors corresponds to the maximal number of factors considered in the corresponding principal components estimation. BN: Bai and Ng (2002), ICp2 information criterion; AH: Ahn and Horenstein (2013), eigenvalue ratio criterion; AW: Amengual and Watson (2007) estimate of dynamic factors given BN.

Appendix B
Model without GDP-Financial Factor Channel

Here we present additional results of the alternative model specification without a GDP financial factor channel, described in section 1.3.3. Tables B1 and B2 show, respectively, the factors' and the variables' forecast error variance decomposition at the 20-quarter horizon. Not surprisingly, by comparing those numbers with respect to the baseline scenario of table 2 in the text, we observe a higher relevance for the growth factor: since it is now the only common force inducing activity contemporaneously, it roughly doubles the variance explained across the set of factors considered.

Table B1. Factors' Variance Decomposition
(model without GDP-Financial factor channel (%))

	Shocks		
	Financial	*Price*	*Growth*
Financial Factor	71.6	24.8	3.6
Price Factor	31.7	60.4	7.9
Growth Factor	16.8	41.7	41.5
Average	40.0	42.3	17.7

Source: Authors' calculations.
Notes: Alternative model specification, with no direct channel between GDP variables and the financial factor. Figures correspond to the share of the 20-period ahead forecast error variance that is attributable to each of the global factors' shocks.

Table B2. Share of Variance Explained by Global Factor Shocks
(model without GDP-Financial factor channel (%, group medians))

	Factor			
	Financial	*Price*	*Growth*	*Total*
All variables				
	13.1	13.7	2.4	38.8
EME variables	6.0	15.0	14.9	35.9
GDP EMEs	2.9	5.2	0.6	8.7
CPI EMEs	16.0	5.6	0.8	22.4
EMBI	46.8	16.2	2.3	65.4
Stock market index				
Global variables				
Import price index	26.6	25.0	4.9	58.9
GDP trade partners	5.9	14.6	14.5	35.0
CPI trade partners	19.7	17.8	3.4	40.9
Exchange rate (local currency/USD	34.7	12.0	1.7	48.4
Commodity prices	13.9	9.4	1.7	28.1
Crude oil	46.2	14.6	3.3	64.0
Copper	47.0	15.3	2.4	64.7
Aluminum	47.7	14.5	2.7	65.0

Source: Authors' calculations.
Notes: Alternative model specification, with no direct channel between GDP variables and the financial factor. See notes in Table B1.

Appendix C
More Robustness Checks

C.1 Blending Growth and Price Factors

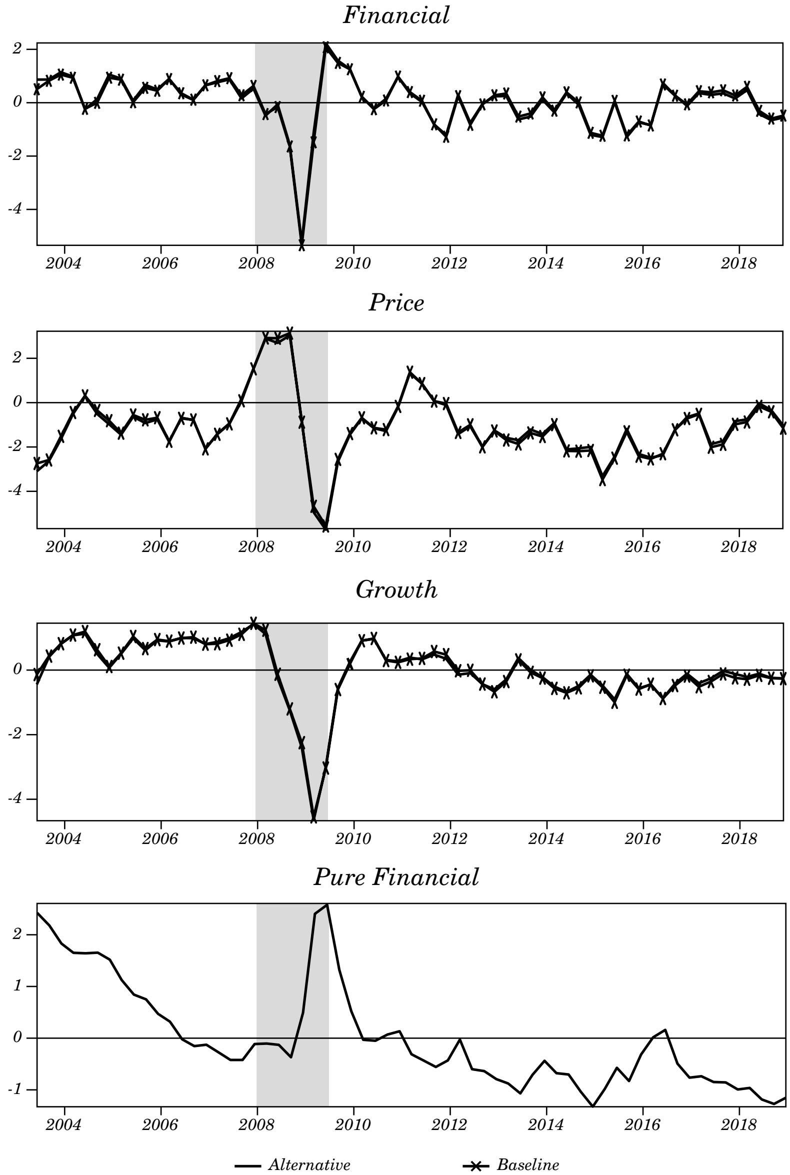
Here given the scant relevance of the growth factor in the baseline scenario, we explore the possibility of blending such factor with the price factor, as we show in table C1. What we get is a decrease of roughly three percentage points for the total median variance explained by the aggregation of factors. On the other hand, the explanatory power of the financial factor increases substantially, while the combined factor sees its variance explained eroded by seven points on average. These results therefore suggest that the separation of the growth and price factors catalyze a better identification and transmission of shocks.

Table C1. Alternative Specification: Joint Growth and Price Factor (Two-Factor Model)
(restrictions on loading matrix)

	Factor	
	Financial	*Price*
EME variables		
GDP EMEs	●	●
CPI EMEs	●	●
EMBI	●	○
Stock market index	●	○
Global variables		
Import price index	●	●
GDP trade partners	●	●
CPI trade partners	●	●
Exchange rate	●	○
Commodities	●	●
Shadow FFR	●	○

Source: Authors' calculations.
Notes: White circles refer entries in the Λ matrix that are set to zero, whereas black circles correspond to unconstrained entries.

Figure C1. Alternative Specification: Additional 'Financial' Factor
(comparison of Estimated Factors with those of the Baseline Model)



Source: Authors' calculations.
Notes: All factors have been centered and scaled such that s.d. = 1.

C.2 An Additional "Purely Financial" Factor (Four-factor model)

Given the fact that the financial factor in the baseline model does not strictly discern a global interpretation from a strictly financial one, we take a look at the scenario in which we disentangle such global factor from a strict common force that only affects financial market variables. Table C2 shows the identifying details.

In figure C1 we plot the consequent time series of the factors we got. The salient feature is the overall stability of the new factors with respect to the baseline results. In terms of quantitative results, even though the aggregate variance explained increases, the actual combined variance explained by the global, price, and growth factors remains intact, which broadly suggests that the global/financial factor in the baseline scenario actually captures common forces across all the variables of the model, while exclusively financial movements appear to have less relevance.

Table C2. Alternative Specification: Additional "Purely Financial" Factor
(restrictions on loading matrix)

	Factor			
	Financial	*Pure Financial*	*Price*	*Growth*
EME variables				
GDP EMEs	●	○	○	●
CPI EMEs	●	○	●	●
EMBI	●	●	○	○
Stock market index	●	●	○	○
Global variables				
Import price index	●	○	●	○
GDP trade partners	●	○	○	●
CPI trade partners	●	○	●	○
Exchange rate	●	●	○	○
Commodities	●	○	●	○
Shadow FFR	●	●	○	○

Source: Authors' calculations.
Notes: White circles refer entries in the Λ matrix that are set to zero, whereas black circles correspond to unconstrained entries.

Sovereign-Debt Crises and Floating-Rate Bonds

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The choice of sovereign-debt maturity in countries at risk of default represents a complex set of competing forces. The tradeoffs reflect the underlying frictions present in international sovereign-debt markets.

The primary frictions are the lack of state contingency in debt contracts and the inability of the government to commit to future actions. These generate two forces in terms of maturity choice. The first is that long-term bonds may be a useful tool for a government to hedge shocks to the cost of funds, say arising from business cycle fluctuations. However, the lack of contingency opens the door to default occurring in equilibrium. Because of the government's inability to commit to future fiscal decisions, bondholders are subject to future dilution of their claims. This generates an opposite force: short-term bonds provide protection from future dilution and, as we shall see, provide better incentives to the government to minimize the costs of default.

This trade-off between insurance and incentives is fundamental to the maturity choice but misses another element. The presence of a

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significant stock of debt in short-maturity by itself generates another potential risk: it leaves the government vulnerable to self-fulfilling rollover crises. This is probably the main drawback of short-term debt —and perhaps the reason why so many restructurings involve lengthening the maturity structure.

In this paper, we explore the advantage to a country of issuing long-maturity debt with a floating-rate coupon. Through both a simple analytical framework, as well as in a richer quantitative framework, we explore the relative benefits of such bonds. We show that having a coupon on a long-term bond indexed to one-period-ahead default probabilities provides all the incentive properties of one-period bonds, without the vulnerability to rollover risk. This can be implemented by indexing the coupon to the auction price of a small amount of one-period bonds.

The framework we explore has both dilution and rollover risk. Dilution risk is well-known in the literature.[1] Aguiar and others (2019) argue that when default risk is high, it is optimal for the government to issue only short-term bonds. This is the case in many real-world crises, as originally documented by Broner and others (2013). Indeed, Bocola and Dovis (2019) argue that the observed shortening of maturity of new issuances of Italian bonds implies a limited role for rollover risk in the European debt crisis. This runs counter to the conventional wisdom that developed in the wake of Mario Draghi's "Whatever it takes" speech in the summer of 2012.

That wisdom holds that the crisis was a self-fulfilling run by creditors that was solved by the European Central Bank stepping in as the lender of last resort.

Rollover risk was a prominent theme after Mexico's 1994–95 crisis. Cole and Kehoe (1996) and Cole and Kehoe (2000) used that crisis as a launching point for their model of rollover risk. Alesina and Tabellini (1990) provide an earlier analysis of self-fulfilling failed auctions. In fact, our discussion of dilution versus rollover risk mirrors that of Alesina and Tabellini (1990), who discuss the experience of floating-rate Italian nominal bonds as the best response to weak inflation credibility and rollover risk.

Aguiar and Amador (2023) provide some evidence of the presence of rollover risk. In particular, they analyze market swaps that involve issuing long-term bonds to repurchase short-maturity bonds. For a case

1. Chatterjee and Eyigungor (2012), Hatchondo and Martínez (2009), Arellano and Ramanarayanan (2012).

involving the Dominican Republic in 2020, they show that the price of *all* bonds increases at the time of the swap, including those of the long-term bonds being issued. They use an analytical framework similar to the one used below to argue that this is evidence that rollover risk is a prominent feature of the data. The environments we study hew fairly close to the quantitative sovereign-debt literature. The main source of risk is endowment risk, to which we add the possibility of a self-fulfilling failed auction. The calibration is based on the benchmark long-term bond paper, Chatterjee and Eyigungor (2012). We find that issuing floating-rate bonds eliminates the risk of a self-fulfilling run while preserving the incentives of one-period bonds. In particular, the government's welfare in the floating-rate bond model in the presence of rollover risk is similar to that of a government with one-period bonds and zero chance of a rollover crisis. Moreover, the floating-rate model dominates the fixed-rate long-term bond model. Welfare gains of switching to floating-rate bonds at zero debt are roughly one percent of consumption. A few caveats are in order to temper these conclusions. One is that we assume the government can auction small amounts of one-period bonds in order to index the coupon payments on the long-term floating-rate bond. This abstracts from liquidity issues in bond markets. Moreover, Alesina and Tabellini (1990) argue that there is evidence that the Italian benchmark-bond auctions may have been manipulated by the government, a possibility we omit from the analysis. Finally, we incorporate the hedging benefits of long-maturity bonds by having persistent income shocks. However, this omits other sources of risk that can be hedged by long-term bonds, such as shocks to risk premia or the risk-free rate.

While we focus on floating-rate bonds, other bond covenants can be used to deal with both dilution and rollover risk. Floating-rate debt is subject to its own source of multiplicity, as studied by Calvo (1988) and, more recently, Ayres and others (2018). Calvo argues that refusing to issue at a high interest rate can help select the best equilibrium. In this spirit, a cap on the coupon can mitigate the risk of this multiplicity, something we also discuss and incorporate in our analysis. Hatchondo and others (2016) discuss covenants that compensate legacy lenders for capital losses as a solution to dilution.

Finally, beyond contract covenants, fiscal rules[2] have been proposed as the solution to dilution, and alternative auction protocols[3] have

2. For example, Hatchondo and others (2012).
3. For example, Chamon (2007).

been proposed to remove rollover risk. The advantage of floating-rate bonds is that they do not require a commitment to enforce fiscal rules or other nonmarket mechanisms; instead, they rely only on competitive markets to deliver the beneficial features.

The paper is organized as follows. Section 1 introduces the general framework absent rollover risk, section 2 provides some analytical results on the efficiency of one-period bonds, section 3 introduces rollover risk, section 4 presents the results of the quantitative exercises, and section 5 concludes.

1. A General Framework

Our framework is based on the standard environment popular in the quantitative sovereign-debt literature.[4] We extend this framework slightly by allowing for floating-rate-coupon bonds. We also alter the model to allow for rollover risk. For expositional reasons, we hold off on the rollover risk extension until after discussing key properties of the baseline model.

Consider a discrete-time, small open economy model. Time is indexed by $t = 0,1,2...$ and the state of nature in time t is given by $s_t \in \mathbb{S}$. The state will index output, default penalties, and, in the extension, include a sunspot that coordinates lenders' beliefs. The state s_t follows a first-order Markov process. In each period, the economy receives a stochastic endowment $y_t = y(s_t)$ that takes values in some discrete, strictly positive, bounded set.

The economy is run by a government with preferences:

$$E\sum_t \beta^t u(c_t),$$

where c_t is consumption of a freely traded good. We assume u is strictly increasing and strictly concave.

The government trades financial assets with competitive, risk-neutral lenders who discount at rate $R^{-1} = (1 + r)^{-1}$. We assume $\beta R \le 1$. Financial trade is restricted to a noncontingent bond. A bond is characterized by a maturity and a coupon. Each unit of debt matures with probability $\lambda \in [0,1]$, which is *iid* across units. In any

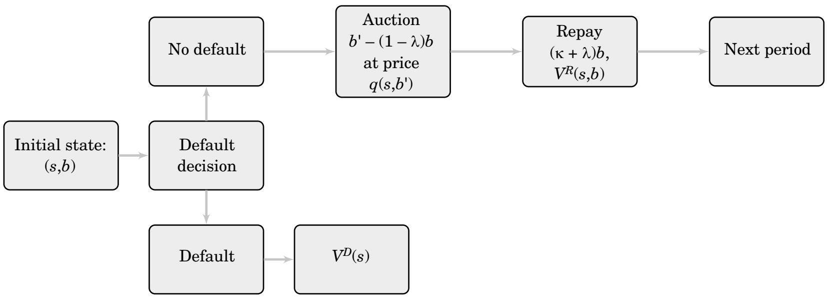
4. See Aguiar and Amador (2021) for a textbook treatment.

nontrivial portfolio, we therefore assume the fraction λ matures and the fraction $1 - \lambda$ remains. The expected maturity is $1/\lambda$. When $\lambda = 1$, we have one-period bonds, and when $\lambda = 0$, we have a perpetuity. Such "perpetual-youth" bonds are a tractable approach to handling bonds of long maturity and have been used by Leland (1994), Hatchondo and Martínez (2009), and Chatterjee and Eyigungor (2012) among others.

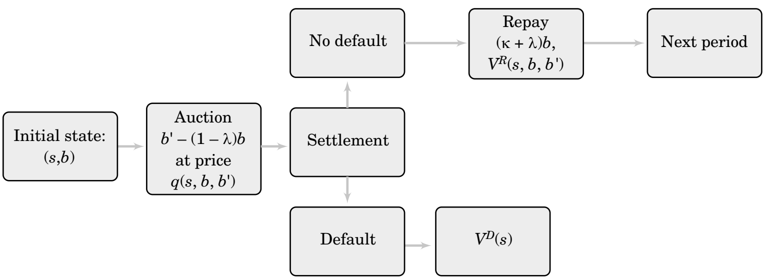
Let $b = b_{t+1}$ be the face value of debt at the end of period t and κ the promised coupon. In $t + 1$ the government owes payments of $(\kappa + \lambda)b$ in every state s_{t+1}. To rule out Ponzi schemes, let $\overline{B}$ denote some arbitrary upper bound on debt issuance and restrict $b \in \mathbb{B} = (-\infty, \overline{B}]$. By making $\overline{B}$ such that promised payments are never greater than the natural debt limit, we ensure it never binds along the equilibrium path, and we will suppress the constraint from the notation going forward.

Figure 1. Within-Period Timing

(a) Eaton-Gersovitz timing



(b) Cole-Kehoe timing



We focus on Markov equilibria, in which equilibrium objects are functions of the exogenous state s_t as well as the government's indebtedness. Let $q : \mathbb{S} \times \mathbb{B} \to [0,1]$ denote the price schedule, and $\mathcal{K} : \mathbb{S} \times \mathbb{B} \to [0, \bar{\kappa}]$ denote the coupon schedule. The coupon is bounded above by a maximum $\bar{\kappa}$, which will be discussed in detail below. For both q and $\mathcal{K}$, the first argument refers to the date on which debt is issued and the coupon is promised, respectively. There is no ex-post contingency in the coupon payment once the state of the next period is realized.

We consider two timing conventions. The first is the "Eaton-Gersovitz" (EG) timing, which is the standard in the literature since Aguiar and Gopinath (2006), Arellano (2008), and Hamann (2002). Under EG timing, depicted in panel (a) of figure 1, the government first observes nature's draw of s, then commits to either repay or default on outstanding debt and then auctions off new bonds. In the alternative, "Cole-Kehoe" (CK) timing, the government, after observing s, first auctions new debt and then decides whether to repay or default on outstanding debt. The key distinction is whether the result of the auction plays a role in the repayment decision. In EG timing, repayment is independent of the realized auction price, while in CK repayment is contingent on the success or failure of a bond auction. We begin by discussing the equilibrium under EG timing.

1.1 The Government's Problem

If the government defaults at time t in state s, we assume it receives value $V^D(s)$. In particular,

$$V^D(s) = u\left(y^D(s), s\right) + \beta \mathbb{E}_s\left[\theta V(s', 0) + (1-\theta) V^D(s')\right]. \tag{1}$$

The term $y^D(s)$ is the endowment received in default when the state is $s \in \mathbb{S}$. This captures any punishment in terms of loss of endowment due to default as well as the fact that the government must consume hand-to-mouth while excluded from financial markets. With probability θ, the government regains access to bond markets and starts anew with zero debt and value $V(s', 0)$ in state s'. With probability $1 - \theta$, the government remains in the default state.

If the government has opted to repay, the government's value satisfies the following Bellman equation:

$$V^{R}\left(s,b,\kappa\right)=\max_{b'}\left\{u\left(y\left(s\right)-\left(\kappa+\lambda\right)b+q\left(s,b'\right)\left(b'-\left(1-\lambda\right)b\right)\right)\right.$$
$$\left.+\beta\mathbb{E}\max\left\{V^{R}\left(s',b',\mathcal{K}\left(s,b'\right)\right),V^{D}\left(s'\right)\right\}\right\}. \tag{2}$$

Here, the government takes the schedules q and $\mathcal{K}$ as given and optimally chooses b'. The continuation value reflects that the government has the option to default next period after observing s'. Given that $\kappa = \mathcal{K}(s,b)$ is pinned down in equilibrium by the states, we can redefine the government's value as a function of (s, b) and the lagged state, (s_{-1}). Henceforth, we write $V^R(s_{-1}, s, b)$, with $\kappa = \mathcal{K}(s_{-1},b)$ being the coupon that is due in the current period.

Let $\mathcal{B}:\mathbb{S}\times\mathbb{S}\times\mathbb{B}\rightarrow\mathbb{B}$ denote the optimal policy function of the government. Implicitly in problem (2), we are assuming that there exists a b' such that it is feasible to repay; that is, $y(s) - (\kappa - \lambda)\, b + q\,(s, b')\,(b' - (1 - \lambda)b) \geq 0$ for some $b' \in \mathbb{B}$. If this is not the case, we set $V^R = -\infty$ so that the government defaults whenever repayment is infeasible.

Define $V(s_{-1}, s, b) \equiv max\{V^R(s_{-1}, s, b), V^D(s)\}$ to be the government's value at the start of the period. The government repays if $V^R(s_{-1}, s, b) \geq V^D(s)$ and defaults otherwise. Let $\mathcal{D}:\mathbb{S}\times\mathbb{S}\times\mathbb{B}\rightarrow\{0,1\}$ denote the optimal default policy, with the value one indicating default and zero indicating repayment.

1.2 The Lenders' Break-Even Condition

The restriction on equilibrium prices is that lenders break even in expectation. In particular:

$$q\left(s,b\right)=R^{-1}\mathbb{E}\left[\left(1-\mathcal{D}\left(s,s',b\right)\right)\left(\left(\kappa+\lambda\right)+q\left(s',\mathcal{B}\left(s,s',b\right)\right)\left(1-\lambda\right)\right)\right], \tag{3}$$

where $\kappa = \mathcal{K}\,(s, b)$.

We consider two alternative coupon schedules. The standard approach is a constant coupon. In particular, define the "fixed-rate coupon schedule" as $\mathcal{K}\,(s, b) = \kappa$ for all $(s, b) \in \mathbb{S}\times\mathbb{B}$ for some constant $\kappa \leq \overline{\kappa}$.

The second is a floating-rate coupon. In particular, consider the equilibrium price of a one-period, zero-coupon bond given the equilibrium behavior of the government:

$$q_S\left(s,b\right) \equiv R^{-1}\mathbb{E}_s\left(1 - D\left(s,s',b\right)\right). \tag{4}$$

Note that q_s lies between zero and R^{-1}. Define the "floating-rate coupon schedule" as:

$$K\left(s,b\right) = min\left\{\frac{1}{q_S\left(s,b\right)} - 1, \overline{\kappa}\right\}. \tag{5}$$

This coupon compensates the bondholder for the one-period-ahead risk of default. It is important to keep in mind that the equilibrium behavior is for an environment with a single bond of inverse maturity λ and coupon $\mathcal{K}$; unless $\lambda = 1$, there is no short-term bond actively traded. Nevertheless, given this equilibrium behavior, we can construct a q_s and $\mathcal{K}$. In particular, q_s is the price that would obtain in equilibrium if an infinitesimal amount of one-period bonds were issued along with the benchmark bonds.

The equilibrium in the floating-rate model depends on $\mathcal{K}$, which, in turn, depends on the default policy function. We are looking for a fixed point for this mapping. There may be more than one, as we discuss at the end of this section.

1.3 Equilibrium

We are now ready to define an equilibrium:

Definition: An Eaton-Gersovitz equilibrium is a price schedule q, a coupon schedule $\mathcal{K}$, a value function V^R with associated policies $\mathcal{B}$ and $\mathcal{D}$, and a default value V^D such that: (i) The lenders' break-even condition (3) is satisfied given $\mathcal{B}$, $\mathcal{K}$, and $\mathcal{D}$; (ii) given $\mathcal{D}$, $\mathcal{K}$ is either fixed or determined by equations (4) and (5); (iii) given q and $\mathcal{K}$, V^R solves the government's Bellman equation (2) with optimal policy $\mathcal{B}$, (iv) $\mathcal{D}\ (s, b, \kappa) = 1$ if $V^R(s, b, \kappa) < V^D\ (s)$ and zero otherwise, and (v) given V^R, V^D, solves the recursion (1).

1.4 Prices and Future Fiscal Policies

The two alternative coupon structures have different implications for how future fiscal policy affects bond prices. Under the fixed-rate schedule, equation (3) indicates that for $\lambda < 1$ the debt-issuance policy function $\mathcal{B}(s', b)$ affects the price of the non-maturing bonds next period and hence affects the price of bonds today. This is the standard channel in which lack of commitment to future fiscal policy potentially 'dilutes' existing bondholders and depresses the value of long-term bonds. We shall return to this below.

Now consider the floating-rate coupon. Suppose that in equilibrium $\mathcal{B}$ is such that there is an upper bound on the ergodic distribution of debt, $B_{max} < \overline{B}$. Moreover, suppose that $q_{min} \equiv \min_{s \in \mathbb{S}} q_S(s, B_{max}) > 0$. That is, along the equilibrium path, the government never issues debt to the point that it will default with probability one the next period. Both of these conditions are typically satisfied in standard quantitative sovereign-debt models. Then, if $\overline{\kappa} > 1/q_{min} - 1$, a valid equilibrium price schedule is $q(s,b) = 1$ for all $s \in \mathbb{S}$ and $b \leq B_{max}$. To see this, define the price operator T_q by equation (3):

$$
\begin{aligned}
[T_q q](s,b) &= R^{-1}\mathbb{E}\Big[\big(1-\mathcal{D}(s,s',b)\big)\Big(\kappa+\lambda+q\big(s',\mathcal{B}(s,s',b)\big)(1-\lambda)\Big)\Big] \\
&= R^{-1}\mathbb{E}\left[\big(1-\mathcal{D}(s,s',b)\big)\left(\frac{1}{q_s(s,b)}-1+\lambda+q\big(s',\mathcal{B}(s,s',b)\big)(1-\lambda)\right)\right] \\
&= 1+(1-\lambda)R^{-1}\mathbb{E}_s\big(1-\mathcal{D}(s,s',b)\big)\Big(q\big(s',\mathcal{B}(s,s',b)\big)-1\Big),
\end{aligned}
$$

where the last line uses the definition of q_s. This operator maps bounded functions on the domain $\mathbb{S} \times (-\infty, B_{max}]$ into itself, and satisfies the Blackwell conditions for a contraction. For any $\mathcal{B}$ such that $\mathcal{B}(s_{-1}, s, b) \leq B_{max}$ on this domain, $q = 1$ is the unique fixed point of the price operator. In this scenario, the price is constant and, more importantly, independent of future fiscal policy.

As noted above, in the floating-rate case $\mathcal{K}$ is defined by q_s, which in turn depends on equilibrium behavior. The latter depends on $\mathcal{K}$. There may be multiple fixed points of this mapping. This is multiplicity in the spirit of Calvo (1988). In particular, without the upper bound $\overline{\kappa}$, there is an equilibrium with zero borrowing. To see this, posit the schedule $\mathcal{K}(s, b) = \infty$ for all $s \in S$ and $b > 0$. For any $b > 0$, it is infeasible for the government to repay, and hence the government will default with probability one, validating $q_s = 0$ and $\mathcal{K} = \infty$. For this reason, we introduce the cap on coupons to rule out this extreme equilibrium. At this stage, we do not have sufficient conditions to ensure that there is a unique floating-rate equilibrium.

2. One-Period Bonds as a Planning Problem

With long-term fixed-rate bonds, the existing bondholders are at the mercy of future fiscal policy. One-period fixed-rate bonds do not feature this risk. A useful way to see this advantage of one-period bonds is to consider the dual of problem (2), as done in Aguiar and Amador (2019).

Specifically, consider problem (2) for the case of $\lambda = 1$ and normalize $\kappa = 0$. Then (2) can be written as:

$$V^R(s,b) = \max_{c,b'} \left\{ u(c) + \beta \mathbb{E}_s \max\left\{ V^R(s',b'), V^D(s') \right\} \right\}$$
$$\text{subject to } c \leq y(s) - b + q(s,b')b'.$$

Because $\kappa = 0$, we can drop the lagged s_{-1} as an argument for this exercise. As shown by Aguiar and Amador (2019), on the relevant domain for bonds,[5] $V^R(s, b)$ is strictly decreasing in b for each s. Let $B(s, v)$ denote the inverse of V^R. That is,

$$V^R(s, B(s, v)) = v.$$

Given the strict monotonicity of V^R, B solves the dual problem:

$$B(s,v) = \max_{c,b'} \left\{ y(s) - c + R^{-1} b' \mathbb{E}_s \mathbf{1}_{\left\{V^R(s,b') \geq V^D(s')\right\}} \right\}$$
$$\text{subject to } v = u(c) + \beta \max\left\{ V^R(s,b'), V^D(s') \right\},$$

5. By relevant domain, we mean the domain on which the government can feasibly repay. See Aguiar and Amador (2019) for more details.

where 1_x is the indicator function that equals one when x is true and zero otherwise, and where we have used the equilibrium condition $q = R^{-1} 1_{\{V^R \geq V^D\}}$. As $V^R(s, b')$ is strictly decreasing, the choice of b' is also the choice of the government's continuation value. In particular, we can think of adding $v(s')$ as a choice variable subject to the constraint that $v(s') = V^R(s', b')$ for all s' such that $V^R(s', b') \geq V^D(s')$. This constraint is equivalent to $B(s', v(s')) = b'$ for all s' such that $v(s') \geq V^D(s')$. This leads to the following problem:

$$B(s,v) = \max_{c,b',\{v(s')\}} y(s) - c + R^{-1} b' \mathbb{E}_s 1_{\{v(s') \geq V^D(s')\}} \tag{6}$$

subject to:

$$v = u(c) + \beta \mathbb{E}_s \max\{v(s'), V^D(s')\}$$

$$b' = B(s', v(s')) \text{ for all } s' \text{ such that } v(s') \geq V^D(s').$$

Problem (6) is similar to an optimal contracting problem. The principal (lender) chooses a sequence of consumption and continuation values for the agent (the government) subject to a promise-keeping constraint and the 'spanning' condition $b' = B(s', v(s'))$. This last condition restricts the span of continuation values and reflects that the one-period bond is noncontingent.

The spanning constraint contains an equilibrium object (the inverse value function). An alternative maturity structure would involve a different restriction on spanning. It may be the case that long-term bonds allow for better hedging of risk, and a true planning problem will not be constrained from implementing such an allocation.

Aguiar and Amador (2019) note that equation (6) defines an operator that maps B in the spanning constraint into the B that equals the maximized payoff to lenders. They show that this mapping is a contraction and therefore there is a unique equilibrium in the one-period bond model.[6]

Note that the Principal cannot prevent the government from walking away from the contract and taking the outside option V^D.

6.To do this, it is first necessary to relax the spanning condition to an inequality. See that paper for details. In addition, the result requires that there is no re-entry to financial markets after a default, that is, $\theta = 0$; so that v^D is exogenously given. For an alternative contraction mapping approach, see also Bloise and Vailakis (2022).

Nevertheless, absent default, the choice of c and b' maximizes the joint surplus conditional on the spanning condition. In particular, the equilibrium is the same regardless of whether the government or the lenders set fiscal policy, reflecting that incentives are aligned with one-period bonds.

This alignment of incentives is not true for long-term bonds, and we cannot write the long-term bond equilibrium as a pseudo-planning problem like (6). One way to see why not mechanically, is that there are three relevant variables for long-term bonds: the face value of debt b, the government's value v, and the market value of debt $q \times b$. With long-term bonds, the equilibrium q depends on future policies that are beyond the control of current actors (either lenders or the incumbent government). In the one-period bond model, absent default, the market value and face value coincide at the start of the period.

2.1 An Example

To provide a little more insight into why incentives are aligned regarding fiscal policy in the one-period bond model, we shut down the endowment fluctuation; that is, $y(s) = y$ for all $s \in S$. The only risk is the value of default $V^D(s)$, which we allow to vary with the state. Let s be *iid* over time and be such that $V^D(s) = v^D$ is drawn from a continuous distribution with CDF $F(v^D)$ and support $[\underline{V}, \overline{V}]$.

With this *iid* shock process, once the government decides to repay, the realized value of s is irrelevant, and we can drop it as a state variable. That is, $V^R(s, b)$ can be written $V^R(b)$, and its inverse is $B(v)$. In the dual problem, there is a single continuation value v' and the spanning condition becomes $b' = B(v')$. In this case, we can substitute the spanning condition into the objective and use the fact that the government repays if $v^D \leq v'$ to write the dual problem as:

$$B(v) = \max_{c,v'} y - c + R^{-1}F(v')B(v')$$

subject to: $v = u(c) + \beta\left(F(v')v' + \int_{v'}^{\overline{V}} v^D dF(v^D)\right).$

This is a true planning problem, subject to limited participation of the government. The key distinction between this problem and the original (6) is that, without income fluctuations or persistence in the outside option, there is no risk that can be hedged. Bonds of any maturity will either be defaulted on or will have a price that is invariant to v^D conditional on repayment.

The planner's inverse Euler equation for this problem (assuming $v' \in (\underline{V}, \overline{V})$) is:

$$\frac{1}{u'(c')} = \frac{\beta R}{u'(c)} + \frac{f(v')B(v')}{F(v')}, \tag{7}$$

where $f = dF/dv$ and c' is next period's consumption conditional on repayment and the optimal choice v'. To gain some intuition, set $\beta R = 1$ and let $u(c) = log(c)$. We then have

$$c' = c + \frac{f(v')B(v')}{F(v')}.$$

The second term on the right-hand side is the marginal probability of default times the amount of debt. If this is strictly positive, then the optimal plan sets $c < c'$. That is, the optimal plan is to save. And the rate of saving is determined by the marginal decline in default probability. The greater $f(v')/F(v')$, the stronger the incentive to save at the margin. This reflects that the risk to the lender is the amount of debt outstanding times the probability of default. The optimal contract internalizes that saving reduces this risk.

Now recall that the optimal contracting problem is just an alternative view of the equilibrium in which the government makes all decisions. Why does the government want to reduce the risk of default? Keep in mind that the government strategically defaults, so at the moment of default, it captures an increase in value. Why not just wait for a high v^D (say a bailout or forgivable default) and then default?

In equilibrium, it is the price schedule that aligns incentives. Specifically, $q(b') = R^{-1}F(V^R(b'))$. Differentiating:

$$q'(b') = R^{-1} f(v') V^{R'}(b'),$$

where $v' = V^R(b')$. From the envelope condition, $V^R(b') = -u'(c')$ Substituting in, equation (7) becomes:

$$u'(c)\left(1 + \frac{q'(b')b'}{q(b')}\right) = \beta R u'(c').$$

In the equilibrium, the government saves because $q'(b') < 0$, and it understands that, by saving, it will issue/roll over its bonds at a higher price. In particular, the government captures the *entire* benefit of reducing default risk via high prices, and therefore incentives are aligned between borrower and lender to minimize the risk of default.

Now, it is also the case that $q'(b) < 0$ with long-term bonds. However, the government is not rolling over its entire stock of debt. Thus, it does not internalize the entire cost of default to the lender, which involves new bonds as well as legacy bonds, and hence does not capture the entire benefit of reductions in default risk. At the extreme of a perpetuity, the government does not have to roll over any debt and has no incentive to reduce the risk of default. This is the sense that fiscal policy is inefficient with long-term bonds.

2.2 Floating Rate Bonds

With these insights in hand, we can now see one of the advantages of floating-rate bonds. In particular, if the coupon on the entire stock of debt reflects the default probability, the government has the same incentive to save as in the case of one-period bonds.

More formally, consider the case discussed in the previous section in which $q(\mathrm{s}, b) = 1$ in a floating-rate equilibrium for a domain that encompasses the ergodic support, $b \leq B_{max}$. The government's value conditional on repayment is $V^R(s_{-1}, s, b)$. Recall that the original value function was written $V^R(s, b, \kappa)$. For an equilibrium $\mathcal{K}$, we replaced κ with s_{-1}. To construct a pseudo-planning problem, we do not substitute out $\mathcal{K}$ but include it explicitly as a constraint in the dual problem. Specifically, let $B(s, v, \kappa)$ be the inverse of $V^R(s, b, \kappa)$. The government's budget constraint (with $q(\mathrm{s}, b') = 1$) is:

$$c = y(s) - (\kappa + \lambda)b + b' - (1 - \lambda)b.$$

Let $\tilde{B}(s,v,\kappa) = (1+\kappa)B(s,v,\kappa)$ and $\tilde{b} = (1+\kappa)b$. The dual problem becomes:

$$\tilde{B}(s,v,\kappa) = \max_{c,b',\kappa',\{v(s')\}} \left\{ y(s) - c + \frac{\tilde{b}'}{1+\kappa'} \right\} \tag{8}$$

subject to:

$$v = u(c) + \beta \mathbb{E}_s max\{v(s'), V^D(s')\}$$

$$\tilde{b}' = \tilde{B}(s', v(s'), \kappa') \text{ for all } s' \text{ such that } v(s') \geq V^D(s')$$

$$\kappa' = \mathcal{K}(s, b'),$$

where we have suppressed the ergodic set constraint that $v(s')$ must be such that $b' = B(s', v(s'), \kappa') \leq B_{max}$, as it should not bind in this case.

If we allow the pseudo-planner to "see through" the equilibrium $\mathcal{K}$, we can characterize the best equilibrium with a planning problem that is isomorphic to (6). This resolves the Calvo multiplicity in favor of the efficient outcome. Specifically, recall that

$$\frac{1}{1+\mathcal{K}\left(s,b'\right)} = R^{-1}\mathbb{E}_s \mathbf{1}_{\left\{V^R(s,b',\kappa)\geq V^D(s')\right\}}.$$

Replacing $V^R(s', b', \kappa)$ with $v(s')$, we obtain:

$$\frac{1}{1+\mathcal{K}\left(s,b'\right)} = R^{-1}\mathbb{E}_s \mathbf{1}_{\left\{v(s')\geq V^D(s')\right\}}.$$

Inspection of the value function (8) shows that we can drop κ as an argument of $\tilde{B}$. Let $\overline{B}$ (s, v) represent the best possible equilibrium, then we have that $\overline{B}$ solves:

$$\tilde{B}\left(s,v\right) = \max_{c,b',\{v(s')\}} \left\{ y\left(s\right) - c + R^{-1}b'\mathbb{E}_s \mathbf{1}_{\left\{v(s')\geq V^D(s')\right\}} \right\} \tag{9}$$

subject to:

$$v = u\left(c\right) + \beta\mathbb{E}_s max\left\{v\left(s'\right), V^D\left(s'\right)\right\}$$

$$\tilde{b}' = \overline{B}\left(s', v\left(s'\right)\right) \text{ for all } s' \text{ such that } v\left(s'\right) \geq V^D\left(s'\right).$$

This is the same as problem (6). Thus, conditional on selecting the best equilibrium, the floating-rate bond provides all the same incentive and spanning features as the one-period bonds. The one caveat about the mapping from floating-rate to one-period bonds is the potential for Calvo multiplicity.

3. Rollover Risk

To introduce rollover risk, we alter the timing within a period as in Cole and Kehoe (2000). The government first auctions debt and then decides to repay maturing debt. This timing makes the repayment decision contingent on the outcome of the auction.[7]

We begin with the fixed-rate coupon environment. Working backward through the period, suppose the government has issued $b'-(1-\lambda)$ bonds at price q during the auction. At the time of settlement, the government's value of repayment is:

$$V^R(s,b,b',q) = u\left(y(s) - (\kappa+\lambda)b + q \times \left(b' - (1-\lambda)b\right)\right) + \beta \mathbb{E}_s V(s',b'),$$

where we have repurposed the notation to fit the current environment. We can let s index the price as well, so that $q = q(s, b, b')$ and drop q as an argument of the repayment value.

The default payoff is the same as in the EG benchmark.[8] The government defaults if $V^R(s, b, b') < V^D(s)$. The government's problem at the time of auction is:

$$V(s,b) = \max\left\{\max_{b'} V^R(s,b,b'), V^D(s)\right\}.$$

Note that there is perfect foresight within a period, and hence the government knows what the payoffs to repayment and default are. Let $\mathcal{B}(s,b) = argmax_{b'} V^R(s,b,b')$ denote the debt-issuance policy, and $\mathcal{D}(s,b) = 1$ if $\max_{b'} V^R(s,b,b') < V^D(s)$ and zero otherwise.

To see the indeterminacy in this environment, consider fixing the continuation equilibrium. Specifically, hold the function $\mathbb{E}_s V(s', b')$ constant in the government's problem, as well as future policies. Let $\bar{q}(s, b')$ be the break-even price conditional on repayment in the current period. That is,

$$\bar{q}(s,b') = R^{-1}\mathbb{E}_s\left[\left(1 - \mathcal{D}(s',b')\right)\left[\kappa + \lambda + (1-\lambda)q\left(s',b',\mathcal{B}(s',b')\right)\right]\right].$$

Note that this is identical to (3); the only difference is that the policy functions may differ in an environment with rollover risk. This is the 'good' equilibrium.

To see the 'crisis' equilibrium, suppose that $q(s, b, b') = 0$ for all $b' \geq 0$. In this case, for $b' \geq 0$,

$$V^R(s,b,b') = u\left(y(s) - (\kappa+\lambda)b\right) + \beta E V\left(s',(1-\lambda)b\right).$$

7. See panel (b) of figure 1.

8. For simplicity, we assume that if the government auctions debt at a positive price and then defaults, the auction proceeds are lost to both parties. On the equilibrium path, this never occurs.

The government must pay the entire amount of maturing debt plus coupon out of current endowment. It then carries over non-maturing debt into the next period. This is a failed auction. If $u(y(s) - (\kappa + \lambda) b) + \beta \mathbb{E} V(s',(1 - \lambda)b) < V^D(s)$, then a zero price is consistent with the lenders' break-even condition.[9] The lenders see that the government will default at settlement, and refuse to pay a positive price at auction. Such a scenario is possible if $(\kappa + \lambda)b$ is large relative to y.

For pairs of b such that both $q = 0$ and $q = \bar{q}$ are possibilities, following Cole and Kehoe (2000) we let a sunspot coordinate beliefs. That is, s contains a random variable that takes a value of one for a crisis and zero otherwise.

In the case of short-term debt, $\lambda = 1$, the debt burden is particularly painful after a rollover crisis. This is the logic behind why short-term debt makes a government particularly vulnerable to a rollover crisis. Conversely, if $\lambda = 0$, for a given face value the repayment burden is light, and a crisis is possible only for very large b.

This sets up the canonical maturity dilemma. On the one hand, short-term debt provides correct incentives. On the other, it exposes the country to rollover risk, and, perhaps, offers less spanning of income risk. A floating-rate coupon bond provides the same incentives as one-period debt, but defers the maturity payments, mitigating rollover risk. Indeed, if we ignore spanning (as in our simple model without income risk), then the floating-rate perpetuities offer the best of both worlds—correct incentives but limited rollover risk.[10]

The only drawback is that a long-term bond may provide better hedging of income and other potential risks, but this is a quantitative question. In the next section, we therefore turn to a quantitative model that incorporates floating-rate debt and noninsurable income risk.

4. A Quantitative Model

In this section, we introduce income risk as well as rollover risk in a quantitative model. We explore five alternatives: a one-period bond EG model (henceforth EG-ST); a one-period bond model with rollover risk (CK-ST); the same two environments but with long-term

9. We assume the government cannot repurchase long-term bonds at zero price. See Aguiar and Amador (2013) for how this can be supported in equilibrium.

10. A floating-rate equilibrium is constructed in the presence of rollover risk along the same lines as in the benchmark EG model. That is, we price a one-period bond, which now must compensate lenders for rollover risk as well as 'fundamental' default risk and set the coupon to compensate lenders for that risk.

fixed-rate bonds (EG-LT and CK-LT); and finally a long-maturity floating-rate bond (FR) with rollover risk. As we shall see, the long-maturity floating-rate bond eliminates the risk of a rollover crisis, so we do not need to present the floating-rate bond under the Eaton-Gersovitz timing in addition to the Cole-Kehoe timing.

The benchmark parameterization is the same as Chatterjee and Eyigungor (2012) (henceforth, CE12).[11] The model is quarterly. The underlying process for log income follows:

$$\ln y_t = x_t + z_t$$
$$x_t = \rho x_{t-1} + \varepsilon_t$$
$$\varepsilon \sim N\left(0, \sigma_\varepsilon^2\right)$$
$$z \sim \text{Truncated } N\left(0, \sigma_z^2\right).$$

Following CE12, we set $\rho = 0.95$, $\sigma_\varepsilon = 0.027$ and $\sigma_z = 0.01$. The persistent process x is approximated by Tauchen's method with a span of three standard deviations of the long-run distribution. The *idd* shock z is a truncated Normal with support $[-2\sigma_z, +2\sigma_z]$, and is included for computational reasons, as discussed by CE12. In default, the endowment is reduced by a quadratic factor. Specifically,

$$\ln y_t^D = x_t^D + z_t$$
$$x_t^D = x_t - \max\left\{0, -0.189 x_t + 0.246 x_t^2\right\}.$$

In the first period of default, we set $z = \underline{z}$, its minimum value.

The government's preferences consist of a constant relative risk aversion felicity with a risk-aversion parameter 2 and a discount factor $\beta = 0.95$. The risk-free interest rate is $R = 1.01$.

The benchmark maturity is $\lambda = 0.05$ or an expected maturity of 20 quarters. For the one-period bond models, we set $\lambda = 1.0$. We set $\kappa = 0.01$ for all models with fixed-rate bonds. And let $\overline{\kappa} = 0.06$ in the baseline specification of the floating-rate bonds.

Finally, in the environments with rollover risk, we set the probability of a sunspot to 10 percent quarterly, although the frequency of crises will be lower in equilibrium. A rollover crisis occurs only if the sunspot is realized and debt is large enough. We assume the

11. The code and additional computational information is available at https://github.com/manuelamador/floating-rate-debt.

probability of a crisis is *iid* over time. See Bocola and Dovis (2019) for a quantitative model in which the probability of a crisis follows a persistent process.

In table 1, we report ergodic moments for the five models, plus an additional floating-rate model in which $\overline{\kappa}$ is set at 0.015, fifty basis points higher than the risk-free net interest rate of 0.01. A few things stand out. One is that with short-term debt, the presence of rollover risk looms large. Comparing EG-ST with CK-ST, debt is much lower in the latter, and one hundred percent of the defaults are due to self-fulfilling runs. With long-term debt, rollover risk is essentially nonexistent, but default is more frequent.

The floating-rate model generates few defaults, with the floating-rate coupon addressing dilution and the long maturity essentially eliminating rollover risk. The corresponding moments for the floating rate and the EG-ST models are very similar.[12] However, in the last column, we report the floating-rate model with $\overline{\kappa} = 0.015$ and the outcome is quite different. The hard cap binds, and this opens the door to dilution risk.

Our focus is on the welfare of the government under alternative arrangements. To evaluate this, we present the value at zero debt for alternative endowments: $V(\cdot, 0)$. In figure 2 panel (a), we plot the value function for the two one-period models (EG-ST and EG-CK) as well as the floating-rate model. The horizontal axis traces out alternative initial endowment states.

The EG-ST and FR values are indistinguishable, while the Cole-Kehoe short-term bond model has a distinctly lower value. The fact that rollover risk lowers welfare is intuitive, particularly with short-maturity bonds. As anticipated by the analytical models, the floating-rate model preserves the good features of the one-period model while eliminating the vulnerability to rollover risk.

In panel (b), we plot the consumption equivalent welfare gain between the CK-ST model and the FR model. For low-endowment states, welfare increases by slightly more than one percent, while for high-endowment states the gain is an order of magnitude less. Recall that welfare is evaluated at zero debt, and hence the likelihood of default (whether fundamental or self-fulfilling) lies well in the future.

12. This result however depends on the value of $\overline{\kappa}$. For example, with a value of 1.0, there are more noted differences between the two models. Therefore, there is an intermediate range of values for $\overline{\kappa}$ for which the environments align.

Table 1. Moments of the Ergodic Distribution

	FR $\bar{\kappa}$ –0.060	*EG-ST*	*CK-ST*	*EG-LT*	*CK-LT*	*FR* $\bar{\kappa}$ –0.015
$\mathbb{E}[\frac{b'}{y}]$	0.82	0.82	0.38	0.94	0.94	0.87
$\mathbb{E}[\frac{q\times b'}{y}]$	0.82	0.82	0.38	0.72	0.72	0.78
Default Rate$^{(*)}$	0.003	0.003	0.002	0.067	0.067	0.033
$\mathbb{E}[r-r^*]^{(*)}$	0.003	0.003	0.002	0.080	0.080	0.038
$StDev[r-r^*]^{(*)}$	0.005	0.004	0.003	0.044	0.044	0.029
$\mathbb{E}\kappa$	0.011	0.010	0.010	0.010	0.010	0.011
$StDev(\kappa)$	0.001	0	0	0	0	0.002
Runs/Defaults	0.087	0	1.000	0	0.003	0.003

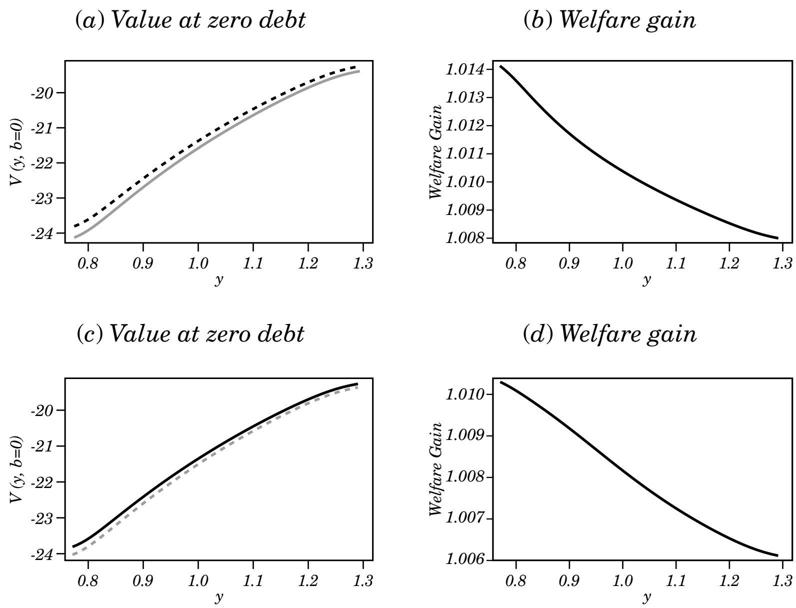
Source: Authors' calculations.
Note: This table reports key moments from the ergodic distribution of each model. All moments are conditional on being in good credit standing for the prior 20 quarters. The first row is the average level of debt issued as a fraction of the endowment. The second row is the average market value of debt issuance, again normalized by the level of endowment. The third row is the annualized frequency of default. The fourth and fifth rows are the mean and standard deviation of implied spreads, respectively. Spreads are computed in annualized form as $(1/q)^4 - R^4$. The sixth and seventh rows are the mean, and standard deviation of the coupon, respectively. The final row is the fraction of defaults that occur due to a self-fulfilling rollover crisis.
(*) Annualized.

In panels (c) and (d) of figure 2 we repeat the same exercises for the long-term bond models. In panel (c), the EG-LT and CK-LT models generate the same value for the government. The reason is that the long-term bonds eliminate the vulnerability to rollover risk. However, the FR model dominates in welfare. This is because the long-term fixed-rate models suffer from dilution risk, something not present with a floating-rate coupon. Panel (d) presents the consumption equivalent welfare gain between FR and CK-LT(=EG-LT). We see that, at low-endowment states, the welfare gain is roughly one percent.

Another approach to evaluating the efficiency of alternative debt instruments is to trace out the frontier between lenders' payoffs and the government's value at different levels of debt. Specifically, consider a state (y_{-1}, y, b). The government's value is $V(y_{-1}, y, b)$, where y_{-1} is an irrelevant state in the fixed-rate environments. The lenders' market value at the start of the period is:

$$MV(y_{-1}, y, b) = \left(1 - \mathcal{D}(y_{-1}, y, b)\right) \times b \times \left[\kappa + \lambda + (1-\lambda)\, q\left(y, \mathcal{B}(y_{-1}, y, b)\right)\right].$$

Figure 2. Government Welfare



Source: Authors' calculations.
Note: Panel (a) depicts the equilibrium value function at zero debt as a function of current endowment. The solid black line represents the floating-rate bond model, the dashed white line represents the short-term EG model, and the solid gray line represents the short-term CK model. Note that the black and the dashed white lines are identical. Panel (b) represents the consumption-equivalent welfare gain for the government between the floating-rate model and the short-term CK model. Panel (c) repeats panel (a) but with the long-term versions of EG and CK. In this case, the EG and CK models are identical. Panel (d) repeats panel (b) comparing the floating-rate model with the long-term CK model.

The value is zero if the government defaults ($\mathcal{D}(\cdot) = 1$). Otherwise, lenders receive the coupon and principal $(\kappa + \lambda)b$, and the market value of non-maturing debt is $q(y, b') \times b$, where b' is the equilibrium debt-issuance policy.

Figure 3 traces out the frontier between *MV* on the vertical axis and *V* on the horizontal axis as we vary b and hold y and y_{-1} at the mean value. Panel (a) contains the short-term fixed-rate models, and panel (b) the long-maturity environments, with both panels containing the floating-rate case as well.

For each frontier, the point furthest to the left on the horizontal axis is the default value for the government. This point represents all b such that the government defaults and lenders receive zero. Note that the default value varies across environments due to the probability of re-entry. Hence, a lower reentry value lowers the default value.

The remaining points represent positive values for the lenders. In panel (a), we see that the one-period EG model (EG-ST) and the floating-rate model lie on top of each other for this parameterization. The floating-rate frontier depends on the coupon, which in the figure depicted is evaluated at the mean endowment (that is, we assume y_{-1} equals the unconditional mean). The CK-ST bond model is clearly dominated by both. The CK-ST model is prone to rollover risk, which depresses the frontier. However, the low default value (due to the low reentry value) of CK-ST enables the government to sustain lower repayment values without defaulting, thus extending the frontier to the left.

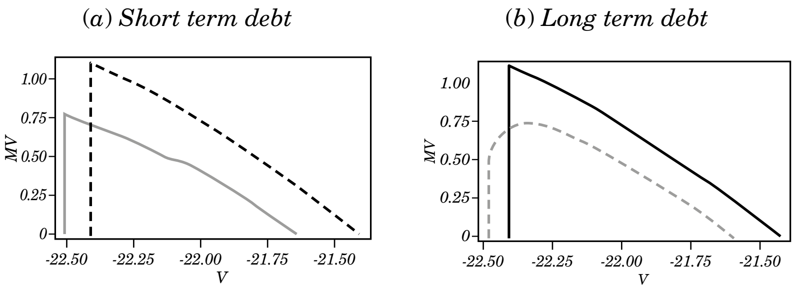
Panel (b) repeats the frontier for the long-term bond model. Recall that in this case, the EG-LT model and the CK-LT model are equivalent, as maturity is such that there is no rollover risk. However, there is the risk of debt dilution. For this reason, the floating-rate frontier dominates the other two. Note that the upward portions of the frontier for the fixed-rate bonds are on the 'wrong' side of the debt Laffer curve. That is, debt forgiveness would increase both lender and government values. This reflects that legacy bondholders are being diluted. Such debt forgiveness is ruled out a priori because it cannot be implemented via voluntary market transactions due to the holdout problem. Hatchondo and others (2014) provide an analysis of negotiated restructurings to alleviate this inefficiency.

5. Conclusion

In this paper we presented analytical and quantitative arguments in favor of long-term bonds with floating-rate coupons. We showed that such bonds combine the incentive properties of one-period bonds with the protection from the rollover risk of a long-term bond. In the presence of rollover and dilution risk, such bonds provide government welfare that dominates both short-term and long-term bonds.

As noted in the introduction, while the analysis includes standard features in the literature, it omits some real-world complications. Perhaps primary among these omissions are the shocks to the global required rate of return.

Figure 3. Pareto Frontiers



Source: Authors' calculations.
Note: This depicts the frontier between lenders' value (vertical axis) and government value (horizontal axis) as b varies, evaluating y and lagged y at the mean. Panel (a) depicts the one-period bond models as well as the floating-rate model. Panel (b) compares the long-term bond models with the floating-rate model. In each panel, the black line is the floating-rate model, the dashed white line is the EG model, and the solid gray line is the CK model.

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*KF*star and Portfolio Inflows: A Focus on Latin America

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Policymakers faced with volatile capital flows may desire a method to identify the level of flows likely to persist in the medium run. In a series of papers (Burger, Warnock, and Warnock, henceforth BWW, 2018, 2022), we have developed an estimate of the natural or equilibrium level of capital flows (*KF*star or *KF**) that provides guidance on the likely amount of portfolio inflows countries can expect over a one- to two-year period.

*KF** is an easy-to-construct slow-moving supply-side benchmark that approximates the level flows should converge to over a medium-term horizon and thus helps gauge the amount of gross portfolio inflows countries can expect to receive. *KF** is a supply-side measure in that it is derived from the supply of rest-of-the-world (*ROW*) savings; in simple terms, it is a lagged portfolio weight, constructed by using portfolio liabilities data from Lane and Milesi-Ferretti (2018), multiplied by current *ROW* savings (from the IMF). The underlying theory is from the Tille and van Wincoop (2010) and Devereux and

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Sutherland (2011) incorporation of portfolio choice in open-economy DSGE models and, specifically, their notion of zero-order weights and portfolio growth flows.

In this paper we focus on *KF** applications to Latin American countries. First, we document that Latin American portfolio inflows converge strongly to *KF** over medium-run horizons. Second, we demonstrate that deviations from *KF** help anticipate sudden stops in the region. Third, we show that *KF** acts as an indicator of vulnerability in the face of global shocks. Case studies of the Global Financial Crisis (GFC), post-GFC surge, and Covid-19 pandemic each indicate that, for Latin American countries, *KF** provides useful real-time information on the vulnerability of flows (well beyond that of alternative statistical proxies). Last, we analyze the drivers of short-run deviations in flows from *KF** and document interesting heterogeneity: flows to Brazil, Chile, and Mexico appear closely linked to commodity prices, while flows to Argentina, Costa Rica, and Peru are linked to global risk tolerance.

In Section 1 we present a brief introduction to *KF**. Section 2 documents the tendency of Latin American flows to revert to *KF** over the medium run. BWW (2022) demonstrated, for a large sample of countries, the usefulness of *KF** as an indicator for sudden stops and vulnerability to large global shocks; in Section 3 we show that it also helps predict stops and vulnerability in Latin American countries. Section 4, following analysis in BWW (2018) using annual data, analyzes factors associated with quarterly deviations from *KF**. Section 5 concludes.

1. *KF**

In this section we briefly present *KF**, the natural level of capital flows.[1]

1.1 The Theory behind *KF**

The construction of *KF** is motivated by the open-economy DSGE models with portfolio choice of Tille and van Wincoop (2010).[2] The model leads to two types of flows. *Portfolio growth flows* are simply the gross flows that would occur if new funds are allocated according

1. For more details, see BWW (2022).
2. See also Devereux and Sutherland (2011).

to zero-order portfolio weights. A positive productivity shock leads to increased savings that are deployed mostly at home (there is a portfolio home bias) but also abroad. If the productivity shock is persistent, these so-called portfolio growth flows are also persistent. The other type of flows—*reallocation flows*—is due to time variation in expected returns and risk. Time variation in expected returns impacts cross-border flows only through the effect of savings, as new home savings are invested mainly at home, thus pushing up home asset prices and requiring a decrease in expected returns (and, thus, capital outflows) to clear the asset markets. Time variation in second moments (risk) impacts optimal portfolio weights through changes in two hedge components: the covariance between excess returns (of home relative to foreign equities) and the real exchange rate and the covariance between excess returns and future expected portfolio returns. It is the change in these covariances that generates reallocation flows so, after a potentially large initial shock, the impact on reallocation flows quickly dissipates as future changes become a function of the persistent portfolio growth flows. In sum, zero-order portfolio growth flows—essentially the flows that would occur when the volatility of shocks becomes arbitrarily small—are persistent, owing to the persistence of underlying real-side shocks and hence savings. Reallocation flows can be substantial (and volatile) but, arising primarily from time variation in second moments, transitory.

1.2 Construction of *KF**

*KF** is based on the portfolio growth component of flows. The notion of portfolio growth flows is intuitively appealing, as the flow of new savings is precisely the amount of new funds available for foreign (or domestic) investment. Put another way, new savings are an important source of funds that would be potentially invested, some at home and some abroad. Portfolio growth flows are simply the gross flows that would occur if those new funds are allocated according to zero-order portfolio weights. Accordingly, the natural level of portfolio inflows at time t for a destination country d is

$$KF^{*}_{d,t} = \frac{1}{5} \sum_{i=1}^{5} 1\ \omega_{ROW,d,t-i}\, S_{ROW,t} \qquad (1)$$

where $\omega_{ROW,d,t\text{-}i}$ is the lagged weight of destination country d in rest-of-the-world (ROW) portfolios, defined as ROW holdings of country d

bonds and equities divided by *ROW* financial wealth, and $S_{ROW,t}$ is the contemporaneous flow of *ROW* private savings. Portfolio weights in equation (1) are formed by using Lane and Milesi-Ferretti (2018) data on *ROW* holdings of the destination country's equities and bonds (in balance-of-payments terms, the country's portfolio equity and portfolio debt liabilities), available annually for almost 200 countries starting in roughly 1995 and scaling these investment positions by *ROW* wealth (Davies and others, 2018). Savings, from the IMF's World Economic Outlook (WEO) dataset, is private savings (that is, national savings minus fiscal savings or "General government net lending/borrowing" in the IMF's WEO terms). *ROW* savings is world savings minus the recipient country's savings, and *ROW* wealth is world wealth minus the recipient country's wealth. Throughout, our *ROW* savings and weights (and flows) are 'ex-China' because, over the past two decades, there has been a substantial disconnect between China's savings (sizeable) and its outward portfolio investment (miniscule).[3]

As indicated in equation (1), we operationalize zero-order portfolio weights as a trailing five-year moving average of past portfolio weights. This ad-hoc decision is one that we are comfortable with for a number of reasons.[4] We employ a smoothed portfolio weight that abstracts from volatile transitory demand shocks. Filtering a weight has precedence in another measure, potential GDP: the Congressional Budget Office (CBO) applies a filter to the capital share so the volatility in the capital-share series does not create volatile estimates of potential GDP (Shackleton, 2018). Similarly, in our setting, asset price movements produce period-to-period volatility in portfolio weights; a filter dampens this volatility.[5]

We construct *KF** annually for the 2000 to 2021 period and form a quarterly version by linearly interpolating between year-end values. The number of countries for which we can form *KF** is limited primarily by Lane and Milesi-Ferretti (2018) data on portfolio liabilities; if a country has portfolio liabilities data, *KF** can be created even if the country does not publish flow data. Ninety-one countries

3. See BWW (2022) for details.

4. An alternative of using a theory such as CAPM to construct zero-order portfolio weights is possible but runs into the practical limitation that there is a sizeable home bias in actual data. And modeling higher frequency fluctuations in portfolio weights as in Koijen and Yogo (2019, 2020) would run counter to our focus on the longer-run natural level of flows.

5. It turns out that the smoothing of portfolio weights has no material impact on the performance of *KF**.

have portfolio liabilities data starting in 1995; for these, we can form a five-year lagged rest-of-the-world portfolio weight ($\omega_{ROW,d,t}$) starting in 2000 (i.e., the average weight from 1995 through 1999). For 90 other countries, we can form *KF** beginning later. In all, we create *KF** for 181 countries. This paper provides analysis for Latin American economies.

1.3 *KF** and its Decomposition

By definition,[6] trends in *KF** are given by *ROW* private savings, which is largely common to all destinations, and foreigners' (lagged) weights on stocks and bonds, which can vary substantially across investment destinations.

These components are presented in figure 1. The top left graph shows global (excluding China) private savings, which increased 8.5 percent per year over the 2005 to 2011 period and then was essentially flat from 2011 to 2018, increasing only 0.2 percent per year, before resuming strong increases. Thus, *ROW* savings tended to increase *KF** through 2011, maintained a level effect until 2018, and has since increased.

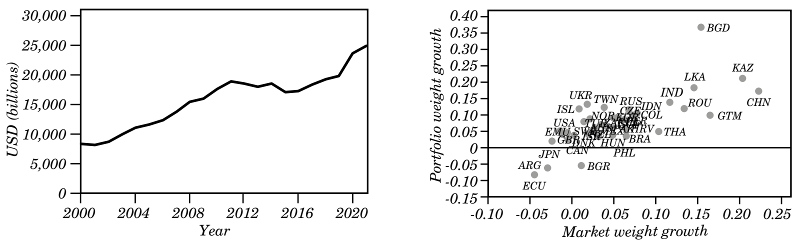
The other component is the portfolio weight (top right graph). The vertical height in the scatterplot shows the average annual increase in portfolio weight from 2000 to 2017. India's portfolio weight, for example, grew about 12 percent per year over that period, while Argentina's portfolio weight fell about five percent per year. While detailed analysis of factors behind countries' changing portfolio weights is beyond the scope of this paper, we note (and display in the scatterplot) that portfolio weights grew with market weights. For example, Peru's market weight increased at an annualized rate of 6.4 percent, while *ROW* investors increased their portfolio weight on Peru by 6.8 percent annually. The coefficient in a simple bivariate regression associated with the scatterplot is 0.85 with R^2 of 0.53.[7]

6. See equation 1.

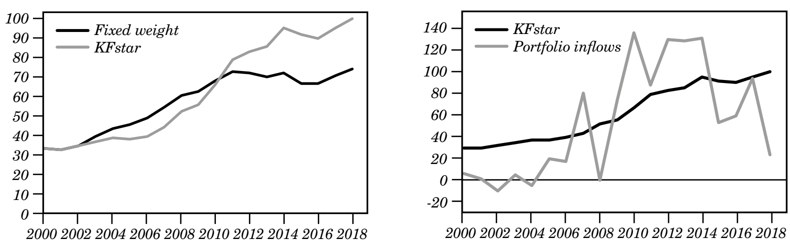
7. A cross-sectional regression (not shown) of the annualized growth rate in *ROW* portfolio weights over 2000–2017 on the average annual growth rate in the market weight (calculated as the sum of country *i*'s equity and bond-market relative to global-market capitalization), the average annual growth rate in country *i*'s share of global GDP, and the 2000–2017 change in the country's financial openness—by using the Chinn Ito (2006) KA measure—indicates that the most powerful explanatory variable is market weight. There is also a positive and statistically significant constant term that represents the broad-based trend toward financial globalization (or reduced home bias), suggesting that, independent of country-specific factors, average *ROW* portfolio weights increased by 2.4% annually.

Figure 1. *KF and Its Decomposition**

(*a*) *Global Private Savings (ex-China)* (*b*)



(*c*) *Latin America and Caribbean* (*d*) *Latin America*



Source: Authors' calculations.
Note: Shown in the top row are two components of *KF**: global private savings (ex-China) and the change in portfolio weights. The scatterplot (upper right) shows the relationship between the growth in portfolio weights and market weights (both expressed as average annual growth rate from 2000–2017). The bottom left graph shows, in billions of U.S. dollars, Latin America's *KF** and *KF** with weights fixed at the 2000 level. The difference between the two lines is a visual representation of the effect of increased weights on *KF**. The bottom right graph shows Latin America *KF** and actual portfolio inflows.

There is a great deal of endogeneity in a regression of portfolio weight on market weight, as inflows can enable market growth, so we view the relationship as illustrative not causal. Nonetheless, one might interpret this result as evidence of 'relative' (as in 'relative' PPP) international capital asset pricing model or international CAPM. We know that portfolio weights are far below market weights (i.e., home bias leads to failure of absolute international CAPM), but the growth rates in portfolio and market weights are highly correlated in the long run. Overall, we conclude that there are persistent movements in portfolio weights that appear to be associated with long-run changes in relative market size and reduced home bias.

The bottom left graph displays annual *KF** and a "fixed-weight" version of *KF**, which together provide a visual depiction of the relative contribution of global savings and changing portfolio weights. *KF** increases through 2011 and especially in the period starting in 2005, in part because of the strong growth in global (excluding China) private savings. The difference between *KF** and the fixed-weight version is a representation of the effect of increased portfolio weights on *KF**. For most regions around the world, portfolio weights increased, but here Latin America is an outlier because its portfolio weight actually decreased in the 2000s, driven mainly by a sharp decrease in Argentina's.

2. *KF** and Medium-Run Capital-Flow Forecasts

The volatility of international capital flows makes it difficult to discern what level of flows will likely persist going forward. figure 2 provides plots of quarterly portfolio flows and *KF** for eight Latin American economies. Visual inspection of the time-series plots in figure 2 reveals that quarterly portfolio flows are extremely volatile, yet they tend to oscillate around *KF** over time. Currently, *KF** suggests that each quarter Argentina, Chile, and Colombia should receive about $2 billion in portfolio inflows, whereas Mexico and Brazil should receive about $10 billion in quarterly inflows.

It is important to note that *KF** in figure 2 is *not* a statistical filter of flows but rather formed by projecting global savings based on historical portfolio weights.[8] A formal statistical assessment of *KF**, following Cogley (2002), is provided by analyzing whether current deviations of flows from *KF** help predict future changes in portfolio flows over the medium run. Focusing on the Latin American region, country-level regressions of the following form are estimated:

$$flows_{i,t+6} - flows_{i,t} = \alpha_i + \beta_i\,(flows_{i,t} - KF^{*}_{i,t}) + \varepsilon_{i,t}. \qquad (2)$$

Equation (2) provides a test of whether flows revert to *KF** over a six-quarter horizon. If current flows are above (below) *KF**, then we expect a future decline (increase) in flows. If flows revert precisely to *KF**, we expect to obtain estimates of $\beta_i = -1$. Note that this analysis is out of sample, as it uses the period t gap between actual flows and the predetermined *KF** to predict the *six-quarter*-ahead change in flows; the closeness of β_i to -1 is in effect a summary measure of its performance.

8. See equation 1.

Figure 2. *KF and Gross Portfolio Inflows**
(2000.IV–2021.IV, billions of U.S. dollars)

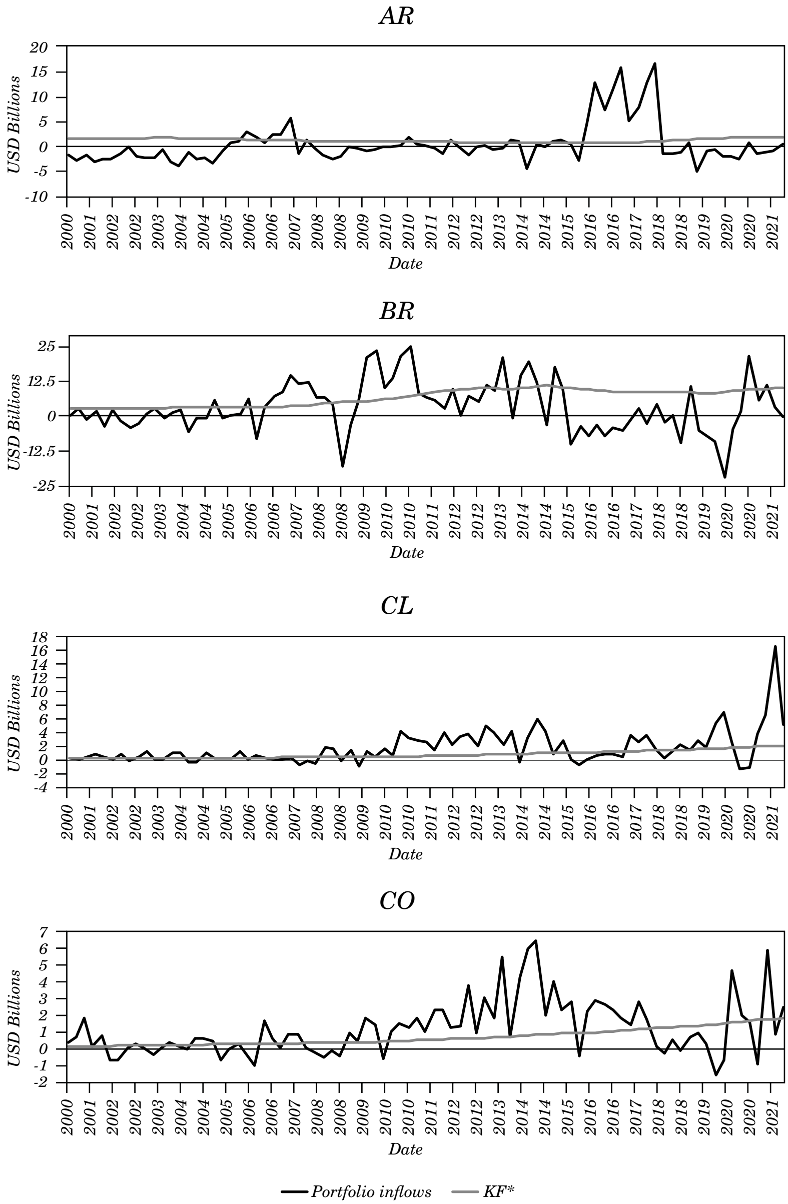
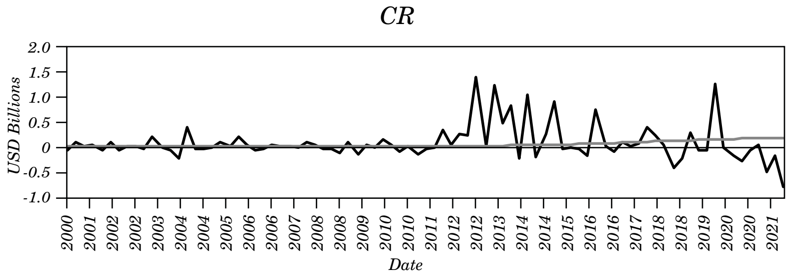
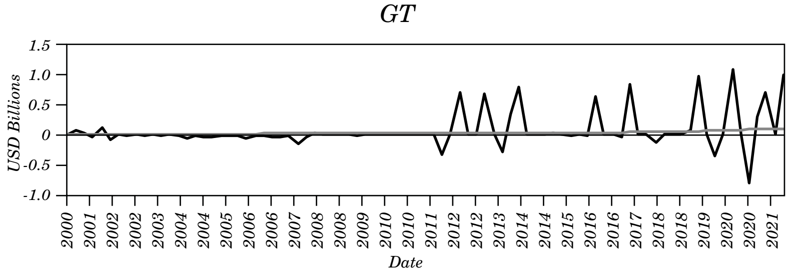
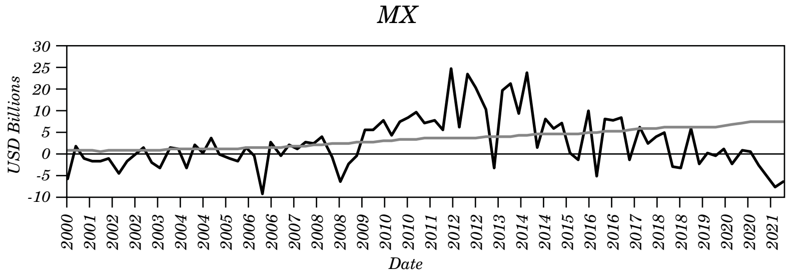
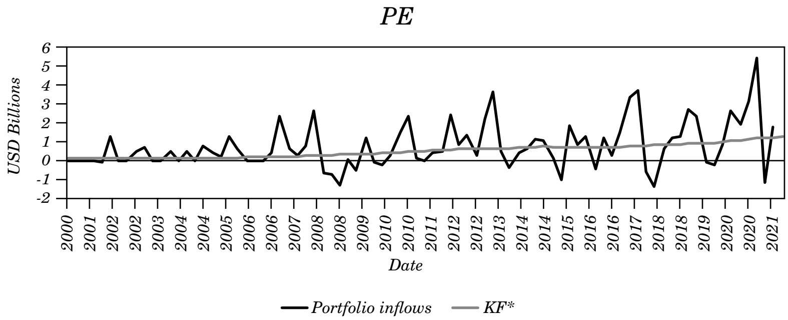


Figure 2. *KF and Gross Portfolio Inflows**
(2000.IV–2021.IV, billions of U.S. dollars) (*continued*)









Source: Authors' calculations.
Note: The graphs show quarterly portfolio inflows (the volatile line) and *KF** (the slow-moving line).

Table 1 provides results from the estimation of equation (2) for the eight Latin American economies with sufficient time-series data for quarterly portfolio flows. The simple regressions presented in Table 1, which forecast the *future* six-quarter change in portfolio flows as a function of the *current* gap in flows from *KF**, generate impressive levels of explanatory power (R^2). Remarkably, for most of the Latin American economies, the tendency of flows to revert to their natural level can explain 34–43 percent of medium-run variation. And for seven of the eight countries, we fail to reject a null hypothesis of β equal to –1, suggesting portfolio flows revert to *KF** over a six-quarter horizon. *KF** performs relatively poorly for Chile although the lack of fit appears driven primarily by recent volatility in flows including an outlier of $16.5 billion in inflows during 2021.III. In fact, truncating the sample at end-2020 greatly improves the fit for Chile (R^2 = 0.31) and yields a β estimate of –0.78.

Table 1. Reversion of Flows to *KF**

Country	*Beta (s.e.)*	R^2	*Observations*
Argentina	-0.788*** (0.159)	0.40	79
Brazil	-0.840*** (0.110)	0.39	79
Chile	-0.413 (0.317)	0.08	79
Chile*	-0.778*** (0.167)	0.31	75
Colombia	-0.748*** (0.160)	0.34	79
Costa Rica	-0.762*** (0.172)	0.34	79
Guatemala	-1.039*** (0.256)	0.43	79
Mexico	-0.503*** (0.131)	0.20	79
Peru	-0.819*** (0.129)	0.34	78

Source: Authors' calculations.
Note: The table presents Cogley regression results based on equation (2): $flows_{i,t+6} - flows_{i,t} = \alpha_i + \beta_i\,(flows_{i,t} - KF^*_{i,t}) + \varepsilon_{i,t}$. The proximity of β_i to –1 effectively summarizes the degree to which portfolio flows revert to *KF** in the medium run (in this case, over six quarters). Sample period for t+6 is 2002.II–2021.IV, except * which is truncated at 2020.IV. *** denotes significance at the 1% level.

3. Applications: *KF** as a Warning Indicator for Sudden Stops and Vulnerability to Large Global Shocks

In the previous section we demonstrated that, for Latin American countries, *KF** helps identify the component in portfolio flows that is expected to persist over medium-run horizons. BWW (2022) showed that, for a large sample of countries, *KF** helps forecast sudden stops and flows during large global shocks. This section assesses whether those results apply to Latin America economies.

3.1 Predicting Sudden Stops

We test whether portfolio flows that are well above *KF** predict an upcoming sharp decline in flows, focusing on the Forbes and Warnock (2012, 2021) extreme capital-flow episodes updated through 2021.IV. That is, does *KF*gap*—the gap between actual flows and *KF**—help predict future sudden stops in Latin America? Following BWW (2022), we estimate models of the form:

$$Prob\ (STOP_{i,t+h} = 1) =$$
$$F\ (KF^{*}gap_{i,t}, Global\ Factors_{t}, Local\ Factors_{i,t}) \qquad (3)$$

where $STOP_{i,t+h}$ is an indicator variable that takes the value of *1* if country *i* is experiencing a sudden stop in capital flows at time *t+h,* and $KF^{*}gap_{i,t}$ is the gap (scaled by GDP) between current flows and *KF**, averaged over the last four quarters. Everything is as in Forbes and Warnock (2021) with three exceptions: our forecast horizon is medium term, whereas Forbes and Warnock (2021) focus on one-quarter-ahead episodes; we extend the dataset through 2021.IV; and we include *KF*gap*. Global factors include global risk (measured as year-over-year change in the volatility index, VIX), global liquidity (measured as the year-over-year percentage growth in the 'global' broad money supply, where global is the sum for the Eurozone, U.S., U.K., and Japan), global monetary policy (measured as the year-over-year change in the average shadow short rate for the U.S., U.K., Eurozone, and Japan), global growth (measured as year-over-year global GDP growth from the IMF's WEO dataset), and the year-over-year percentage change in oil prices. Local factors are, as in Forbes and Warnock (2021), limited to local year-over-year real GDP growth and a regional contagion measure (an indicator equal to one if another country in the region has an episode). Because extreme capital-flow episodes are rare, following Forbes and Warnock (2021), we estimate equation (3) by using the complementary logarithmic framework, which assumes F(·) is the cumulative distribution function of the extreme value distribution.

Results from panel estimation of equation (3) at a six-quarter forecast horizon are presented in table 2. Merging our *KF** dataset with the Forbes and Warnock (2021) capital-flow episodes leaves a sample of eight Latin American economies (same countries listed in Table 1) and 595 quarterly observations. The results in panel A, which are similar to but stronger than those in BWW (2022), indicate flows above *KF**, strong global growth, rising global risk, and rapid growth in the global money supply are each associated with an increased likelihood of a sudden stop in capital inflows in six quarters.

Table 2. *KF and Extreme Capital-Flow Episodes**

Panel A	*Prob (Stop)* t+6 *quarters*
*KF*gap*	52.196***
	(10.341)
Global Variables	
Global GDP Growth	0.483***
	(0.158)
Risk	0.079***
	(0.022)
Liquidity	0.080***
	(0.022)
Oil Prices	-0.008*
	(0.005)
Monetary Policy	0.141
	(0.255)
Local and Contagion Variables	
Local GDP Growth	5.716
	(6.184)
Observations	595
Countries	8

Table 2. *KF and Extreme Capital-Flow Episodes (continued)**

Panel B	*Prob (Stop) t+6 quarters*
*KF*gap / GDP* = 0%	6.7%
*KF*gap/GDP* = 2.4%	21.9%
*KF*gap/GDP* = 4.8%	56.6%
*KF*gap* = 2.4% &	
Global growth = 4.2%	40.6%

Source: Authors' calculations.
Note: Panel A presents regressions forecasting *t*+6 sudden stops for the sample (*t*) 2001.IV–2019.IV. Explanatory variables include period *t KF*gap* (the deviation of actual flows from *KF**, expressed as a share of GDP) and global and local variables. Global variables include global GDP growth (year-over-year), risk (measured as the change in the VIX), liquidity (measured as the year-over-year percentage growth in the 'global' broad money supply, where global is the sum for the U.S., U.K., Eurozone, and Japan), monetary policy (measured as the year-over-year change in the average shadow short rate for the U.S., U.K., Eurozone, and Japan), and the year-over-year percentage change in oil prices. Local factors are year-over-year real GDP growth and a regional contagion measure (an indicator equal to one if another country in the region has an episode). Panel B shows, by using marginal effects from those regressions, the probability of a period *t*+6 sudden stop when (i) *KF*gap* is at its mean (0%) and one and two standard deviations above its mean (2.4% and 4.8%), holding all other variables at their means, and (ii) both *KF*gap* and global GDP growth are one standard deviations above their means.

To get a sense for economic magnitudes we calculate the model's estimated probability of a future stop when *KF*gap* is at its mean (zero) and one and two standard deviations above its mean (2.4 percent and 4.8 percent of GDP), holding all other variables at their means. When *KF*gap* is zero—that is, current flows are equal to *KF**—there is a 6.7 percent probability of experiencing a stop episode six quarters in the future. But when *KF*gap* is one or two standard deviations above its mean, the probability of a stop increases to 21.9 percent and 56.6 percent, respectively. We also find evidence that the combination of strong global growth and a large positive *KF*gap* is a particularly powerful predictor of a coming sudden stop: When *KF*gap* and global growth are both one standard deviation above their respective mean, the probability of a future stop climbs to 40.6 percent. These probabilities are similar to but slightly higher than those reported in BWW (2022) for a broader set of countries.

As in BWW (2022), the story that emerges is similar to the 'gap' analysis that the Bank for International Settlements (BIS) uses to predict banking crises.[9] For example, the BIS uses two 'gaps' as predictors, each defined as an underlying—corporate debt-to-GDP

9. See Aldasoro and others (2018).

or debt-service ratio—growing faster than trend, where trend for the BIS credit gap is estimated by an HP-filter and for the debt-service ratio is a 20-year moving average. The BIS indicators are not based on whether debt levels or debt servicing burdens are high, but whether they are growing faster than in the past. A similar 'gaps' analysis is at work with predicting sudden stops. When *KF** is growing (because global growth and hence global savings are growing) and actual flows are growing even faster (i.e., both global growth and *KF*gap* are above their sample means), a sudden stop is likely in six quarters. One difference from the BIS indicators: Our 'trend' is not a mechanical trend but *KF**.

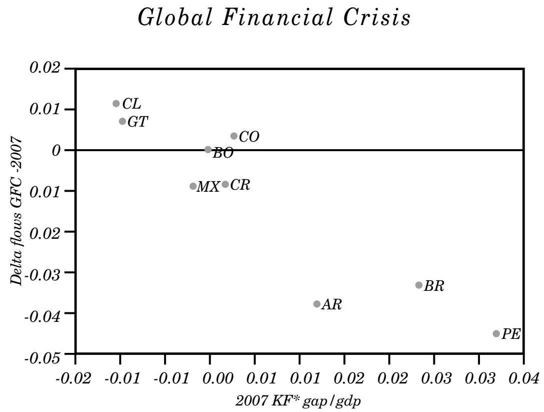
3.2 Vulnerability to Large Global Shocks

The results in Section 3.1 suggest that for Latin America *KF** can be used as a warning signal for future sudden stops. In this section we take a deeper dive to determine how deviations of flows from *KF** provide an indicator of the region's vulnerability to global shocks.

If *KF** represents the natural level of portfolio flows, Latin American countries receiving flows well above *KF** are most likely to experience a sharp reduction in flows in response to an external global shock. For analysis of the GFC, for each country we calculate the average *KF*gap/GDP* over the four quarters of 2007 as a measure of pre-GFC vulnerability. We then calculate the GFC impact of the crisis on flows as average *flows / GDP* during the GFC period (2008.IV—2009.III) minus *flows / GDP* during 2007. Figure 3 provides strong visual evidence in support of the hypothesis that countries with flows well above *KF** during 2007 (e.g., Peru, Brazil, and Argentina) subsequently suffered the largest (scaled by GDP) reductions in flows during the crisis. On the other end of the spectrum, we note that Chile and Guatemala were receiving flows *below KF** during 2007 and subsequently experienced increased flows during the GFC.

In the years following the GFC, capital flows to emerging markets rebounded strongly–especially for many Latin American economies. Figure 2 showed that the post-GFC rebound resulted in flows well in excess of *KF** for many Latin American economies (e.g., Brazil, Chile, Colombia, Costa Rica, and Mexico). A policymaker equipped with *KF** would have been concerned about flows well above equilibrium that were ripe for a reversal. And as it turns out, each of these markets experienced a subsequent sudden stop (2015 for most, 2013 for Chile).

Figure 3. *KF and Portfolio Flows during the GFC**

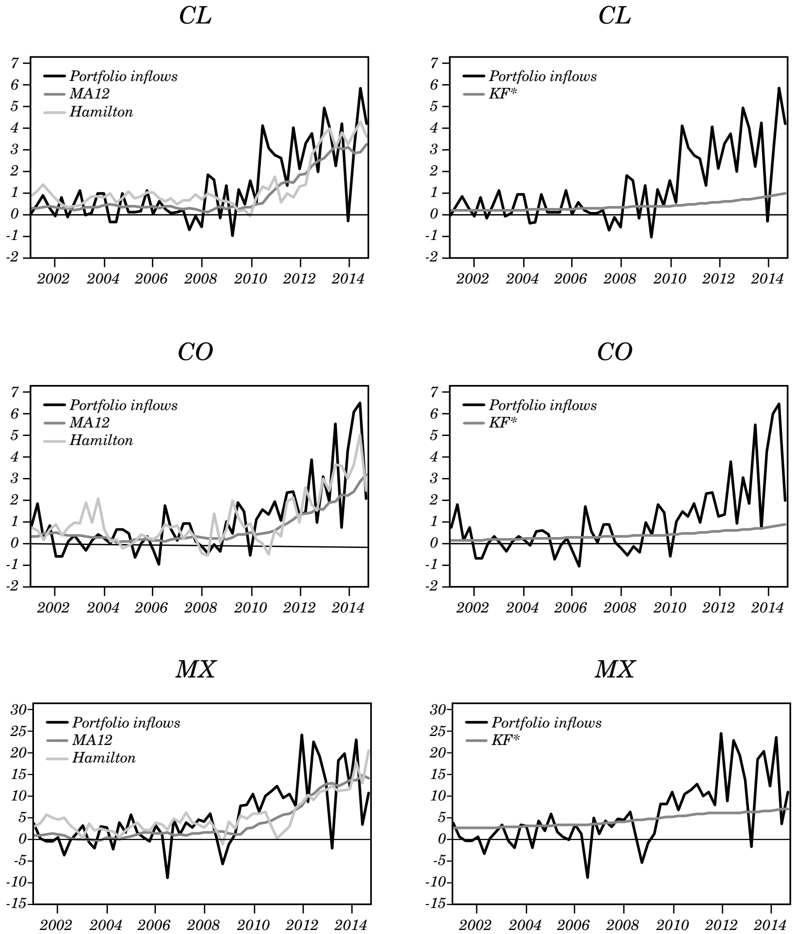


Source: Authors' calculations.
Note: The figure is a scatterplot of the relationship between the average 2007 *KF*gap/GDP* (the deviation of actual flows from *KF**, expressed as a share of GDP) and the subsequent change in portfolio flows during the GFC. The change in flows is calculated as average *flows / GDP* during the GFC period (2008.IV–2009.III) minus *flows / GDP* during 2007.

Although in hindsight it might appear self-evident that Latin American portfolio inflows were unsustainably high in the post-GFC period, real-time analysis is far more challenging. To demonstrate the impressive real-time forecasting properties of *KF**, we compare *KF** with some *statistical* proxies for equilibrium flows. One simple proxy is a 12-quarter moving average of past flows as a proxy for the equilibrium level of flows. If flows surge above the recent past, one might be concerned about the likelihood of a reversal. The Hamilton (2018) linear projection provides a more sophisticated statistical estimate of trend flows that is the fitted values from an OLS regression of a variable at date t on a constant and the four most recent values as of date $t–h$.

The left panel of figure 4 provides plots of actual portfolio flows overlaid against these statistical proxies for the period ending 2014.IV. For Chile, Colombia, and Mexico, we note that a policymaker comparing actual flows to these statistical filters would not have received a clear real-time signal regarding the sustainability of portfolio flows, as flows were oscillating around these filters. By contrast, the plots in the right panel, which overlay actual flows relative to *KF**, give a clear signal: In all three countries, flows greatly exceed their natural level and were therefore susceptible to a shock.

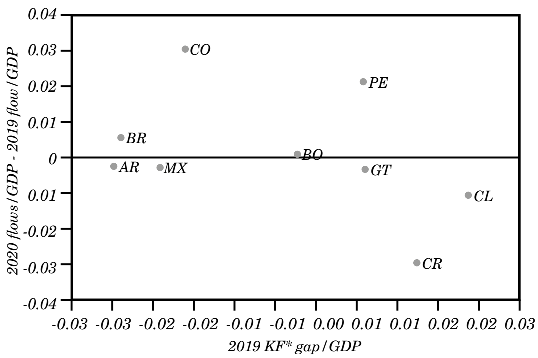
Figure 4. *KF** and Portfolio Flows post GFC



Source: Authors' calculations.
Note: The left panel compares quarterly portfolio flows to two statistical filters: (1) a 12-quarter moving average and (2) Hamilton (2018) linear projection estimated with data through 2014.IV. The right panel compares quarterly portfolio flows to *KF**. The sample for all graphs is 2000.IV–2014.IV.

As a final example of a global shock, we consider the Covid-19 pandemic. Figure 5 plots the relationship between the pre-pandemic (i.e., 2019) deviation of portfolio flows from *KF** (scaled by GDP) and the subsequent change in flows during 2020. Prior to the Covid-19 shock, only Chile and Costa Rica had flows significantly above *KF*,* while Argentina, Brazil, Colombia, and Mexico entered the pandemic with flows already well below *KF**. The Covid-19 shock induced a dramatic period of portfolio outflows—especially from emerging economies—but, consistent with the predictions of *KF*,* the period of outflows was short-lived, and the average change in flows from 2019 to 2020 was relatively small. Moreover, countries with the greatest outflows in 2020 had the largest (positive) *KF*gap* prior to the Covid-19 shock. In other words, *KF** provided a sense of which Latin American countries were most susceptible to outflows in 2020 and, as a guidepost, indicated that for all the countries in the region, the outflows episode should be neither long-lasting nor severe.

Figure 5. *KF and Portfolio Flows during the Pandemic**



Source: Authors' calculations.
Note: The figure is a scatterplot of the relationship between the 2019 (pre-pandemic) *KF*gap/GDP* (deviation of actual flows from *KF**, expressed as a share of GDP) and the subsequent change in portfolio flows during the 2020 pandemic (2020 *flows/GDP* – 2019 *flows/GDP*).

4. An Investigation into Deviations from *KF**

Thus far we have focused on the fact that deviations from *KF** are informative for future changes in flows, especially in the medium run. But these deviations can be sizable and occasionally sustained for significant periods, thus raising the question of what factors might drive flows to stray from their natural level. BWW (2018) conducted such analysis using annual panel data for 19 EMEs and found that higher than normal portfolio inflows occur when growth is strong, equity returns are high, and U.S. Treasury yields and risk measures (BBB-AAA spread or VIX) are low. Here we focus on Latin American countries and, noting that the drivers for deviations from *KF** likely differ by country, we analyze factors associated with the gap between actual and natural flows in country-level regressions. Explanatory variables include the VIX, long-term U.S. interest rates, commodity prices, and local and global GDP growth.

The results in table 3 highlight interesting heterogeneity across Latin American countries. In broad terms, deviations from *KF** are driven by commodity prices for Brazil, Chile, and Mexico—specifically, rising commodity prices are associated with a positive *KF*gap*. In contrast, for Argentina, Peru, and Costa Rica, risk measures are more important. For these countries, 'risk-off' episodes are associated with flows below *KF**.

Table 3. Analysis of deviations from *KF**

	AR	*BR*	*CL*	*CO*	*CR*	*GT*	*MX*	*PE*
Risk	-0.112**	-0.254	-0.012	-0.028	-0.007*	-0.005	-0.108	-0.028*
	(0.056)	(0.191)	(0.023)	(0.020)	(0.004)	(0.003)	(0.123)	(0.016)
U.S. rates	-1.188**	4.421***	-0.224	-0.188	-0.094	-0.042	0.992	-0.066
	(0.588)	(1.448)	(0.213)	(0.284)	(0.064)	(0.036)	(0.882)	(0.150)
Com. prices	0.002	0.044***	0.005**	0.004	-0.000	0.000	0.032***	-0.001
	(0.006)	(0.015)	(0.002)	(0.003)	(0.001)	(0.000)	(0.008)	(0.002)
Local growth	-0.577	102.186***	10.015	11.797	0.910	-1.124	90.317	4.783
	(6.158)	(38.282)	(11.187)	(11.377)	(3.082)	(1.372)	(69.491)	(3.800)
Global growth	0.186	-1.909*	-0.124	-0.221*	-0.023	-0.008	-1.445	-0.002
	(0.280)	(1.014)	(0.170)	(0.122)	(0.036)	(0.015)	(1.227)	(0.090)
R^2	0.20	0.39	0.23	0.16	0.10	0.08	0.21	0.06
N	77	77	77	76	77	72	77	77

Source: Authors' calculations.
Note: The table presents results from country-level regressions (2000.IV–2019.IV), where the dependent variable is *KF*gap* calculated as actual portfolio flows – *KF**. Global explanatory variables include global GDP growth (year-over-year), risk (measured as the change in the VIX), U.S. rates (10-yr Treasury yield), and the yearover-year percentage change in commodity prices. Year-over-year real GDP growth is included as a local factor.
* p<0.1; ** p<0.05; *** p<0.01

5. Conclusion

Latin American portfolio inflows show a strong tendency to revert to a natural level, *KF**, over medium-run horizons. Deviations of actual flows from *KF** provide significant predictive power for future flows – even in the face of large global shocks. Comparing current flows to *KF** provides policymakers with a real-time predictor of future sudden stops and vulnerability to external global shocks. Finally, analysis of short-run deviations of flows from *KF** reveals heterogeneous drivers: commodity prices for Brazil, Chile, and Mexico; risk tolerance for Argentina, Costa Rica, and Peru.

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How Important Is The Commodity Supercycle?

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World commodity prices are known to display long cycles. These cycles have a periodicity of 20 to 30 years and are called commodity-price supercycles. Figure 1 displays the time paths of eleven commodity prices deflated by the U.S. consumer price index over the period 1960 to 2018. All commodity prices appear to have long cycles in accordance with the supercycle hypothesis. In particular, commodity prices display two peaks post 1960—one in the early 1980s and one in the early 2010s. In the academic and financial-industry literature, the upswing

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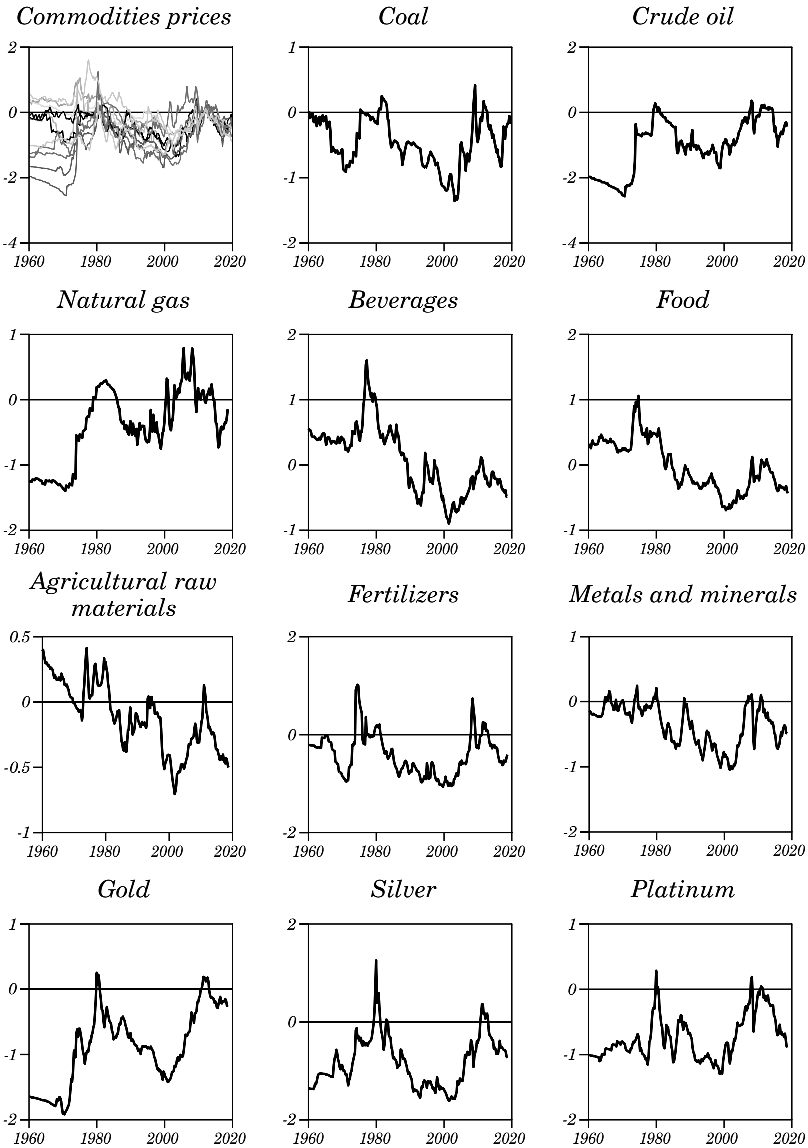
in commodity prices leading to the 1980s peak is typically attributed to the post-World War II reconstruction of Western Europe and Japan and to the cartelization of the crude oil market. The peak in the early 2010s is frequently attributed to the accession of China and other southeast Asian countries to world markets.

The existing literature on commodity-price supercycles has mainly focused on documenting their frequency, amplitude, and turning points. Less work has been devoted to estimating the importance of commodity-price supercycles for economic activity. The contribution of this paper is to identify global disturbances that cause regular cycles and supercycles in world commodity prices and to estimate the contribution of these global shocks to aggregate fluctuations in emerging and developed countries.

The econometric-oriented related literature typically uses spectral analysis to identify commodity-price supercycles. Cuddington and Jerrett (2008) pioneered the use of the asymmetric band-pass filter of Christiano and Fitzgerald (2003) to identify supercycles in commodity prices. They apply this technique to the prices of six metals traded on the London Metal Exchange. Subsequently, Erten and Ocampo (2013) apply this methodology to the identification of supercycles in real non-oil commodity prices.

The present paper proposes a different methodology to identify long cycles in world commodity prices. It identifies the commodity supercycle as a common permanent component in all commodity prices. The proposed common permanent component approach to identifying the commodity-price supercycle has two advantages relative to the spectral analysis approach. First, the spectral approach applies the band-pass filter to individual commodity-price time series separately. As a result, it delivers one supercycle per commodity price. However, the data strongly suggests that long cycles in commodity prices are correlated. The top left panel of figure 1, which plots all eleven commodity prices together, shows that long swings in commodity prices are synchronized. This correlation has been interpreted as reflecting the existence of a common driver. The common permanent component approach we propose delivers this common driver by identifying the nonstationary world shock responsible for the supercycle in all commodity prices.

Figure 1. Eleven Real Commodity Prices, 1960:I–2018:IV



Source: Authors' calculations.
Note: The frequency of the data is quarterly. The eleven commodity prices are deflated by the U.S. consumer price index and expressed in logs. All series are normalized to zero in 2010.IV.

A second desirable property of the common permanent component approach proposed in this paper for the identification of the commodity supercycle is that it allows for the joint estimation of transitory and permanent, domestic and world disturbances affecting aggregate activity in individual countries. As a result, the common permanent component approach provides a natural environment for estimating the contribution of the shocks responsible for the supercycle to explaining variations in output at the country level.

The paper formulates an empirical model that includes eleven commodity prices, the world interest rate, and the output of 24 (quarterly sample) or 41 (annual sample) small open developed and emerging economies. All commodity prices are assumed to be cointegrated with a common nonstationary world shock. In addition, commodity prices and the world interest rate are assumed to be buffeted by stationary world shocks. Output at the individual country level is driven by the nonstationary and stationary world shocks, a nonstationary country-specific shock, and a stationary country-specific shock. Thus, a constellation of stationary and nonstationary world and domestic shocks compete to explain movements in country-specific output. The nonstationary world shock is the one responsible for the commodity-price supercycle. Thus, ascertaining the role of the supercycle in accounting for output movements in a given country amounts to estimating the share of the nonstationary world shock in the variance decomposition of the country's output.

The model is cast in terms of deviations of endogenous variables from their respective stochastic trends and exogenous shocks. Since the exogenous shocks and the stochastic trends are unobservable, the variables in the model are latent variables. The estimation exploits the fact that the model delivers precise predictions for variables that are observed. In particular, the observable variables used in the estimation of the model are the growth rates of the eleven commodity prices in figure 1, the level of the world interest rate, and the growth rates of output of the countries included in the sample. The likelihood of the data is computed by using the Kalman filter, and the econometric estimation employs Bayesian techniques. The model is estimated on quarterly data covering the period 1960 to 2018. For countries for which quarterly output data since 1960 is not available, the model is estimated on annual data.

The paper delivers three main results. First, the common permanent component plays an important role in explaining movements in commodity prices at frequencies typically associated with the supercycle.

Specifically, the common permanent component explains on average across commodities between 67 and 91 percent of the forecast error variance of commodity prices at horizons between five and thirty years.

Second, world shocks that drive commodity prices and the world interest rate are major drivers of aggregate fluctuations in developed and emerging small open economies. Jointly the stationary and nonstationary world shocks explain more than half of the variance of output growth on average across countries. Third, and more importantly, the bulk (more than two thirds) of the explanatory power of world shocks stems from stationary shocks. These results obtain not only unconditionally but also conditionally on time horizons. Even at forecasting horizons typically associated with the supercycle (20 years or longer), stationary world shocks play a larger role than the nonstationary world shock in explaining the forecast error variance of the level of output in individual countries. Taken together, these results suggest that the commodity-price supercycle matters for explaining aggregate activity at the country level, but that its contribution is smaller than that of stationary world shocks.

This paper is related to several strands of literature. The empirical model follows recent studies that identify permanent disturbances and their contribution to business cycles (Uribe, 2018). The main results speak to a body of work on the role of world prices as mediators of world shocks for economic outcomes in small open economies. Mendoza (1995) and Kose (2002), by using calibrated real business-cycle models fed with estimated stochastic processes for the terms of trade, find that disturbances to this international price account for more than thirty percent of fluctuations in aggregate activity. More recently, Miyamoto and Nguyen (2017), and Drechsel and Tenreyro (2018) find similar results by using a Bayesian estimation approach. Schmitt-Grohé and Uribe (2018) apply a more agnostic approach based on structural vector autoregressions and find that the contribution of terms-of-trade shocks to explaining aggregate fluctuations in poor and emerging economies is only ten percent. These authors argue for the need to consider more disaggregated measures of world prices to better capture the transmission of world shocks to individual economies. Fernández and others (2017) employ a similar empirical strategy as Schmitt-Grohé and Uribe (2018) but expand the set of world prices from one to four, three commodity prices, and the world interest rate, and find that world shocks mediated by this set of prices explain one third of output fluctuations on average in a set of 138 countries over the period 1960 to 2015. This figure more than doubles when the estimation is conducted on a more recent sample beginning in the late 1990s, as shown by Shousha (2016), Fernández and others (2018), and Fernández and others (2017). In the papers just cited, shocks to commodity prices, if explicitly modeled, are assumed to be stationary, and as a result, this body of work does not speak directly to the importance of commodity-price supercycles.

As mentioned earlier, key references on the use of spectral analysis for the estimation of commodity-price supercycles are Cuddington and Jerrett (2008) and Erten and Ocampo (2013). These papers and the early work by Pindyck and Rotemberg (1990) speculate on the existence of common drivers of commodity prices, which serves as motivation for the common permanent component approach proposed in the present paper. Alquist and others (2020) apply a factor-based identification strategy to estimate the role of commodity prices in explaining the global economic activity. Benguria and others (2018) identify the commodity-price supercycle by HP filtering and analyze its transmission by using firm-level administrative data from Brazil.

The paper also contributes to a literature assessing the role of transitory and permanent shocks in driving business cycles in developed and emerging economies.[1] It finds that, for both developed and emerging countries, transitory shocks play a larger role than permanent shocks, even if one conditions on world shocks or on country-specific shocks.

Finally, the structuralist literature pioneered by Prebisch (1950) and Singer (1950) argues that secular deterioration in the terms of trade plays a central role in the development of emerging countries. The empirical relevance of this hypothesis has been the subject of debate. One reason is that, as figure 1 suggests, it is not clear that overall commodity prices display secular deterioration. In the context of that literature, deterioration of the terms of trade is understood as primary commodity prices growing on average at a slower pace than prices of nonprimary goods. Our commodity-price data is deflated by the U.S. CPI index, and as such secular deterioration should manifest as a downward trend in raw data. Over the sample period considered in the figure, some commodity prices display a downward trend (e.g., agricultural raw materials) but others display an upward trend (e.g., energy and precious metals). The present analysis does not aim to explain the effect of trending real commodity prices on economic

1. See, among others, Aguiar and Gopinath (2007), García-Cicco and others (2010), Chang and Fernández (2013), and Miyamoto and Nguyen (2017).

development. However, our analysis does have a point of contact with the Prebisch-Singer hypothesis in that it allows for innovations in the permanent component of real commodity prices to have an effect on output growth at the country level.

The remainder of the paper is organized into seven sections. Section 1 presents the empirical model. Section 2 introduces the observables, the priors, and the estimation strategy. Section 3 presents the definitions and sources of the quarterly data on commodity prices, world interest rates, and output in 24 predominantly developed economies spanning the period 1960.I to 2018.I. Section 4 analyzes the estimated commodity-price supercycle. Section 5 presents variance decompositions, forecast error variance decompositions, and impulse response analysis to ascertain the importance of the commodity supercycle for aggregate activity in the small open economies considered. Section 6 estimates the model on annual data from 1960 to 2018 for 24 emerging and 17 developed countries.[2] Finally, section 7 concludes.

1. An Empirical Model of the Commodity Supercycle

The empirical model consists of a world block and a country-specific block. The world block describes the evolution of the vector p_t containing eleven real commodity prices and the gross real interest rate in quarter t, all expressed in logarithms. The commodity supercycle is modeled as a nonstationary exogenous variable X_t^p with the property of being cointegrated with the eleven commodity prices. We can then define a vector of transformed world prices, denoted $\hat{p}_t$, that is stationary as follows:[3]

$$\hat{p}_t = \begin{bmatrix} \hat{p}_t^1 \\ \hat{p}_t^2 \\ \vdots \\ \hat{p}_t^{11} \\ \hat{r}_t \end{bmatrix} \equiv \begin{bmatrix} p_t^1 - X_t^p \\ p_t^2 - X_t^p \\ \vdots \\ p_t^{11} - X_t^p \\ r_t \end{bmatrix}$$

2. For further technicals details see Appendix.

3. For expositional purposes, constant terms are omitted. The model with constant terms is presented in the Appendix.

The identification assumption that all commodity prices have the same cointegrating vector with X_t^p is based on the observation that, in the raw data, commodity prices do not seem to diverge from one another over time (see figure 1, upper left panel).

The vector $\hat{p}_t$ is assumed to be buffeted by a nonstationary shock, given by variations in the growth rate of the permanent component of world prices, $\Delta X_t^p \equiv X_t^p - X_{t-1}^p$, and 12 stationary world shocks denoted z_t^p. The vector of world prices evolves according to the following autoregressive process:

$$\hat{p}_t = \sum_{i=1}^{4} B_{pp}^i \hat{p}_{t-i} + C_{pX^p} \Delta X_t^p + C_{pz^p} z_t^p, \tag{1}$$

where B_{pp}^i for i = 1, 2, 3, 4, C_{pX^p}, and C_{pz^p} are matrices of coefficients of order 12-by-12, 12-by-1, and 12-by-12, respectively. Without loss of generality, assume that C_{pz^p} is lower triangular with ones on the diagonal. This is not an identification restriction. The elements of z_t^p should be interpreted as combinations of stationary world shocks affecting commodity prices and the interest rate. The present study does not aim to identify these shocks individually, but rather to ascertain their joint contribution to explaining movements in world prices and aggregate activity and to compare it to that of the nonstationary world shock X_t^p driving the commodity supercycle.

The domestic block consists of the vector y_t containing real output for 24 small open economies expressed in logarithms and denoted y_t^i, $i = 1,\ldots,24$. In each country, output is assumed to be cointegrated with a linear combination of a country-specific nonstationary shock, denoted X_t^i for $i = 1,\ldots,24$ and the nonstationary component of real-world commodity prices, X_t^p. The rationale behind the assumption that X_t^p enters in the cointegrating relationship of output is that, in models of small open economies with commodity prices, output inherits their stochastic properties. We note that this long-run relationship between output and the nonstationary component of world shocks is not subject to the observation made by Kehoe and Ruhl (2008) that, depending on how real GDP is measured in the data, terms-of-trade shocks may not act like technology shocks, for their observation has to do with the direct effect of terms-of-trade shocks on measured GDP and not with their indirect effect on quantities. The cointegration relationship between y_t^i, X_t^i, and X_t^p is estimated, thus allowing the data to choose the strength of the long-run link between output and commodity prices in each country.

Let $\hat{y}_t$ be a 24-by-1 vector of deviations of output from trend. Then,

$$\hat{y}_t = \begin{bmatrix} \hat{y}_t^1 \\ \hat{y}_t^2 \\ \vdots \\ \hat{y}_t^{24} \end{bmatrix} \equiv \begin{bmatrix} y_t^1 - X_t^1 - \alpha^1 X_t^p \\ y_t^2 - X_t^2 - \alpha^2 X_t^p \\ \vdots \\ y_t^{24} - X_t^{24} - \alpha^{24} X_t^p \end{bmatrix}.$$

For each country i the country-specific shocks consist of the growth rate of the permanent component of output, ΔX_t^i, and a stationary shock, z_t^i. Detrended output is assumed to evolve according to the following autoregressive process:

$$\hat{y}_t = \sum_{i=1}^{4} B_{yp}^i \hat{p}_{t-i} + \sum_{i=1}^{4} B_{yy}^i \hat{y}_{t-i} + C_{yX^p} \Delta X_t^p + C_{yz^p} z_t^p + C_{yX} \Delta X_t + z_t, \tag{2}$$

where $\Delta X_t = [\Delta X_t^1 \ldots \Delta X_t^{24}]'$ and $z_t = [z_t^1 \ldots z_t^{24}]'$. B_{yp}^i and B_{yy}^i, for $i = 1, \ldots, 4$, are 24-by-12 and 24-by-24 matrices of coefficients, respectively, and C_{yX^p}, C_{yz^p}, and C_{yX} are matrices of order 24-by-1, 24-by-12, and 24-by-24, respectively. Matrices B_{yy}^i and C_{yX} are assumed to be diagonal.

Note that the world shocks ΔX_t^p and z_t^p enter directly in the domestic block, as opposed to mediated by $\hat{p}_t$. This flexibility allows for the possibility that ΔX_t^p and z_t^p capture global and regional shocks affecting individual countries both directly and via world prices, such as exogenous global and regional productivity shocks.

The exogenous shocks, ΔX_t^p, ΔX_t, z_t^p, and z_t follow univariate autoregressive processes.

Specifically, let u_t denote the vector of exogenous shocks

$$u_t \equiv \begin{bmatrix} \Delta X_t^p \\ \Delta X_t \\ z_t^p \\ z_t \end{bmatrix}.$$

We assume that u_t obeys the law of motion

$$u_t = \rho u_{t-1} + \psi \nu_t, \tag{3}$$

where $\nu_t \sim i.i.d.N\ (0, I_{61})$. The matrices ρ and ψ are assumed to be diagonal. This implies that the permanent component of world prices, X_t^p, is uncorrelated with the stationary world shocks, z_t^p; that the permanent and transitory country-specific shocks are uncorrelated with each other and with other country-specific shocks; and that country-specific shocks, X_t^i and z_t^i, are uncorrelated with the world shocks, X_t^p and z_t^p. The latter assumption is motivated by the fact that the countries in the sample are small open economies and, as such, their idiosyncratic shocks do not affect world prices. We assume that the correlation of output across countries stems from world shocks. Accordingly, the matrices B_{yy}^i for i = 1,..., 4, as well as the matrix C_{yX}, are restricted to be diagonal. The assumption that the world shocks X_t^p and z_t^p (as opposed to the contemporaneous world prices $\hat{p}_t$) enter directly in the domestic block, equation (2), allows for the possibility that world shocks affect country-level output, both directly and indirectly, mediated by world prices. A direct effect of world shocks on country-level output could occur, for example, via productivity shocks that are correlated across countries.

2. Observables, Priors, and Estimation Strategy

All variables in the system (1), (2), and (3), except for the interest rate, r_t, are latent variables and therefore unobservable. As a result, the system cannot be directly estimated on data. However, we will exploit the fact that the model has precise predictions for variables that are observable. Specifically, the data used in the estimation includes the growth rates of the commodity prices, Δp_t^i for $i = 1,...,11$, the level of the world interest rate, r_t, and the growth rates of output, Δy_t^i for $i = 1,...,24$. The observable variables are related to the latent variables through the following identities:

$$\Delta p_t^i = \Delta \hat{p}_t^i + \Delta X_t^p; i = 1,\ldots,11, \tag{4}$$

$$\Delta y_t^i = \Delta \hat{y}_t^i + \Delta X_t^i + \alpha^i \Delta X_t^p; i = 1,\ldots,24, \tag{5}$$

and

$$r_t = \hat{r}_t. \tag{6}$$

The observable variables are assumed to be measured with error. Letting o_t denote the vector of observed variables, we have that

$$o_t = \begin{bmatrix} \Delta p_t^1 \\ \vdots \\ \Delta p_t^{11} \\ r_t \\ \Delta y_t^1 \\ \vdots \\ \Delta y_t^{24} \end{bmatrix} + \mu_t, \tag{7}$$

where μ_t is a 36-by-1 vector of measurement errors distributed i.i.d. $N(0, R)$ and R is a diagonal matrix. We restrict the measurement errors to explain no more than ten percent of the variance of the data.

The relationship between the observables and the latent variables, the fact that the model is linear, and that all innovations are Gaussian, make it possible to compute the likelihood of the data, which in turn allows for the estimation of the parameters of the model. To calculate the likelihood, it is convenient to express the model in state-space form. To this end, let

$$\hat{x}_t = \begin{bmatrix} \hat{p}_t & \hat{y}_t \end{bmatrix}' \text{ and } \xi_t = \begin{bmatrix} \hat{x}_t & \hat{x}_{t-1} & \hat{x}_{t-2} & \hat{x}_{t-3} & u_t \end{bmatrix}'.$$

Then the state-space representation of the model, equations (1)–(7), is given by

$$\xi_{t+1} = F\xi_t + P\nu_{t+1}, \tag{8}$$

and

$$o_t = H'\xi_t + \mu_t, \tag{9}$$

where the matrices F, P, and H are known functions of the matrices B_{pp}^i, B_{yp}^i, B_{yy}^i for $i = 1,...,4$, C_{pX^p}, C_{pz^p}, C_{yX^p}, C_{yz^p}, C_{yX}, ρ and ψ. The model is estimated with Bayesian techniques. Draws from the posterior distribution are obtained by applying the Metropolis-Hastings algorithm. We construct an MCMC chain of 2.5 million draws and discard the first 1.5 million.

The prior distributions of the estimated parameters are summarized in table 1. We impose normal prior distributions to all elements of B_{pp}^i, B_{yy}^i, and B_{yp}^i, for $i = 1,...,4$. In accordance with the

Minnesota prior, we assume that, at the mean of the prior parameter distribution, the elements of $\hat{x}_t$ follow univariate autoregressive processes. So, when evaluated at their prior mean, only the main diagonals of B^1_{pp} and B^1_{yy} take nonzero values, and all other elements of B^i_{pp} and B^i_{yy}, and all elements of B^i_{yp} for $i = 1,\ldots,4$ are nil. We impose an autoregressive coefficient of 0.95 in all equations so that all elements along the main diagonal of B^1_{pp} and B^1_{yy} take a prior mean of 0.95. We assign a prior standard deviation of 0.5 to these elements, which implies a coefficient of variation close to one half (0.5/0.95). Also, along the lines of the Minnesota prior, we impose lower prior standard deviations on all other estimated elements of the matrices B^i_{pp}, B^i_{yy}, and B^i_{yp} for $i = 1,\ldots,4$, and set them to 0.25.

All estimated elements of the matrices C_{pX^p}, C_{pz^p}, C_{yX^p}, C_{yz^p}, and C_{yX} are assumed to have normal prior distributions with mean zero and unit standard deviation, with one exception: the diagonal elements of the diagonal matrix C_{yX}, which govern the responses of $\hat{y}^i_t \equiv y^i_t - X^i_t - \alpha^i X^p_t$ to an innovation in ΔX^i_t for $i = 1,\ldots,24$, are assumed to have a prior mean of –1. This means that a shock that increases output in country i in the long run by one percentage point, under the prior, has a zero-impact effect. This prior is motivated by a strand of the business-cycle literature suggesting that the impact effect on output of a permanent productivity shock could have either sign depending on the strength of the wealth effect on labor supply.[4]

The diagonal elements of the diagonal matrix ψ representing the standard deviations of the innovations in the exogenous shocks are all assigned Gamma prior distributions with mean and standard deviations equal to one. We impose non-negative serial correlations on the exogenous shocks (the diagonal elements of ρ) and adopt Beta prior distributions for these parameters. We assume relatively small means of 0.3 for the prior mean of the serial correlations of the nonstationary shocks (ΔX^p_t and ΔX^i_t for $i = 1,\ldots,24$) and a relatively high mean of 0.7 for the prior mean of the serial correlations of the stationary shocks (z^p_t and z_t). The prior distributions of all serial correlations are assumed to have a standard deviation of 0.2. The variances of all measurement errors (the diagonal elements of the matrix R) are assumed to have a uniform prior distribution with lower bound zero and upper bound of ten percent of the sample variance of the corresponding observable indicator. Although not explicitly discussed thus far, the estimated

4. See, for example, Galí (1999).

model includes constants, which appear in the observation equation (9).[5] These constants represent the unconditional means of the 36 observables. They are assumed to have normal prior distributions with means equal to the sample means of the observables and standard deviations equal to their sample standard deviations divided by the square root of the sample length (231 quarters).

Table 1. Prior Distributions

Parameter	*Distribution*	*Mean*	*Std. Dev.*
Main diagonal elements of B^1_{pp} and B^1_{yy}	Normal	0.95	0.5
All other estimated elements of B^1_{pp}, B^1_{yy}, B^1_{yp}	Normal	0	0.25
Estimated elements of B^i_{pp}, B^i_{yy}, B^i_{yp} $i = 2,3,4$	Normal	0	0.25
Estimated elements of C_{pX^p} and C_{pz^p}	Normal	0	1
Diagonal of C_{yX}	Normal	–1	1
Elements of C_{yX^p} and C_{yz^p}	Normal	0	1
Diagonal of $\rho(1{:}25,1{:}25)$	Beta	0.3	0.2
All other diagonal elements of ρ	Beta	0.7	0.2
Diagonal of ψ	Gamma	1	1
α_i, $i = 1,\ldots,24$	Normal	0	1
Diagonal elements of R	Uniform $[0,\frac{var(o_t)}{10}]$	$\frac{var(o_t)}{(10*2)}$	$\frac{var(o_t)}{(10*\sqrt{12})}$
Elements of A	Normal	mean(o_t)	$\sqrt{\frac{var(o_t)}{T}}$

Source: Authors' calculations.
Notes: T denotes the sample length, which equals 231 quarters. The vector A denotes the mean of the vector o_t and is defined in the Appendix.

5. See the Appendix for details.

Posterior means and error bands around the impulse responses shown in later sections are constructed from a random subsample of the MCMC chain of length 100 thousand with replacement.

3. The Data

The model is estimated on quarterly data on eleven world commodity prices, the world interest rate, and the gross domestic product of 24 small open economies. The sample period is 1961.I to 2018.IV. The eleven commodity prices included in the estimation are beverages, food, agricultural raw materials, fertilizers, metal and minerals, gold, platinum, silver, coal, crude oil, and natural gas. The raw data is monthly and expressed in current U.S. dollars. The source is the World Bank's Commodity-Price database (the Pink Sheet) except for coal prices, which come from Global Financial Data (GFD). The GFD coal price is identical to the one in the Pink Sheet, except that it begins in 1960.1, whereas it begins only in 1970.1 in the Pink Sheet. Quarterly real commodity-price indices are constructed by first deflating the monthly nominal price indices by the monthly CPI index of the United States, then taking a simple average of the deflated values across the corresponding months in each quarter. The data are normalized by dividing by each series' 2010.IV observation. The World Bank publishes data on the prices of 40 individual commodities. The aggregation into eleven prices responds to the need to economize on degrees of freedom in the estimation. The individual commodities whose prices are aggregates are beverages, food, agricultural raw materials, fertilizers, and metal and minerals. The aggregate price indices are taken from the Pink Sheet. Thus the eleven commodity prices included capture information from all 40 commodity prices in the World Bank's Commodity-Price database.

The quarterly time series for the world real interest rate, r_t, is constructed as $1+r_t=(1+i_t)E_t\frac{1}{1+\pi_{t+1}}$, where i_t denotes the nominal interest rate on three-month U.S. Treasury bills, and $1+\pi_{t+1}=\frac{P_{t+1}}{P_t}$, denotes the gross growth rate of the consumer price index, P_t, as measured by the U.S. CPI index. The expected value of the inverse of gross inflation, $E_t\frac{1}{1+\pi_{t+1}}$ is approximated by the fitted component of an OLS regression of $\frac{1}{1+\pi_{t+1}}$ onto a constant, $\frac{1}{1+\pi_t}$ and $\frac{1}{1+\pi_{t-1}}$.

Output is measured by seasonally adjusted real gross domestic product from the quarterly national accounts of the OECD.[6]

For a country to be included in the sample, we require at least 50 years of quarterly observations of real output. The rationale behind this restriction is that identifying the real effects of the commodity supercycle requires observing the behavior of output over a relatively long period. In addition, since commodity prices and the world interest rate are assumed to be exogenous to the country, we exclude large economies. These selection criteria result in the following 24 countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Greece, Iceland, Ireland, Italy, Korea, Luxembourg, Mexico, Netherlands, New Zealand, Norway, Portugal, South Africa, Spain, Sweden, Switzerland, Turkey, and the United Kingdom. The panel includes only four emerging economies, Korea, Mexico, South Africa, and Turkey. Section 6 estimates the model on annual data, which allows for the inclusion of a larger number of emerging economies.

4. The Commodity-Price Supercycle

Figure 2 displays the estimated common permanent component, X_t^p, of real commodity prices. It is constructed by Kalman smoothing at the posterior mean of the parameter estimate. The figure also displays the eleven observed commodity prices. We interpret the variable X_t^p as the commodity supercycle. The figure suggests that this interpretation is sensible as X_t^p appears to capture well the low-frequency comovement of the individual commodity prices. Over the period 1960 to 2018, commodity prices display two distinct supercycles—one peaking in 1980 and the other in 2008. The rapid growth in X_t^p between the early 1970s and 1980 coincides with the OPEC oil-price crises. As the market power of the oil cartel weakened in the 1980s and the supply of other countries (e.g., the United States and those located around the North Sea) rose, the downswing of the supercycle began. The expansionary phase of the second commodity-price supercycle begins around the time of China's accession to the WTO in 2001 and the peak is reached with the onset of the Global Financial Crisis of 2008. The prediction of two commodity supercycles

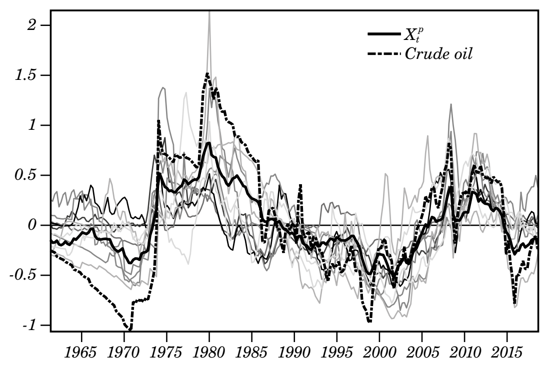
6. The OECD series name is VOBARSA. For Greece and Iceland, the data appears not to have been seasonally adjusted at the source. Therefore, these two series were adjusted by using the X-13 ARIMA-SEATS software produced, distributed, and maintained by the U.S. Census Bureau.

post 1960 and their dating is in line with the estimates reported in Erten and Ocampo (2013) using an asymmetric bandpass filtering approach on real non-oil commodity prices that picks out cycles with periodicity between 20 and 70 years.

The permanent component of commodity prices, X_t^p, plays a significant role in explaining movements in these variables. Table 2 displays the fraction of the variance of changes in commodity prices accounted for by changes in their permanent component. On average across prices, ΔX_t^p explains more than one fourth of the variance of changes in commodity prices. The variance shares are estimated with precision, with standard deviations equal to two percentage points on average. The permanent component plays the largest role in explaining movements in crude oil prices with a variance share of 60 percent.

Estimating the price block of the model separately from the output block (not shown) yields similar results for the time path of X_t^p. Also, this estimation approach yields a similar result for the average share of the variance of the growth rates of world prices explained by ΔX_t^p (21 percent when the price block is estimated separately versus 27 percent when it is estimated jointly with the output block). However, estimating the price block by using only information on prices yields a smaller role for the permanent component, ΔX_t^p, in explaining the variance of the growth rate of crude oil prices (35 versus 60 percent). This finding suggests that, even though the price block is independent of the output block, data on country-level output is informative for the estimation of the parameters governing the dynamics of world prices.

Figure 2. The Commodity-Price Supercycle



Source: Authors' calculations.
Notes: The permanent component of the eleven real commodity prices, X_t^p, is computed by Kalman smoothing using the posterior mean of the parameter estimates. The thin solid lines are the eleven observed real commodity prices (beverages, food, agricultural raw materials, fertilizers, metal and minerals, gold, platinum, silver, coal, crude oil, and natural gas). All-time series are constructed as cumulative demeaned growth rates.

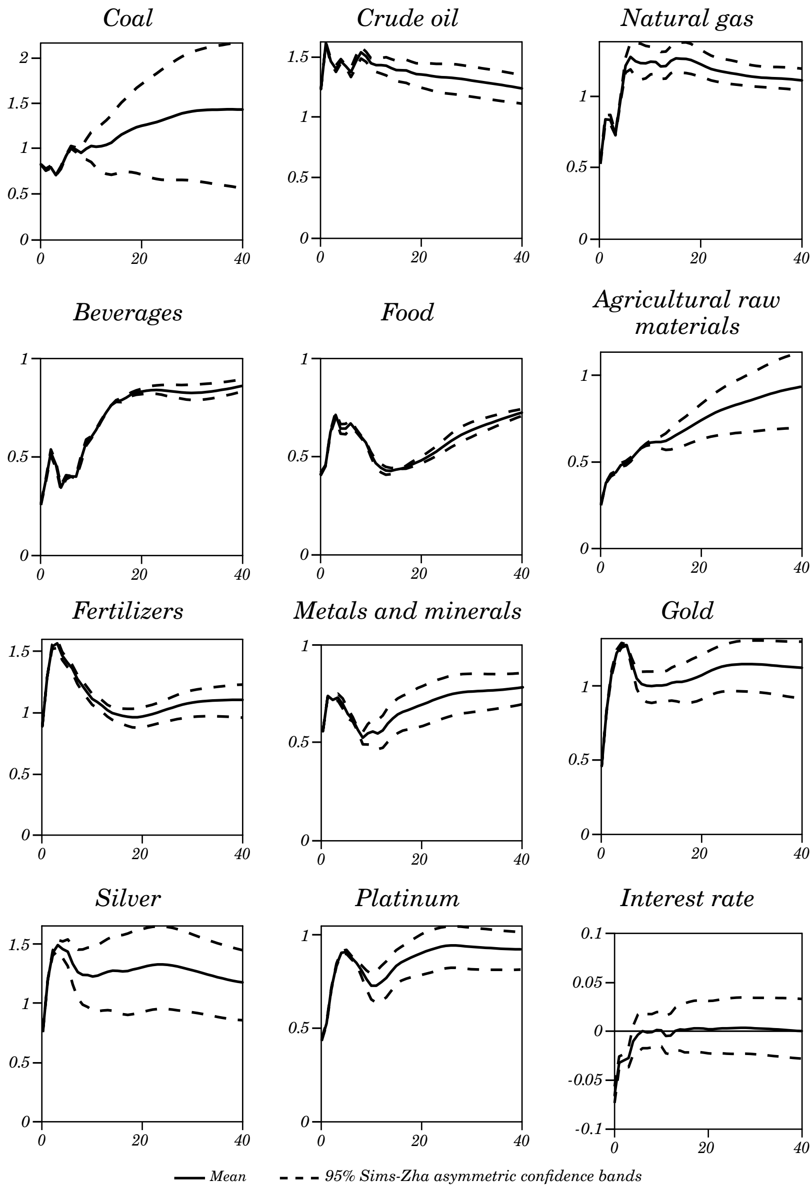
Table 2. Percent of Variance of the Growth Rate of Real Commodity Prices Explained by ΔX_t^p

Price of	*Mean*	*Std. Dev.*
Coal	26	3
Crude Oil	60	2
Natural Gas	20	3
Beverages	11	1
Food	21	2
Agr. Raw Materials	20	3
Fertilizers	33	2
Metal and Minerals	30	2
Gold	30	2
Silver	28	2
Platinum	19	1
Mean across prices	27	2
Median of prices	26	2
Real Rate	15	8

Source: Authors' calculations.
Note: The reported figures are based on 100,000 draws from the posterior distribution of the variance decomposition.

Figure 3 presents the impulse responses of the eleven real commodity prices and the world interest rate to a unit long-run increase in the permanent component X_t^p along with 95-percent asymmetric confidence bands, computed by using the methodology proposed by Sims and Zha (1999). For most commodity prices, a positive innovation in X_t^p has a positive but less than unity impact effect and induces a slow convergence to the permanently higher level, which by construction is equal to one. Exceptions are crude oil, which displays overshooting on impact and convergence from above, and natural gas, fertilizers, gold, and silver, which display delayed overshooting.

Figure 3. Impulse Responses of World Prices to a Long-Run Increase in X_t^p of Unity



Source: Authors' calculations.

Notably, an increase in the permanent component of commodity prices has a negative effect on the world interest rate. The estimated negative conditional comovement between commodity prices and interest rates is important for commodity exporters with external debt because it suggests that, when commodity prices increase, the country also benefits from favorable conditions in international financial markets. Similarly, during a downturn in commodity prices, the costs of external debt rise. This result is in line with the work of Shousha (2016), who finds that movements in the interest rate are in part driven by variations in commodity prices. The novel aspect of the result documented here is that the negative comovement between commodity prices and interest rates is conditional on a permanent change in commodity prices. The finding that interest rates fall when commodity prices increase can also be interpreted as representing a particular manifestation of a phenomenon that Kaminsky and others (2005) refer to as 'When it Rains, it Pours.'

5. Importance of the Commodity Supercycle for Economic Activity

Thus far we have documented that the permanent component of commodity prices explains a sizeable fraction, over one fourth, of movements in commodity prices. In other words, we have documented that there is a significant commodity supercycle. We now wish to ascertain the role of the commodity supercycle in explaining business-cycle fluctuations in individual countries.

Table 3 displays the variance decomposition of output growth for the 24 countries in the sample. On average across countries, the permanent component of commodity prices explains only eight percent of the overall volatility of output growth. By contrast, the transitory components of commodity prices jointly explain 62 percent of the variance of output growth. This result suggests that world shocks are important in explaining output movements in small open economies. However the vast majority of the movements stems from stationary world disturbances. In this sense, the role of the commodity supercycle is modest in accounting for business cycles. The importance of the commodity supercycle in explaining output fluctuations does not vary much across countries. The cross-sectional standard deviation of the variance share of output growth accounted for by ΔX_t^p is only 2.4 percentage points. This means that the relatively modest role played by the supercycle is not just valid on average but applies to most countries in the sample.

Table 3 also speaks to a large literature assessing the role of permanent versus transitory shocks in accounting for aggregate fluctuations in emerging and developed countries.[7] It shows that, in the present sample of 24 countries, the vast majority of fluctuations in output growth is driven by stationary shocks. Jointly the domestic and world stationary shocks (z_t^i and z_t^p) explain 80 percent of the variance of output growth on average across countries. Noticeably this result is obtained not only for the developed countries in the sample but also for the emerging ones (Korea, 80 percent; Mexico, 93 percent; South Africa, 91 percent; and Turkey, 95 percent). The finding that stationary shocks explain the lion's share of output fluctuations in emerging countries is in line with those reported in Garcia-Cicco and others (2010), Chang and Fernández (2013), and Singh (2020).

It is of interest to ascertain the effects of the commodity-price supercycle on commodity prices and aggregate activity at different time horizons. Table 4 presents forecast error variance decompositions of the level of commodity prices and the level of output at horizons of 5, 10, 20, and 30 years computed at the posterior mean of the parameter estimate. The top panel of the table shows that the commodity supercycle plays a sizeable role in explaining the level of commodity prices at all forecasting horizons considered. The median share of X_t^p in the forecast error across the eleven commodity prices ranges from 67 percent at the five-year horizon to 93 percent at the 30-year horizon. This suggests that the commodity supercycle affects commodity prices not just at its own frequency of 20 years or higher but also at shorter frequencies of five to ten years.

By contrast, the commodity supercycle appears to play a secondary role in explaining movements in the level of output at horizons of five and ten years, which are typically associated with business-cycle fluctuations. The bottom panel of the table shows that the contribution of X_t^p in accounting for the forecast error variance is at most 12 percent at horizons of ten years or less. At horizons of 20 and 30 years, which fall into the range of frequencies of the commodity supercycle itself, the contribution of X_t^p to explaining the variance of forecast errors of output increases to 19 percent. By contrast, stationary world shocks, the elements of the vector z_t^p, account for the majority of the forecast error variance of output at all horizons considered. Their median

7. See, for example, Aguiar and Gopinath (2007).

contribution ranges from 75 percent at the five-year forecasting horizon to 58 percent at the 30-year horizon. This indicates that the economic impact of the commodity supercycle on output, relative to that of stationary world shocks, is small at business-cycle frequencies (ten years or less) and moderate at its own frequency (20 years or more).

Table 3. Variance Decomposition of Output Growth

	Shock			
Country	ΔX_t^p	z_t^p	ΔX_t^i	z_t^i
Australia	7	61	1	32
Austria	10	67	1	22
Belgium	8	84	7	1
Canada	10	71	1	19
Denmark	7	65	0	28
Finland	6	68	17	8
France	8	60	1	31
Greece	7	63	30	0
Iceland	5	47	45	2
Ireland	6	42	51	2
Italy	10	74	0	17
Korea, Rep.	11	60	10	20
Luxembourg	10	50	23	18
Mexico	7	71	0	22
Netherlands	8	58	33	1
New Zealand	5	51	36	8
Norway	4	55	19	22
Portugal	13	63	0	24
South Africa	9	61	0	29
Spain	12	69	0	19
Sweden	8	54	0	37
Switzerland	6	62	0	31
Turkey	4	51	0	44
United Kingdom	7	74	1	19
Mean	8	62	12	19
Median	8	62	1	20

Source: Authors' calculations.
Notes: The table presents the share (expressed in percent) of the total variance of output growth explained by shocks to the permanent component of commodity prices, ΔX_t^p, all 12 transitory commodity-price shocks taken together, z_t^p, the country-specific nonstationary shock, ΔX_t^i, and the country-specific stationary shock, z_t^i. The reported numbers are averages over 100,000 draws from the posterior distribution of the variance decomposition.

Table 4. Forecast Error Variance Decomposition of the Level of Commodity Prices and Output

Shock	X_t^p				z_t^p				X_t^i				z_t			
Horizon (in years)	5	10	20	30	5	10	20	30	5	10	20	30	5	10	20	30
Coal	52	66	77	82	48	34	23	18	0	0	0	0	0	0	0	0
Crude Oil	86	89	93	95	14	11	7	5	0	0	0	0	0	0	0	0
Natural Gas	73	83	90	93	27	17	10	7	0	0	0	0	0	0	0	0
Beverages	46	70	85	91	54	30	15	9	0	0	0	0	0	0	0	0
Food	38	52	76	85	62	48	24	15	0	0	0	0	0	0	0	0
Agr. Raw Materials	54	74	87	92	46	26	13	8	0	0	0	0	0	0	0	0
Fertilizers	68	78	87	91	32	22	13	9	0	0	0	0	0	0	0	0
Metal and Minerals	47	63	79	87	53	37	21	13	0	0	0	0	0	0	0	0
Gold	71	83	90	93	29	17	10	7	0	0	0	0	0	0	0	0
Silver	67	78	86	89	33	22	14	11	0	0	0	0	0	0	0	0
Platinum	67	81	90	93	33	19	10	7	0	0	0	0	0	0	0	0
Median	67	78	87	91	33	22	13	9	0	0	0	0	0	0	0	0
Interest Rate	8	7	6	7	92	93	94	93	0	0	0	0	0	0	0	0
Australia	5	2	2	7	82	89	87	80	5	6	8	11	8	3	2	2
Austria	19	27	33	36	74	70	65	63	1	1	1	1	6	2	1	1
Belgium	9	10	13	13	88	87	82	79	3	3	5	8	0	0	0	0
Canada	2	12	33	41	92	85	64	56	3	2	3	3	3	1	0	0
Denmark	6	13	22	24	88	84	76	74	0	0	0	0	6	3	2	2
Finland	5	8	11	10	60	59	50	42	34	32	39	48	1	0	0	0
France	20	27	36	41	70	69	61	56	2	2	2	2	8	2	1	1
Greece	1	2	5	5	95	96	92	91	3	2	2	3	0	0	0	0
Iceland	0	1	1	1	25	22	16	12	65	70	78	84	9	8	5	4
Ireland	4	2	1	2	26	18	11	7	69	79	88	91	1	1	0	0

Table 4. Forecast Error Variance Decomposition of the Level of Commodity Prices and Output (continued)

Shock	X_t^p				z_t^p				X_t^i				z_t			
Horizon (in years)	5	10	20	30	5	10	20	30	5	10	20	30	5	10	20	30
Italy	12	14	15	14	81	83	83	84	0	0	0	1	7	3	1	1
Korea, Rep.	1	0	1	2	19	11	6	4	80	88	93	94	1	0	0	0
Luxembourg	29	35	36	33	43	36	27	22	25	28	36	44	3	1	1	1
Mexico	1	1	1	3	85	93	94	93	0	0	1	1	13	5	3	3
Netherlands	19	27	30	29	75	67	60	55	4	5	10	16	1	0	0	0
New Zealand	1	1	3	5	21	18	11	8	74	80	85	86	3	2	1	1
Norway	6	4	5	4	77	77	70	63	14	17	24	32	3	2	1	1
Portugal	24	23	25	25	69	74	74	73	0	0	0	0	7	3	2	1
South Africa	43	42	38	50	49	55	59	47	1	1	1	1	6	3	1	1
Spain	7	18	30	33	87	80	69	65	0	0	1	1	6	1	1	1
Sweden	8	36	59	67	74	58	38	30	2	2	1	2	16	5	2	1
Switzerland	8	15	29	33	75	75	65	61	1	1	2	3	16	9	4	3
Turkey	1	7	7	9	67	74	77	75	0	0	1	1	31	19	15	15
United Kingdom	17	28	39	42	77	67	56	52	2	3	4	5	4	2	1	1
Mean	10	15	20	22	67	64	58	54	16	18	20	22	7	3	2	2
Median	6	12	19	19	75	72	65	58	3	2	2	3	6	2	1	1

Source: Authors' calculations.
Note: All shares are computed at the posterior mean of the estimated parameters and are expressed in percentage points.

Figure 4. Impulse Response of Output to a Unit Long-run Increase in X_t^p

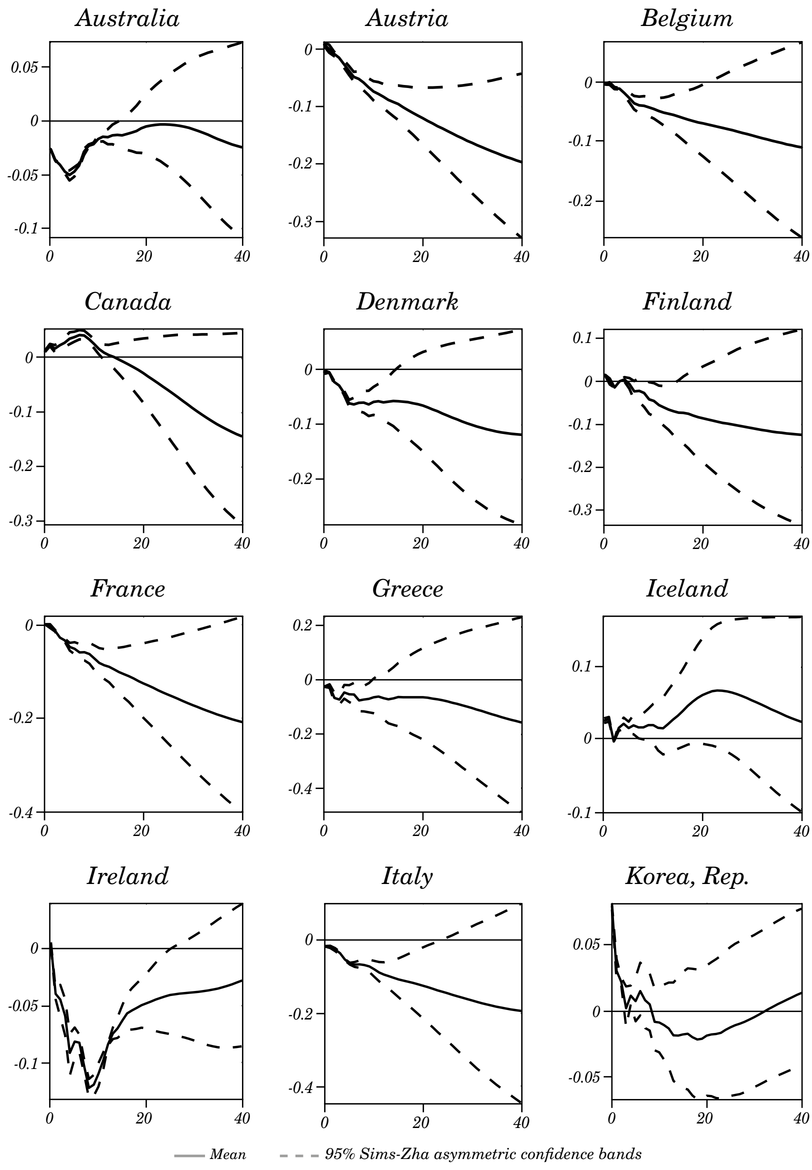
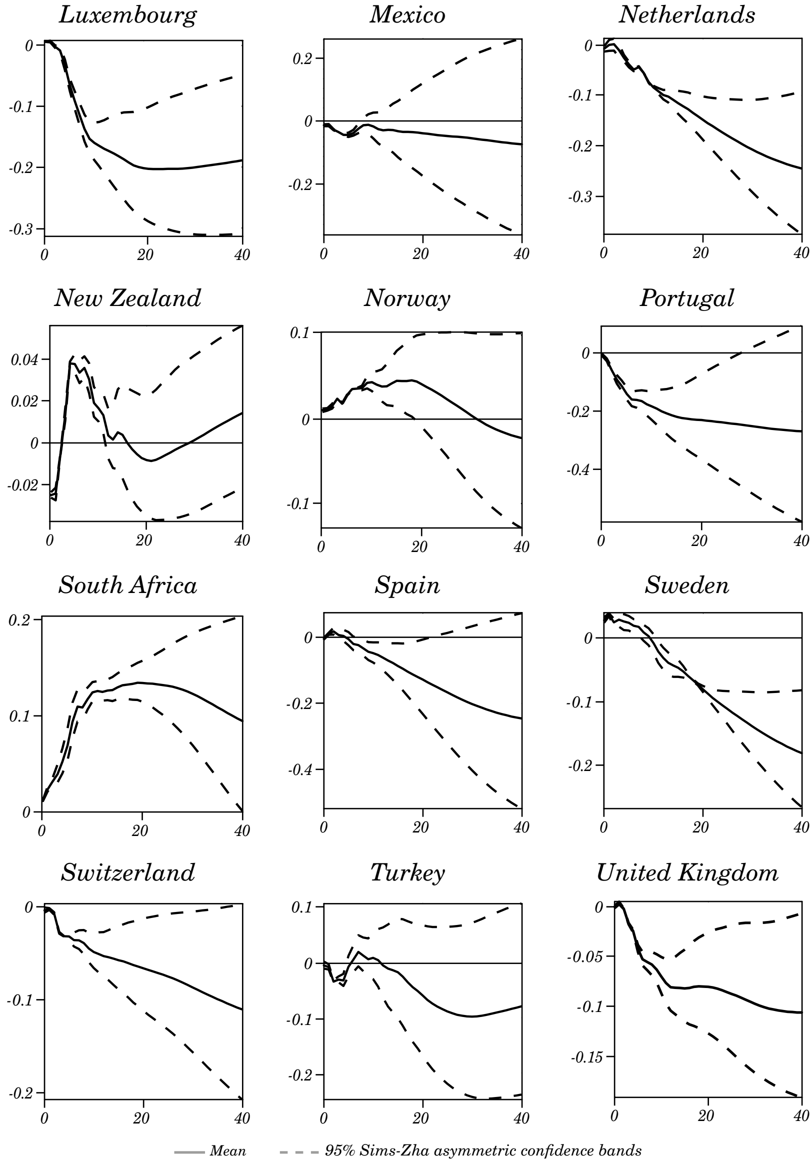


Figure 4. Impulse Response of Output to a Unit Long-run Increase in X_t^p (continued)



Source: Authors' calculations.

The fact that the world and domestic stationary shocks, z_t^p and z_t^i, jointly explain the majority of the forecast error variance of output even at horizons of 20 and 30 years, 65 and 60 percent on average, respectively, indicates that the world and domestic nonstationary components, X_t^p and X_t^i, are not the dominant drivers of movements in output.

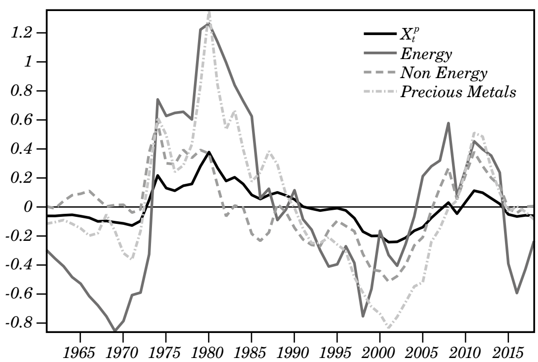
Figure 4 displays the impulse responses of the level of output, y_t, in each of the 24 countries, to a permanent world shock, X_t^p, that increases commodity prices in the long run by one percent. In most countries, the permanent commodity-price increase is contractionary. One possible explanation for this finding is that the sample includes mostly developed open economies that are not important primary commodity producers. As we will see in section 6, in emerging countries the output response to an increase in the permanent component of world prices is in general positive. But even for primary commodity producers, an increase in X_t^p could have an ambiguous effect on output for at least two reasons. One is that when X_t^p goes up, all commodity prices go up. To the extent that some commodities are imported and used as intermediate inputs in domestic production, an increase in X_t^p would result in an increase in marginal costs, which in turn, may lower domestic employment. The second reason is that because X_t^p represents a permanent increase in real commodity prices, it might entail a large positive wealth effect for the commodity-producing country. In turn, this positive wealth effect could lead to a contraction in labor supply and in this way lower equilibrium employment.

6. Emerging Countries

As mentioned earlier, long quarterly time series for output are available mostly for developed countries. As a result, emerging countries are underrepresented in the sample. To shed light on the importance of the commodity supercycle in emerging countries, this section turns to an analysis based on annual data for which the coverage of this group of countries is more comprehensive. The empirical model is the one described in section 1 except for the number of lags and the number of commodity prices included. Because the data is annual, the model includes only one lag of prices and output, $\hat{p}_t$ and $\hat{y}_t$. To economize on the number of parameters estimated, the eleven commodity prices are aggregated into three indices—energy, non-energy, and precious metals—, following the Pink Sheet aggregation scheme. Energy commodities include coal, crude oil, and natural gas. Non-energy commodities include beverages, food, agricultural raw materials, fertilizers, and metals and minerals. And precious metals comprise gold, silver, and platinum. The sample includes 24 emerging countries and 17 developed countries, which are listed in table 6. The output data comes from World Development Indicators. The selection of countries follows a number of criteria which include data availability since 1960, a population of more than three million in 2018, not having transitioned from a planned to a market economy, and having a common secondary data source for output. As in the analysis using quarterly data, the model is estimated using Bayesian techniques. The prior distributions for the model parameters are the same as those presented in table 1.

Figure 5 plots with thin lines the three real commodity-price indices and with a thick line their estimated permanent component, X_t^p. As in the case of the estimation on quarterly data, the commodity supercycle is a smooth stochastic trend of the three prices and displays two peaks since 1960—one in 1980 and the other in 2012. The peaks and troughs of the estimated commodity-price supercycle line up with the ones identified using quarterly data on the more disaggregated commodity prices plotted in figure 2.

Figure 5. The Commodity-Price Supercycle in Annual Data



Source: Authors' calculations.
Notes: The permanent component of the three aggregate commodity-price indices, X_t^p, is computed by Kalman smoothing using the posterior mean of the parameter estimates. All time series are constructed as cumulative demeaned growth rates.

Table 5. Percent of Variance of the Growth Rate of Annual World Prices Explained by ΔX_t^p

Price of	*Mean*	*Std. Dev.*
Energy Commodities	98	1
Non-Energy Commodities	94	2
Precious Metals	94	1
Mean	95	1
Median	94	1
Real Rate	25	18

Source: Authors' calculations.
Note: The reported figures are based on 100,000 draws from the posterior distribution of the variance decomposition.

Table 5 shows that the permanent component, X_t^p, explains more than 90 percent of the variation in the growth rate of the three commodity indices. Thus, as in the quarterly estimation, the commodity supercycle is an important driver of commodity prices. A difference is that now the share of the variance of the growth rate of prices explained by the permanent component is much larger than the one estimated in quarterly data. This is to some extent expected since aggregation across time and commodities tends to average away the effects of commodity-specific and transitory disturbances. The table also shows that the commodity supercycle explains 25 percent of movements in the world interest rate, a share somewhat higher than the one obtained in the estimation on quarterly data (15 percent).

Table 6 displays the variance decomposition of output growth in the 24 emerging and 17 developed countries considered. As in the case of the analysis using quarterly data on a sample of mostly developed economies, all world shocks taken together, X_t^p and z_t^p, play a major role in explaining the variance of output growth. It also continues to be the case that, of the contribution of world shocks to output fluctuations, the majority is attributed to stationary disturbances, z_t^p.

This pattern applies to a large extent when one limits attention to emerging countries. Within this group, on average world shocks explain more than 50 percent of the variance of output and, of this fraction, almost two thirds are attributable to stationary world shocks.[8]

The fact that stationary world shocks, z_t^p, explain a much larger share of the variance of output growth than of the variance of the growth rate of prices indicates that these world shocks may be only partially mediated through commodity prices. An example of a world shock that could have an output effect both directly and through world commodity prices are productivity shocks that are correlated across countries.

Table 6 also speaks to the literature on the role of stationary and nonstationary shocks in explaining business cycles in emerging countries. The posterior mean joint contribution of stationary shocks, z_t^p and z_t^i, to the variance of output growth is 57 percent with the remaining 43 percent explained by nonstationary shocks, $X_t^p$and X_t^i. This result suggests that the majority of fluctuations in aggregate activity in the emerging countries considered stems from stationary domestic and world disturbances.

The preponderance of stationary world shocks in accounting for movements in output in emerging economies also manifests itself at different forecasting horizons. Table 7 displays the forecast error variance decomposition of the level of output at horizons 5, 10, 20, and 30 years. At forecasting horizons of five and ten years, which are typically associated with business-cycle frequencies, the mean share of variance explained by the stationary world shocks, z_t^p, is 40 and 38 percent, respectively, compared to 19 and 24 percent explained by the nonstationary world shock, X_t^p. At longer forecasting horizons of 20 and 30 years, the role of nonstationary world shocks increases, as expected, but does not clearly dominate that of stationary world shocks. Specifically, the variance of the forecasting error of output explained by X_t^p has a mean of 30 and 34 percent at horizons 20 and 30 years, compared to 34 and 31 percent for stationary world shocks.

8. The results are robust to estimating the model by maximum likelihood, which indicates that the findings are not due to the choice of priors. The maximum likelihood estimate assigns more importance (about 10 percentage points) to world shocks in explaining the variance of output growth. For emerging countries, of this, about one third is accounted for by the supercycle and two thirds by stationary world shocks. For developed countries, the relative importance of the supercycle is somewhat larger, with innovations to ΔX_t^p explaining about forty percent of the variance of output growth accounted for by world shocks.

Table 6. Variance Decomposition of Output Growth — Annual Data

	Shock			
Country	ΔX_t^p	z_t^p	ΔX_t^i	z_t^i
Mean Emerging	18	32	24	25
Mean Developed	19	48	13	20
Argentina	16	8	74	1
Bangladesh	8	17	73	1
Bolivia	26	55	0	19
Brazil	21	33	0	45
Chile	10	18	0	71
Colombia	26	28	2	44
Costa Rica	25	42	28	4
Dominican Republic	8	8	0	84
Ecuador	34	32	33	1
Guatemala	20	78	1	1
India	10	24	63	2
Indonesia	15	50	32	2
Korea, Rep.	19	55	26	1
Malaysia	24	54	0	21
Mexico	12	40	47	1
Pakistan	8	35	57	1
Panama	14	16	0	70
Paraguay	30	19	0	50
Peru	18	19	0	63
Philippines	20	16	60	2
South Africa	35	27	1	36
Thailand	13	60	6	20
Turkey	4	16	79	0
Uruguay	18	21	0	61
Australia	6	27	64	4
Austria	25	51	1	22
Belgium	21	64	1	13
Canada	16	43	1	40
Denmark	21	54	2	21
Finland	17	60	19	3
France	18	77	2	4
Greece	22	47	0	30
Iceland	8	17	0	75
Italy	25	58	0	17
Luxembourg	28	22	50	1
Netherlands	18	50	0	32
Norway	10	39	3	49
Portugal	19	58	22	1
Spain	23	53	0	24
Sweden	14	63	21	2
United Kingdom	30	33	34	2

Source: Authors' calculations.

Notes: The table presents the share (expressed in percent) of the total variance of output growth explained by shocks to the permanent component of commodity prices, ΔX_t^p, all stationary world price shocks taken together, z_t^p, the country-specific nonstationary shock, ΔX_t^i, and the countryspecific stationary shock, z_t^i. The reported numbers are averages over 100,000 draws from the posterior distribution of the variance decomposition.

Figures 6 and 7 display the impulse response of output in the 17 developed and 24 emerging economies, respectively, to a shock in X_t^p that increases energy, non-energy, and precious metal prices in the long run by one percent. The figures also include 95-percent confidence bands. In line with the results obtained in section 5 using quarterly data (figure 4), in developed economies, a permanent increase in world commodity prices is contractionary for most countries. By contrast, for most emerging countries a permanent increase in commodity prices is expansionary. As pointed out in section 5, a possible explanation for this difference could be that in emerging countries the production of primary commodities represents a larger share of total output than it does in developed countries.

Table 7. Forecast Error Variance Decomposition of the Level of Output — Annual Data

Shock	X_t^p				z_t^p				X_t^i				z_t			
Horizon (in years)	5	10	20	30	5	10	20	30	5	10	20	30	5	10	20	30
Argentina	45	33	20	14	10	6	4	3	45	61	76	83	1	0	0	0
Bangladesh	15	12	9	9	26	41	38	33	55	45	52	57	4	2	1	1
Bolivia	1	6	48	65	63	69	41	30	0	1	1	1	36	25	10	5
Brazil	23	28	31	32	39	51	57	58	0	0	1	1	38	21	12	9
Chile	5	5	11	20	15	16	15	14	0	1	3	4	80	78	71	63
Colombia	34	42	58	69	29	30	25	18	1	2	2	2	37	26	15	10
Costa Rica	11	10	38	60	68	61	35	18	17	26	26	21	3	2	1	0
Dominican Republic	20	33	43	49	15	19	19	17	0	0	0	1	65	48	38	34
Ecuador	59	67	66	63	34	26	23	21	7	7	11	16	1	1	0	0
Guatemala	17	13	10	11	82	86	86	84	1	1	3	5	1	1	0	0
India	9	10	16	17	54	64	64	61	32	24	20	21	5	2	1	1
Indonesia	21	32	44	46	66	57	39	32	12	10	16	21	2	1	1	0
Korea, Rep.	9	7	5	9	77	54	49	45	12	38	45	46	3	1	0	0
Malaysia	16	18	31	42	50	43	36	32	0	1	2	2	33	38	31	24
Mexico	19	22	24	24	70	64	57	53	10	14	19	23	1	0	0	0
Pakistan	0	2	16	28	71	67	48	37	27	29	35	35	2	2	1	0
Panama	13	26	36	36	14	14	13	14	0	1	1	2	73	59	50	47
Paraguay	37	54	65	68	18	21	22	22	0	0	0	0	45	25	13	10
Peru	24	38	38	35	13	12	20	27	0	0	1	1	63	50	42	37
Philippines	19	19	12	10	5	5	10	14	74	75	77	75	2	1	1	1
South Africa	45	54	59	60	20	18	17	16	1	1	2	4	35	27	22	20
Thailand	1	1	6	14	66	63	67	65	1	3	7	8	32	32	20	13
Turkey	4	3	4	6	35	22	12	8	61	75	84	86	1	0	0	0

Table 7. Forecast Error Variance Decomposition of the Level of Output — Annual Data (continued)

Shock	X_t^p				z_t^p				X_t^i				z_t			
Horizon (in years)	5	10	20	30	5	10	20	30	5	10	20	30	5	10	20	30
Uruguay	22	31	33	30	13	12	18	25	0	1	1	2	64	56	48	43
Mean-Emerging	19	24	30	34	40	38	34	31	15	17	20	21	26	21	16	13
Australia	0	0	0	0	69	62	46	36	24	34	52	62	7	4	2	1
Austria	4	2	1	2	86	93	95	94	1	1	1	2	9	4	2	2
Belgium	9	7	9	13	86	90	88	83	1	1	2	3	4	2	1	1
Canada	15	15	15	14	50	59	64	65	1	1	2	3	35	25	19	17
Denmark	13	12	8	7	68	75	81	82	2	3	5	6	16	9	6	5
Finland	7	4	3	3	86	87	85	81	3	6	10	14	4	3	2	2
France	2	1	5	13	94	96	92	84	1	2	2	2	3	2	1	1
Greece	21	30	38	41	55	59	57	55	0	0	0	0	24	11	5	4
Iceland	8	14	26	35	27	41	46	43	0	0	1	1	65	44	28	21
Italy	5	5	18	29	84	90	80	70	0	0	0	0	11	4	2	1
Luxembourg	28	37	33	27	47	32	22	18	23	31	45	55	1	1	0	0
Netherlands	3	5	6	6	72	81	85	86	0	1	1	1	24	13	8	7
Norway	1	6	17	24	48	66	69	65	3	4	4	4	48	23	10	7
Portugal	4	3	2	3	91	92	92	90	4	4	5	7	1	1	0	0
Spain	6	15	21	22	69	72	72	72	0	0	1	1	25	12	7	6
Sweden	4	3	6	10	89	88	77	66	5	8	16	23	2	2	1	1
United Kingdom	29	36	37	35	52	39	31	26	17	24	32	39	2	1	0	0
Mean-Developed	9	12	14	17	69	72	70	66	5	7	11	13	17	9	5	4

Source: Authors' calculations.
Note: All shares are computed at the posterior mean of the estimated parameters and are expressed in percentage points.

Figure 6. Impulse Responses of Output to a Unit Long-Run Increase in X_t^p — Developed Economies

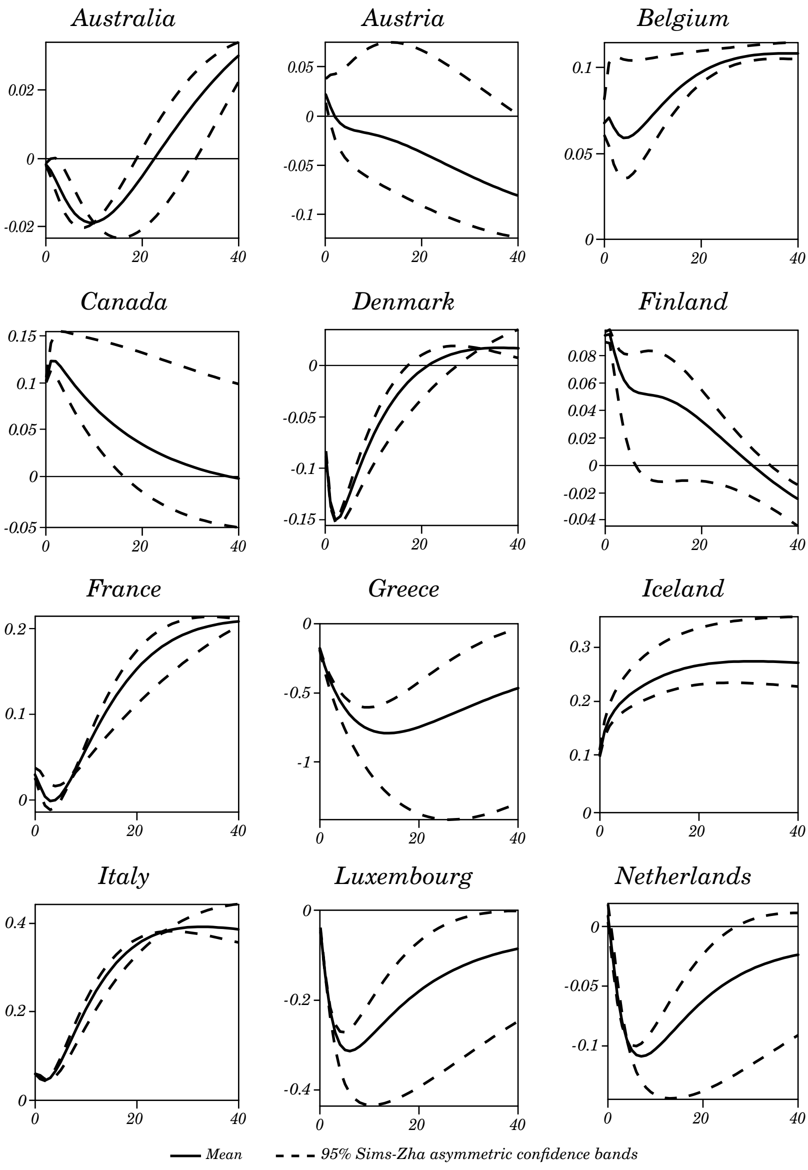
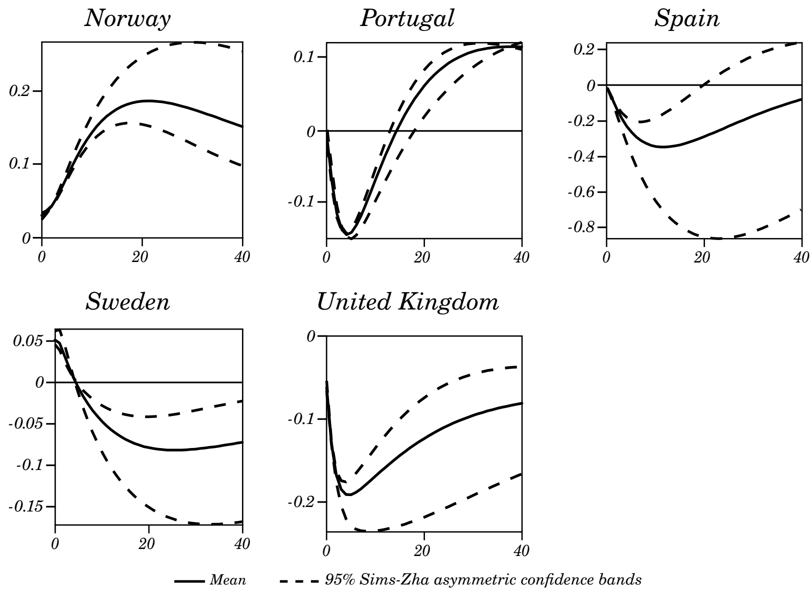


Figure 6. Impulse Responses of Output to a Unit Long-Run Increase in X_t^p — Developed Economies (continued)



Source: Authors' calculations.

Figure 7. Impulse Response of Output to a Unit Long-Run Increase in X_t^p: Emerging Economies

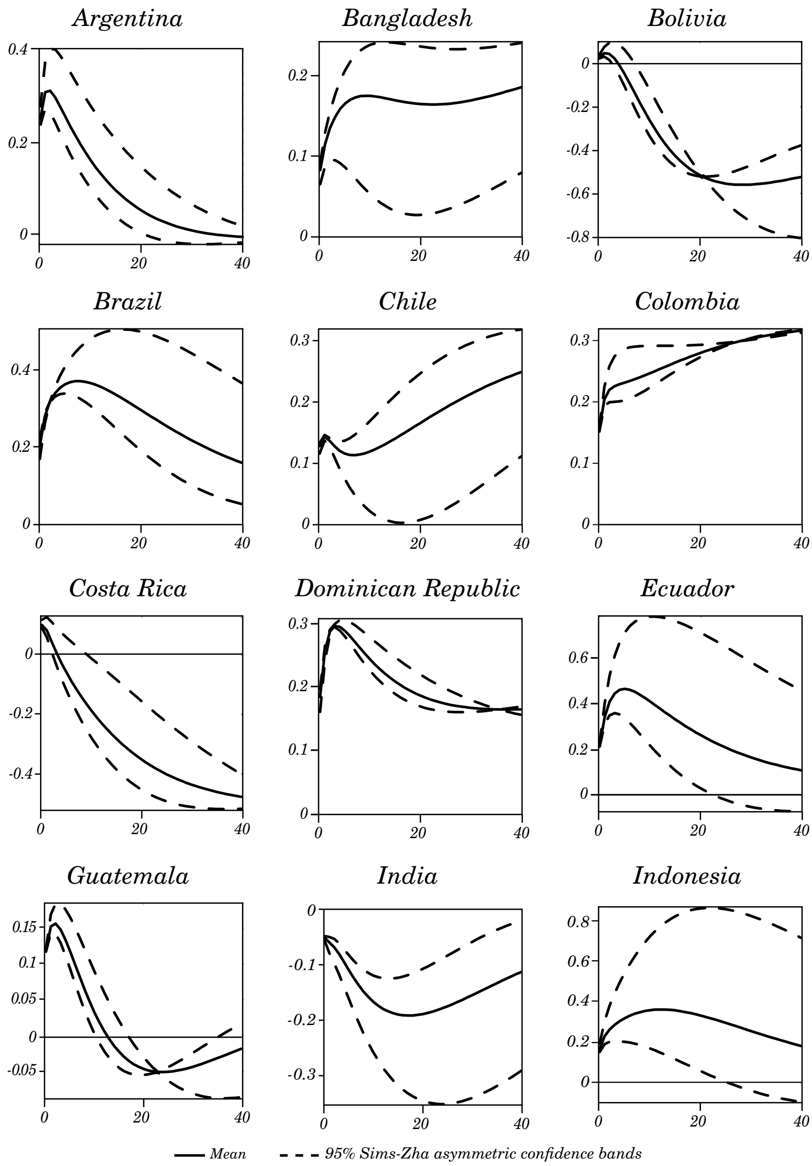
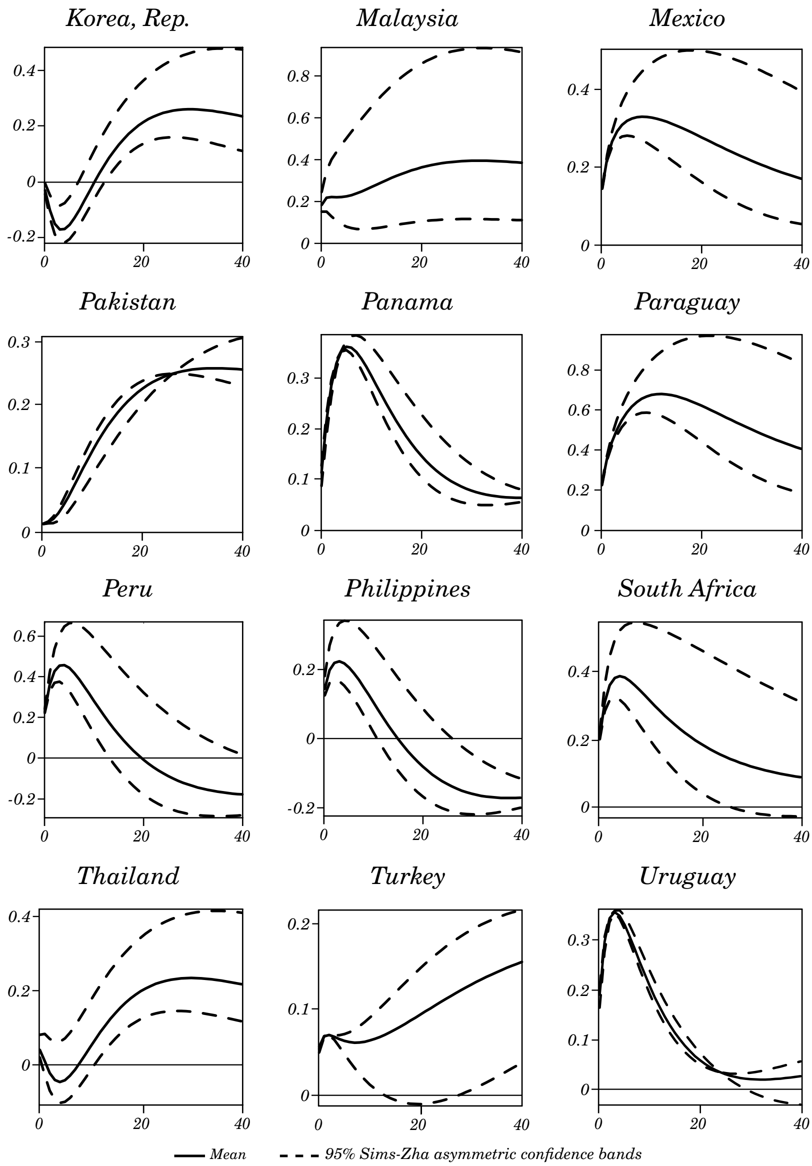


Figure 7. Impulse Response of Output to a Unit Long-Run Increase in X_t^p: Emerging Economies (continued)



Source: Authors' calculations.

7. Conclusion

This paper aims to fill a gap in the literature on the transmission of world shocks through commodity prices in open economies. An existing literature has documented the presence of a commodity-price supercycle. An empirical technique employed in many of these studies is based on spectral analysis and identifies one supercycle per commodity price. The resulting supercycles are positively correlated across commodities suggesting a common driver.

The first contribution of the present paper is to propose an alternative definition of the commodity-price supercycle consisting in representing it as the common stochastic trend in all commodity prices. The so-identified supercycle turns out to share a number of key characteristics with the ones obtained using spectral analysis. An advantage of the common permanent component approach is that it lends itself to a joint estimation of the contributions of domestic and foreign transitory and permanent shocks to aggregate fluctuations in open economies.

The results of the paper suggest that world shocks are responsible for more than half of observed variations in aggregate activity in developed and emerging economies. However, more than two thirds of the contribution of world shocks is due to temporary disturbances, leaving less than one third to the permanent world shock that drives the commodity supercycle. This result is obtained both unconditionally and conditional on forecasting horizons. Importantly, even at horizons of 20 and 30 years, which are typically associated with the periodicity of the commodity-price supercycle, the permanent world shock does not clearly dominate temporary world shocks in accounting for variations in aggregate activity.

Taken together, these findings indicate that the permanent world shock that drives the commodity-price supercycle does matter but does not play the central role in shaping short- or medium-run business-cycle fluctuations.

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Appendix

In section 1, the presentation of the model omitted constant terms to facilitate the exposition. This Appendix presents the model including those omitted constant terms. As we will see, this will introduce a vector of constants, denoted A, into the observation equation (9). We will also derive expressions for the matrices F, P, and H of the state-space representation of the model, equations (8) and (9).

Redefine the vectors $\hat{p}_t$, $\hat{y}_t$, and u_t as deviations from their respective means:

$$\hat{p}_t = \begin{bmatrix} p_t^1 - X_t^p - E\left(p_t^1 - X_t^p\right) \\ p_t^2 - X_t^p - E\left(p_t^2 - X_t^p\right) \\ \vdots \\ p_t^{11} - X_t^p - E\left(p_t^{11} - X_t^p\right) \\ r_t - Er_t \end{bmatrix};$$

$$\hat{y}_t = \begin{bmatrix} y_t^1 - X_t^1 - \alpha^1 X_t^p - E\left(y_t^1 - X_t^1 - \alpha^1 X_t^p\right) \\ y_t^2 - X_t^2 - \alpha^2 X_t^p - E\left(y_t^2 - X_t^2 - \alpha^2 X_t^p\right) \\ \vdots \\ y_t^{24} - X_t^{24} - \alpha^{24} X_t^p - E\left(y_t^{24} - X_t^{24} - \alpha^{24} X_t^p\right) \end{bmatrix}$$

and

$$u_t \equiv \begin{bmatrix} \Delta X_t^p - E\left(\Delta X_t^p\right) \\ \Delta X_t - E\left(\Delta X_t\right) \\ z_t^p - E\left(z_t^p\right) \\ z_t - E\left(z_t\right) \end{bmatrix}.$$

The evolution of the vector $\hat{p}_t$ is

$$\hat{p}_t = \sum_{i=1}^{4} B_{pp}^i \hat{p}_{t-1} + C_{pX^p}\left(\Delta X_t^p - E\left(\Delta X_t^p\right)\right) + C_{pz^p}\left(z_t^p - E\left(z_t^p\right)\right). \tag{10}$$

The evolution of the vector $\hat{y}_t$ is

$$\begin{aligned}\hat{y}_t = & \sum_{i=1}^{4} B_{yp}^{i}\hat{p}_{t-i} + \sum_{i=1}^{4} B_{yy}^{i}\hat{y}_{t-i} \\ & + C_{yX^p}\left(\Delta X_t^p - E\left(\Delta X_t^p\right)\right) + C_{yz^p}\left(z_t^p - E\left(z_t^p\right)\right) \\ & + C_{yX}\left(\Delta X_t - E\left(\Delta X_t\right)\right) + \left(z_t - E\left(z_t\right)\right).\end{aligned} \tag{11}$$

The evolution of the exogenous shocks, u_t, is

$$u_t = \rho u_{t-1} + \psi \nu_t. \tag{12}$$

Let

$\hat{x}_t = \left[\hat{p}_t \ \hat{y}_t\right]'$; and $\xi_t = \left[\hat{x}_t \ \ \hat{x}_{t-1} \ \ \hat{x}_{t-2} \ \ \hat{x}_{t-3} \ \ u_t\right]'$.

The system of equations (10), (11), and (12) can then be expressed as:

$$\hat{x}_{t+1} = B\begin{bmatrix}\hat{x}_t \\ \hat{x}_{t-1} \\ \hat{x}_{t-2} \\ \hat{x}_{t-3}\end{bmatrix} + Cu_{t+1}, \tag{13}$$

where

$$B \equiv \begin{bmatrix} B_{pp}^1 & \varnothing_{12\times 24} & B_{pp}^2 & \varnothing_{12\times 24} & B_{pp}^3 & \varnothing_{12\times 24} & B_{pp}^4 & \varnothing_{12\times 24} \\ B_{yp}^1 & B_{yy}^1 & B_{yp}^2 & B_{yy}^2 & B_{yp}^3 & B_{yy}^3 & B_{yp}^4 & B_{yy}^4 \end{bmatrix}$$

and

$$C \equiv \begin{bmatrix} C_{\hat{p}X^p} & \varnothing_{12\times 24} & C_{\hat{p}z^p} & \varnothing_{12\times 24} \\ C_{\hat{y}X^p} & C_{\hat{y}X} & C_{\hat{y}z^p} & I_{24\times 24} \end{bmatrix}.$$

The vector ξ_t evolves over time as

$$\xi_{t+1} = F\xi_t + P\nu_{t+1},$$

where

$$F = \begin{bmatrix} B & C\rho \\ I_{108\times 144} & \varnothing_{108\times 61} \\ \varnothing_{61\times 144} & \rho \end{bmatrix}; \text{ and } P = \begin{bmatrix} C\psi \\ \varnothing_{108\times 61} \\ \psi \end{bmatrix}.$$

Given the redefinition of $\hat{p}_t$, the observation equations (4), (5), and (6) become

$$\Delta p_t^i = \Delta \hat{p}_t^i + \left(\Delta X_t^p - E\left(\Delta X_t^p\right)\right) + E\left(\Delta X_t^p\right); i = 1,\ldots,11,$$

$$\Delta y_t^i = \Delta \hat{y}_t^i + \left(\Delta X_t^i + \alpha^i \Delta X_t^p - E\left(\Delta X_t^i + \alpha^i \Delta X_t^p\right)\right)$$
$$+ E\left(\Delta X_t^i + \alpha^i \Delta X_t^p\right); i = 1,\ldots,24,$$

and

$$r_t = \left(r_t - Er_t\right) + Er_t.$$

In vector form, the observation equations can be expressed as

$$o_t = A' + H'\xi_t + \mu_t$$

where

$$A = \left[E\left(\Delta X_t^p\right)\ldots E\left(\Delta X_t^p\right)\ Er_t\ \ E\left(\Delta X_t^1 + \alpha^1 \Delta X_t^p\right)\ \ldots\ E\left(\Delta X_t^{24} + \alpha^{24} \Delta X_t^p\right)\right],$$

$$H' = \left[I_{36\times 36} \quad H_2 \quad \varnothing_{36\times 72} \quad H_4 \quad H_5 \quad \varnothing_{36\times 36}\right]$$

with

$$H_2 = -\begin{bmatrix} I_{11\times 12} & \varnothing_{11\times 24} \\ \varnothing_{1\times 12} & \varnothing_{1\times 24} \\ \varnothing_{24\times 12} & I_{24\times 24} \end{bmatrix}$$

$$H_4 = \left[1_{1\times 11}\ \ 0\ \ \alpha^1\ \ \ldots\ \alpha^{24}\right]'$$

$$H_5 = \begin{bmatrix} \varnothing_{12\times 24} \\ I_{24\times 24} \end{bmatrix}.$$

Cross-Border Corporate Control: Openness and Tax Havens

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Cross-border corporate control is a major facet of globalisation. In roughly one out of four listed controlled companies in 2012, control was exercised by a foreign entity or family/individual. Controlling—and passive—ownership stakes are often hidden in complex structures, involving pyramids and chains of intermediate firms. Besides, shareholders often use shell companies incorporated in financial offshore centres. As we demonstrate in this paper, even locals use firms in tax-haven jurisdictions as conduits of their (controlling) equity stakes in domestic firms. However, international corporate control is not well-understood due to the esoteric corporate holding schemes and the complex network of equity holdings. We take a first step in understanding cross-border corporate control by documenting some broad patterns, based on our ongoing research of the drivers of the internationalisation of corporate control (Fonseca and others, 2022).

We are thankful to Carolina Villegas-Sanchez, Şebnem Kalemli-Özcan, Winfrid Blaschke, and participants at the Annual Conference of the Central Bank of Chile 2021 for their useful comments and valuable feedback. We thank Divyakshi Jain and, in particular, Andreas Miyashiro for excellent research assistance. This paper should not be reported as representing the views of the European Central Bank (ECB). The views expressed are those of the authors and do not necessarily reflect those of the ECB. This paper was written while Luís Fonseca was a PhD student at London Business School.

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By compiling new ownership data for almost 90 percent of the world market capitalisation of listed firms in 2012, we provide a mapping of corporate control, zooming into the role of tax havens, whose role, while prominent, is not well-understood due to secrecy, lack of data and transparency.

In section 1, we discuss the data compilation, which extends our earlier work (Aminadav and Papaioannou, 2020). Relying on a plethora of sources (e.g. regulatory filings, company reports, financial media), we augment, update, and revise the ORBIS database on corporate ownership to identify ultimate controlling shareholders for 25,884 listed firms in 86 jurisdictions in 2012. The 83,942 shareholders and ultimate owners come from 90 territories. We give examples of ownership structures for various controlling entities, individuals, banks, governments, and nonbank financial institutions, distinguishing between three nationality types for the ultimate controlling entity and the immediate controlling shareholding entity: (a) domestic, (b) foreign, and (c) foreign tax-haven.[1] We also compare our newly compiled proxies of international corporate ownership of listed firms with the widely used external wealth of nations statistics of Lane and Milesi-Ferretti (2018, 2021) and find strong correlations.

In section 2, we provide an anatomy of international corporate control and cross-border ownership of listed corporations. First, we uncover large differences in cross-border corporate ownership and control around the world. The degree of 'openness', reflecting the share of market capitalisation (and share of listed firms) controlled by foreign entities differs considerably, even when looking at countries of similar income and in the same region. Second, when we tabulate differences across income groups and explore the role of market size, we find that foreign control is less common in richer and more populous countries, echoing the international trade and portfolio investment evidence.

In section 3, we zoom into the role of tax-haven-incorporated vehicles in the exercise of control. The use of tax havens in 2012 appears, on average, moderate, but quite heterogeneous, even within regions. We find evidence that lower-income countries have higher shares of control of their companies by or via entities in tax havens, but not that poorer countries are more likely to exert control through tax havens. We find that, in a few countries, domestic entities, including

1. Following the classification of the OECD (2000) and Tørsløv and others (2018).

families and individuals, hold controlling equity stakes in firms listed in the local stock exchange by using intermediate firms incorporated in tax-haven jurisdictions. This pattern is higher in Ukraine, Russia, Greece, and Serbia, as well as in China. The exercise of control by or via tax-haven-incorporated vehicles appears to be low in the United States. This may be so because our data do not distinguish the state of incorporation, which would be useful due to the case of Delaware, which has been identified as a tax haven (Michel, 2021).

Our paper relates and contributes to various strands of research in the literature on international economics and corporate finance:

First, our paper mostly connects to the voluminous literature on trade, foreign direct investment (FDI), portfolio, and bank flows (Lane and Milesi-Ferretti, 2007 and 2008; Portes and Rey, 2005; Wei, 2000; Aviat and Coeurdacier, 2007; Papaioannou, 2009; Hau and Rey, 2008; Alfaro and others, 2008). Rather than looking at volatile capital flows, we examine international corporate control, which is more persistent. Examining corporate control allows for a more in-depth mapping of global market integration. Our data and effort here and in our companion papers (Aminadav and Papaioannou, 2020; Fonseca and others, 2022) have been on mapping actual ties and incorporating indirect links; for example, a Russian national controlling a Brazil-incorporated listed corporation via a Cypriot or Maltese 'shell' company. We try addressing a major shortcoming of most international asset holdings and liabilities positions datasets—IMF International Financial Statistics (IFS), U.S. Treasury International Capital (TIC) System—that, following the *residence principle*, misses indirect exposure. While international institutions, policymakers, and researchers increasingly acknowledge this issue, there has been limited progress in capturing indirect exposure, which anecdotal evidence and case studies suggest is becoming extensive. Important exceptions are the parallel and independent works of Coppola and others (2021), and Damgaard and others (2019). The former study international bond and equity issuance via special purpose vehicles (SPV) documenting the chief role of tax havens. The latter combine foreign direct investment data from various sources to approximate real and 'phantom' FDI, often channelled via countries with low-tax systems tailored for multinationals. Rather than looking at corporate debt issuance and multinationals' activities, we look at corporate ownership and control, major facets of globalisation that have not been much researched.[2]

Second, our findings that a non-negligible portion of international corporate control gets through offshore financial centres contribute to a nascent but fast-growing research agenda on their increasing role in the global economic system (Hines and Rice, 1994; Zucman, 2015; Tørsløv and others, 2018). The literature focuses on how corporations shift earnings across jurisdictions (Johannesen and others, 2020; Guvenen and others, 2017), how tax havens allow hiding assets (Alstadsæter and others, 2018), and even money laundering and criminal activity (Andersen and others, 2020). We show that offshore financial centres play a crucial conduit role in the internationalisation of corporate control.

Third, our paper adds to research in corporate finance studying cross-country differences in corporate control (La Porta, Lopez-De-Silanes, and Shleifer, 1999; Claessens and others, 2000; Faccio and Lang, 2002; Laeven and Levine, 2007; Franks and others, 2012). This research mostly works with relatively small samples and countries. We take a panoramic view covering the vast majority of listed corporations across the world. We revise, clean, and extend the dataset of Aminadav and Papaioannou (2020), who in turn have expanded the ORBIS dataset, to identify control from the often obscure structures of corporate ownership. We zoom in on the internationalisation of corporate control, which has not been much studied—except for the parallel and independent work of De La Cruz and others (2019).

1. Data and Methodology

In this section, we first go over the ownership data used to identify corporate control of public (listed) corporations. Second, we discuss our methodology to identify ultimate controlling shareholders from obscure structures of corporate ownership. Third, we present, providing company examples, our methodology to classify domestic, foreign, and tax-haven control and direct ownership. Fourth, we discuss our

2. For example, Coppola and others (2021) are able to record both direct U.S. investments into the Brazilian corporate-bond market and indirect investments via subsidiaries in the Cayman Islands and Bermuda. Likewise, we are able to trace direct equity stakes of U.S. nationals to Brazil, as well as indirect links via private companies in offshore financial centres (e.g. Panama), but also other jurisdictions (e.g. Chile).

aggregation of corporate control across countries, distinguishing between destination and source. Fifth, we present tabulations comparing our measures of international corporate control with the widely used data compiled by Lane and Milesi-Ferretti (2018).

1.1 Ownership Data

The corporate ownership and control data we use builds on the work in Aminadav and Papaioannou (2020), who in turn extend, clean, and update the ORBIS dataset.[3]

1.1.1 Procedure

We proceed as follows.

• We start with Bureau van Dijk's (BvD) ORBIS database.[4] BvD collects ownership information from company reports, financial news, private correspondence, and local specialised agencies. BvD reports shareholder's voting rather than cash-flow rights, taking into account dual shares, "golden shares", and other special share types. This makes them suitable for identifying control.[5] We extract information for publicly traded corporations from ORBIS. We correct inconsistencies, omissions, and errors (e.g. double entries).

• We then match ORBIS' corporate ownership information with Datastream (Thompson Reuters) and Compustat (North America and Global) to get firms' market capitalisation, industry, and other information.

• ORBIS data have gaps on shareholders for many private companies, which prevents tracing ultimate controllers of listed companies. We manually checked and added information on control for firms with incomplete coverage. This work started with Aminadav and Papaioannou (2020), who gathered information on ultimate control for 10,857 listed companies whose ultimate controller could not be traced from ORBIS for 2004–12; they obtained ownership information for

3. Aminadav and Papaioannou (2020) goal was to re-examine the link between corporate control and legal origin and institutions for the largest possible sample of publicly traded firms. We refer interested readers to Aminadav and Papaioannou (2020) main paper and appendix for details on the data.

4. Kalemli-Özcan and others (2015) discuss practical details in building samples from this database.

5. See also Massa and Zaldokas (2016), Kalemli-Özcan and others (2015), Franks and others (2012).

about 7,000 private firms, which appear in ORBIS as main shareholders of listed companies. They relied on financial data providers (Bloomberg, Dun & Bradstreet, Google Finance, Credit Risk Monitor, and Forbes), government publications, reports from regulatory agencies, news, and data made available by the *International Consortium of Investigative Journalists*. For the current paper, we focus on 2012 and we expanded the search into the corporate ownership structure of 4,002 listed firms that Aminadav and Papaioannou (2020) could not trace control. These firms had 3,695 unique controllers, usually private firms. We traced new ultimate controllers for 3,387 of these private firms. Though in our search we may find information about multiple links in the chain of control, our dataset captures only the immediate shareholders and the ultimate controller and does not record further intermediate links.

In 2012, the full dataset contains 27,315 listed firms in 126 jurisdictions.[6] To ensure reasonable coverage across countries and meaningful country statistics, we drop:

• Companies with a market capitalisation below 1 million U.S. dollars. Doing so, we lose 956 companies from 48 (typically very small) jurisdictions.

• Companies for which our database registers aggregate ownership stakes of one percent or less. This drops 300 companies from 49 jurisdictions.

• Companies from jurisdictions with ten or fewer public companies. This leads to the loss of 113 listed companies from 40 jurisdictions.[7]

• Ownership stakes held by entities from jurisdictions when shareholders from those jurisdictions hold stakes in ten or fewer

6. Throughout the paper, we use jurisdiction and country as synonyms.

7. These are: Anguilla, Bahamas, Barbados, Belize, Benin, Botswana, Cambodia, Cameroon, Curaçao, Ecuador, Faroe Islands, Gabon, Gambia, Georgia, Gibraltar, Iraq, Isle of Man, Jamaica, Jersey, Kazakhstan, Kyrgyzstan, Lebanon, Liberia, Liechtenstein, Macao, Malawi, Monaco, Mongolia, Namibia, Niger, North Macedonia, Palestinian Territories, Panama, Papua New Guinea, Rwanda, Senegal, Sudan, Tanzania, Trinidad & Tobago, and Uganda.

companies. This excludes 56 jurisdictions from statistics related to direct ownership stakes.[8]

Companies whose controller is from a jurisdiction that controls five or fewer distinct companies. This drops 75 companies from 48 jurisdictions and drops 37 jurisdictions as controllers.[9]

1.1.2 Sample

The final sample consists of 25,884 public firms located in 86 jurisdictions in 2012. These countries represent approximately 96 percent of global GDP. Our sample accounts for about 87 percent (81 percent) of the total global market cap in Datastream (World Bank). There are 81,192 distinct shareholders; we have information on the nationality of percent of these, accounting, however, for the overwhelming majority of equity (percent). Shareholders come from 90 jurisdictions. We have 8,048 unique ultimate controllers; we have information on the nationality of percent of these, accounting for 97 percent of the controlled market capitalisation, and they come from 81 jurisdictions. The combined market capitalisation is USD 41.35 trillion, and the database captures about half (19.62 trillion) of the value of the voting right stakes. There is strong home bias, as domestic entities hold stakes worth around USD 13.88 trillion.

8. The dropped jurisdictions are Algeria, Andorra, Angola, Anguilla, Azerbaijan, Barbados, Belize, Benin, Botswana, Brunei, Burkina Faso, Cambodia, Congo - Kinshasa, Costa Rica, Côte d'Ivoire, Czechia, Ecuador, Gambia, Ghana, Gibraltar, Guinea-Bissau, Iran, Iraq, Jamaica, Jersey, Kazakhstan, Kyrgyzstan, Macao SAR China, Madagascar, Malawi, Moldova, Mongolia, Mozambique, Myanmar (Burma), Namibia, Nepal, New Caledonia, North Korea, North Macedonia, Papua New Guinea, Puerto Rico, São Tomé & Príncipe, Seychelles, St. Kitts & Nevis, St. Lucia, St. Vincent & Grenadines, Sudan, Tajikistan, Tanzania, Togo, Trinidad & Tobago, Uganda, Uruguay, Vanuatu, Yemen, and Zambia. There are 167 affected listed firms; these firms are however not fully dropped from the sample (only the stakes from shareholders from these countries), as the goal is only to avoid computing statistics on ownership and control of countries with little representation in the sample.

9. These are Andorra, Angola, Azerbaijan, Bahamas, Barbados, Botswana, DR. Congo, Côte d'Ivoire, Curaçao, Ecuador, Ethiopia, Gibraltar, Iran, Jamaica, Kazakhstan, Kenya, Kyrgyzstan, Lebanon, Liberia, Libya, Liechtenstein, Marshall Islands, Monaco, Mongolia, Myanmar, Namibia, North Macedonia, Palestinian Territories, Panama, Seychelles, Sierra Leone, Tanzania, Togo, Trinidad & Tobago, Uruguay, Zambia, and Zimbabwe.

1.1.3 Shareholder Types

BvD classifies entities into 19 types, which we aggregate into six major categories:

• Bank: Banks

• Nonbank Finance: Financial companies; insurance companies; mutual & pension funds / nominees / trusts / trustees; private equity firms; venture capital; hedge funds

• Industry: Industrial companies

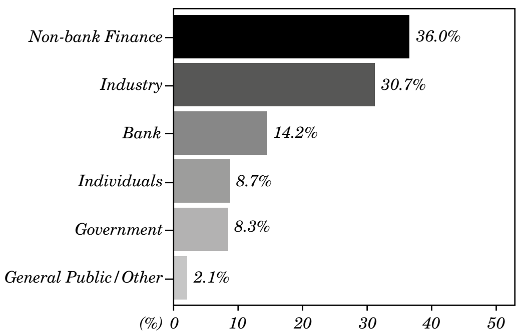
• General Public / other: Foundations / research institutes; public; other unnamed shareholders, aggregated; branches; marine vessels

• Government: Public authorities, states, governments

• Individuals: Individuals; employees / managers / directors; self-ownership; unnamed private shareholders, aggregated

Figure 1 shows the share of equity stakes (controlling and passive) held by each major shareholder type. Nonbank finance and industrial companies are the largest shareholders, each holding around one-third of the equity stakes in our sample.

Figure 1. Share of the Market Capitalisation Value of the Direct Stakes by each Shareholder Entity Type



Source: Authors' calculations.

Note: The sample consists of 25,884 publicly traded firms located in 86 jurisdictions in 2012.

1.2 Controlled and Widely Held Corporations

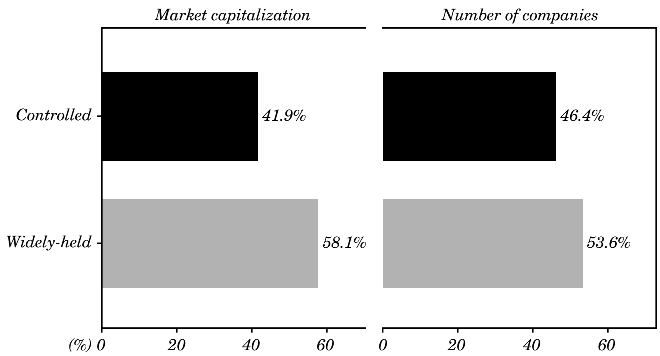
We follow the corporate finance literature and apply a -percent voting right cutoff to identify controlled, as opposed to widely held, companies (e.g. La Porta, Lopez-De-Silanes, Shleifer, and Vishny, 1999).[10] We classify as controlled listed firms where a shareholder (i.e. individual, family, state, another firm, mutual fund) has voting rights over percent. As in Aminadav and Papaioannou (2020), but in contrast to earlier studies, we aggregate the voting rights of all firms that an individual (family or entity) uses to exercise control and aggregate the voting rights of all family members.[11]

Our algorithm identifies 13,864 widely held corporations with a market cap of about USD 24 trillion and 12,020 firms with a controlling shareholder with a market cap of USD 17 trillion. Figure 2 shows the share of controlled and widely held firms in terms of total market capitalisation and the total number of listed firms. Controlled firms are around 42 percent of the market capitalisation and 46 percent of the number of companies. Figure 3 provides the disaggregation across continents and World Bank income groups. Figure 4 tabulates the share of market capitalisation and the number of companies controlled by entities of each type. Despite individuals and families being a minority in ownership stakes (figure 1), they are the controllers of the majority of firms and control a plurality of market capitalisation. Governments control a similar share of market capitalisation with a much smaller share of the number of companies, as they control large companies.

10. Corporate finance research has employed various cutoffs; for example, Lins and others (2013) employ a cutoff, while Laeven and Levine (2008) use . In Aminadav and Papaioannou (2020) we also estimated Shapley-Shubik voting right power measures that incorporate information of all (main) shareholders (Shapley and Shubik, 1954; and Banzhaf, 1965). This alternative metric is useful for the cases where ownership is dispersed and a majority of investors are small or passive, leading stakes smaller than 20% as effective controllers. The 20% cutoff rule yields are quite similar to the Shapley-Shubik method binary classifications of controlled firms that do not matter much when we aggregate at the country(pair) level. Corporate finance studies often distinguish between widely held firms with and without equity blocks, typically over 5% of firm's voting and cash-flow rights. We abstract from this distinction as our focus is on corporate control.

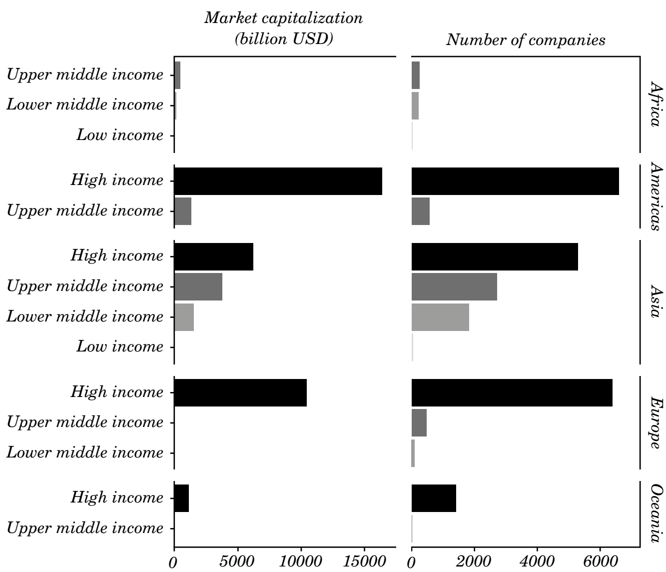
11. In Fiat and BMW, for example, we add the voting shares of all the Agnellis and Quandts.

Figure 2. Share of Controlled and Widely Held Listed Companies in 2012



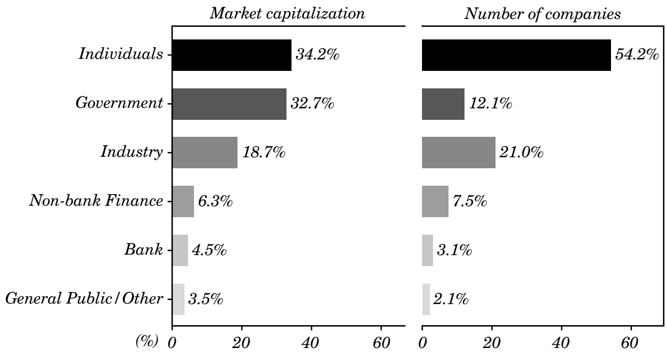
Source: Authors' calculations.
Note: The sample includes 25,884 companies in 86 countries and jurisdictions.

Figure 3. Market Capitalisation and Number of Controlled Firms (with a Shareholder Entity Holding Voting Rights in Excess of 20 Percent) across Income Groups and Regions, Following the Classification of the World Bank



Source: Authors' calculations.

Figure 4. Share of Listed Companies Controlled by each Major Entity Type, as a Share of Total Market Capitalisation and Total Number of Companies



Source: Authors' calculations.
Note: The sample contains 12,020 controlled companies with a total capitalisation of USD 17.3 trillion in 2012.

1.3 International Control

We distinguish between three nationality types for the ultimate controlling entity and for the nationality of the *immediate* controlling (shareholder) entity: (a) domestic, (b) foreign (non-tax-haven), (c) tax-haven (foreign), combining the OECD (2000) list and the classification of Tørsløv and others (2018), which is based on Hines and Rice (1994) and adds Belgium and the Netherlands.[12] Below, we report examples of these different cases.

1.3.1 Widely Held (Noncontrolled)

MercadoLibre Inc., an Argentine company operating online marketplaces is an example of a widely held listed corporation, as its largest shareholder, eBay, held below 20 percent of voting rights (18.4 percent). Marcos Galperin, the company's founder, held a 10.3-percent

12. The jurisdictions in the union of the two classifications are Andorra, Anguilla, Bahamas, Bahrain, Barbados, Belgium, Belize, Bermuda, British Virgin Islands, Cayman Islands, Curaçao, Cyprus, Gibraltar, Hong Kong SAR China, Ireland, Isle of Man, Jersey, Jordan, Lebanon, Liberia, Liechtenstein, Luxembourg, Macao SAR China, Malta, Marshall Islands, Mauritius, Monaco, Nauru, Netherlands, Panama, Samoa, San Marino, Seychelles, Singapore, St. Kitts & Nevis, St. Lucia, St. Vincent & Grenadines, Switzerland, and Vanuatu.

stake, while the remaining shareholders are mostly American investment companies.

Figure 1.3.1

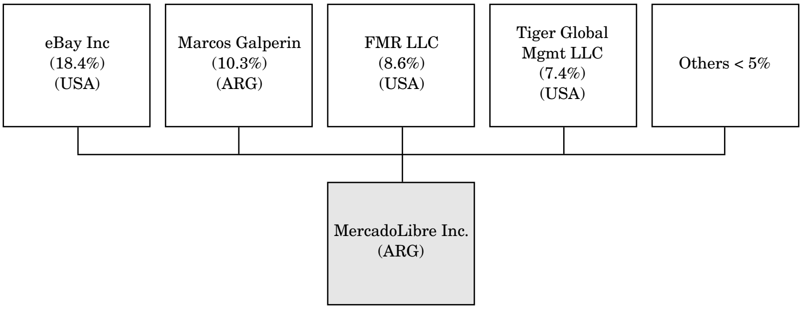


Figure 1.3.2

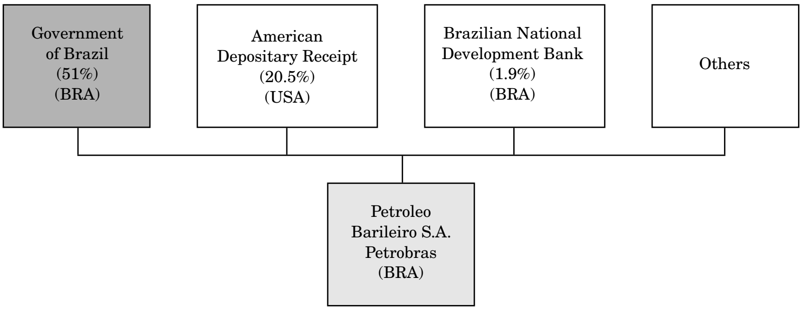


Figure 1.3.3

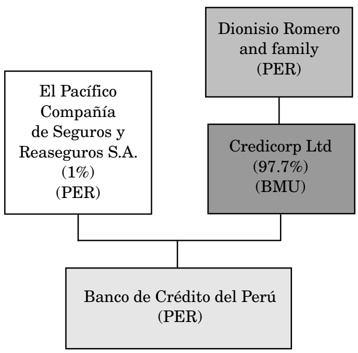
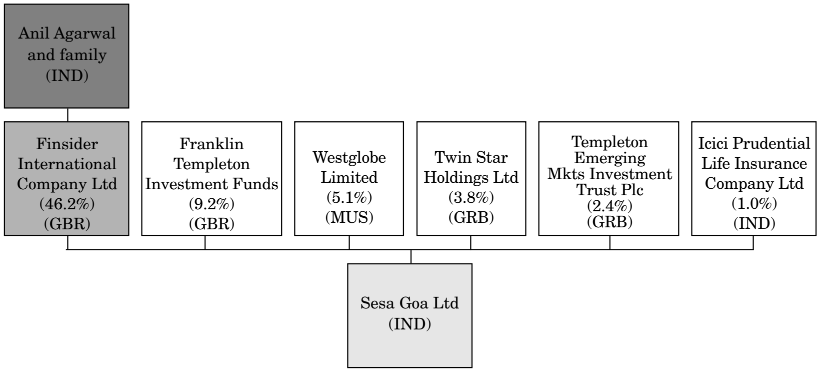


Figure 1.3.4



1.3.2 Domestic Control through Domestic Intermediate or Direct Shareholding

More often than not, listed firms are controlled by a domestic-entity resident either directly or via a local firm. Petrobras, the Brazilian oil and gas giant is an example. The Brazilian government holds an equity stake of above 50 percent. A 20.5-percent stake exists in the form of an American Depository Receipt, which allows the stock to trade in U.S. financial markets.

1.3.3 Domestic Control through Tax Haven

Some firms are controlled by local residents, but the control equity stake goes via an intermediate company, incorporated in financial offshore centres. Banco de Crédito del Perú is an example. The main shareholder, Credicorp Ltd, is incorporated in the Bermuda Islands. This company is in turn owned and controlled by Peruvian citizen Dionisio Romero and his family. A minor stake in the company is held by El Pacífico, a Peruvian insurance company, which is also controlled by Credicorp Ltd.

1.3.4 Domestic Control through Foreign Entity (Non-Tax-Haven)

Often locals control domestic listed corporations by using foreign intermediate firms, which are not necessarily incorporated in tax-haven jurisdictions. Sesa Goa Ltd, an Indian mining company, is an example. The main shareholder is Finsider International Company, a U.K.-based entity, which owns 46.2 percent. Finsider is in turn owned and controlled by Anil Agarwal and his family, also from India. The other main shareholders of Sesa Goa are investment companies from the U.K., Mauritius, and India.

1.3.5 Foreign Tax-Haven Control through Domestic Intermediary

In some cases, firms incorporated in tax havens will have controlling equity stakes in listed corporations by using an intermediate domestic firm. PLDT Communication and Energy Ventures is a listed company on the Philippine Stock Exchange in the communication and energy sectors. In 2012, it was wholly owned by Smart Communications, another Philippine entity, whose controlling shareholder was First Pacific, a Hong-Kong-based and listed investment and management company.

1.3.6 Foreign Tax-Haven Control through Foreign Tax-Haven or Direct Shareholding

It is not uncommon that control exerted by a company in a financial offshore centre is intermediated via a company in another tax-haven jurisdiction. PT Astra International, Tbk. is an Indonesian conglomerate that operates in several sectors, in particular in the automotive industry. Our dataset records a majority stake owned by Jardine Cycle & Carriage, a Singaporean entity, which is ultimately owned by Jardine Strategic Holdings, a Hong-Kong-based entity founded in the 19th century.

Figure 1.3.5

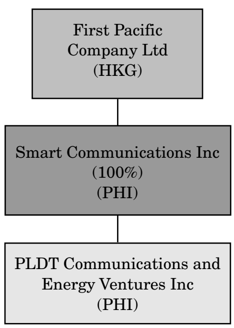


Figure 1.3.6

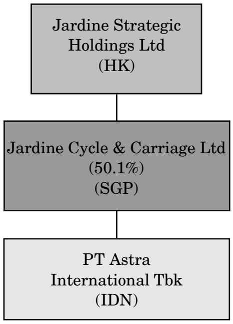
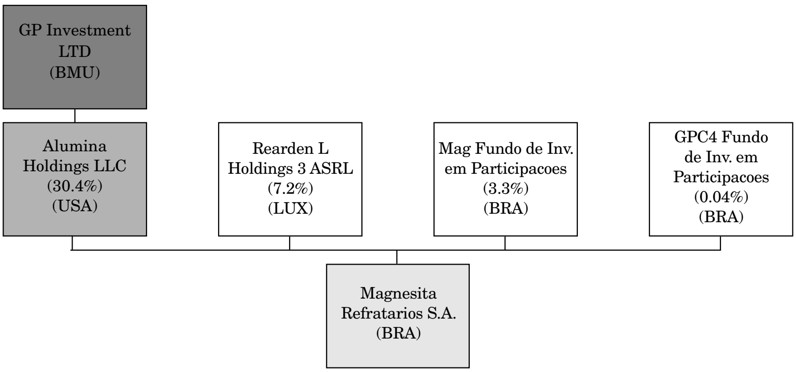


Figure 1.3.7



1.3.7 Foreign Tax-Haven Control through Foreign Non-Tax-Haven

Sometimes controlling equity chains operate via many companies, incorporated both in foreign countries and foreign tax-haven jurisdictions. Take, for example, Magnesita Refratários, a Brazilian company in the refractory industry. Its controlling equity stake is held by Alumina Holdings LLC, a Delaware-based entity,[13] but the intermediate firm is owned by GP Investments LTD, a Bermuda-based entity.

13. Despite the potential classification of Delaware-registered companies as tax-haven companies, our dataset does not allow us to distinguish between different states in the United States.

Figure 1.3.8

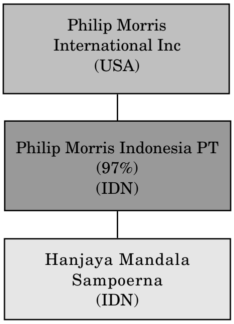


Figure 1.3.9

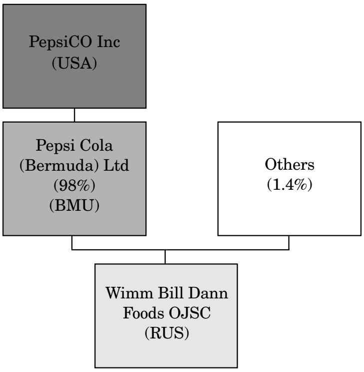
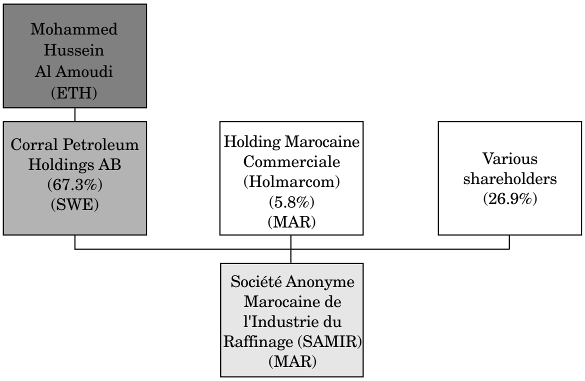


Figure 1.3.10



1.3.8 Foreign (Non-Tax-Haven) Control through Domestic

Often foreign controlling shareholders channel their controlling equity positions via domestic firms. For example, Hanjaya Mandala Sampoerna, an Indonesian tobacco company is owned (97 percent-equity stake) via Philip Morris Indonesia PT, the local subsidiary of Philip Morris International.

1.3.9 Foreign (Non-Tax-Haven) Control through Tax Haven

Often large multinationals and other foreign investors will use an intermediary firm incorporated in a tax-haven jurisdiction. Wimm Bill Dann Foods OJSC, a Russian dairy company is controlled by PepsiCo Inc, the American giant, via a Bermuda-incorporated subsidiary, Pepsi Cola Bermuda Ltd.

1.3.10 Foreign Non-Tax-Haven Control through Foreign Non-Tax-Haven or Direct Shareholding

The final group is for firms held through foreign non-tax-haven entities. Société Anonyme Marocaine de l'Industrie du Raffinage (SAMIR) is a Moroccan firm specialised in refining of petroleum products. Our dataset listed Swedish holding company Corral Petroleum Holdings AB as its main shareholder, holding a stake of 67.3 percent. In addition, a Moroccan holding and various other unidentified shareholders are registered. Corral Petroleum is ultimately held by Ethiopian-Saudi billionaire Mohammed Hussein Al Amoudi.

1.4 Corporate Control across Countries

1.4.1 Measures

As we analyse countries, we discuss the construction of international corporate ownership and control statistics across source and destination countries, by using Argentina as an example. We define the following measures of international corporate ownership and control.

- Cross-border Ownership:
  - Value of direct equity stakes by entities from source jurisdiction in public companies of destination jurisdiction (% of voting stake market capitalisation).

- International Corporate Control:
  - Value of listed firms (market capitalisation) ultimately controlled by entities from source jurisdiction in destination jurisdiction.
  - As our focus is on control, we compile four measures:
    - market capitalisation amount (billion U.S. dollars) and share of total market capitalisation;
    - number of companies and share of total listed companies controlled.

1.4.2 Example: Argentina

Companies in Argentina

Our dataset records 76 companies based in Argentina, 75 of them listed on the local stock exchange, and one listed in the United States, with a total market capitalisation of USD 32 billion in 2012. We classify 71 as controlled, as there is a shareholder (domestic, foreign, or tax-haven) with voting rights in excess of 20 percent. The remaining companies are widely held. The total market capitalisation of controlled firms is USD 26 billion. We assign controlled companies into nine groups (examples above) according to the combination of the ultimate and the main direct shareholder:

- 25 controlled by an Argentine entity, worth USD 13.38 billion.
  - 25 controlled by an Argentine entity through an Argentine entity, worth USD 13.38 billion.
  - 0 controlled by an Argentine entity through a foreign entity.
  - 0 controlled by an Argentine entity through a tax-haven entity.
- 24 controlled by a foreign entity, worth USD 9.54 billion.
  - 10 controlled by a foreign entity through an Argentine entity, worth USD 1.96 billion.
  - 13 controlled by a foreign entity through a foreign entity, worth USD 5.88 billion.
  - 1 controlled by a foreign entity through a tax-haven entity, worth USD 1.7 billion.
- 3 controlled by a tax-haven entity, worth USD 404 million.
  - 1 controlled by a tax-haven entity through an Argentine entity, worth USD 148 million.
  - 2 controlled by a tax-haven entity through a foreign entity, worth USD 256 million.
  - 0 controlled by a tax-haven entity through a tax-haven entity.

• There are 19 domestic listed corporations, worth USD 2.75 billion for which we lack enough information about the nationality of the major entities in the control chain.

Companies controlled by Argentine entities

Argentine entities (individuals/families, banks, government, industry, nonbank finance) control 30 companies worth USD 14 billion.

• 25 domestic firms, worth USD 13.38 billion
– 25 domestic firms controlled through a domestic entity, worth USD 13.38 billion.
– 0 domestic firms controlled through a foreign entity.
– 0 domestic firms controlled through a tax-haven entity.
• 5 foreign firms, worth USD 662 million
– 3 foreign firms controlled through a domestic entity, worth USD 353 million.
– 0 foreign firms controlled through a foreign entity.
– 2 foreign firms controlled through a tax-haven entity, worth USD 309 million.
• 0 tax-haven firms
– 0 tax-haven firms controlled through a domestic entity.
– 0 tax-haven firms controlled through a foreign entity.
– 0 tax-haven firms controlled through a tax-haven entity.

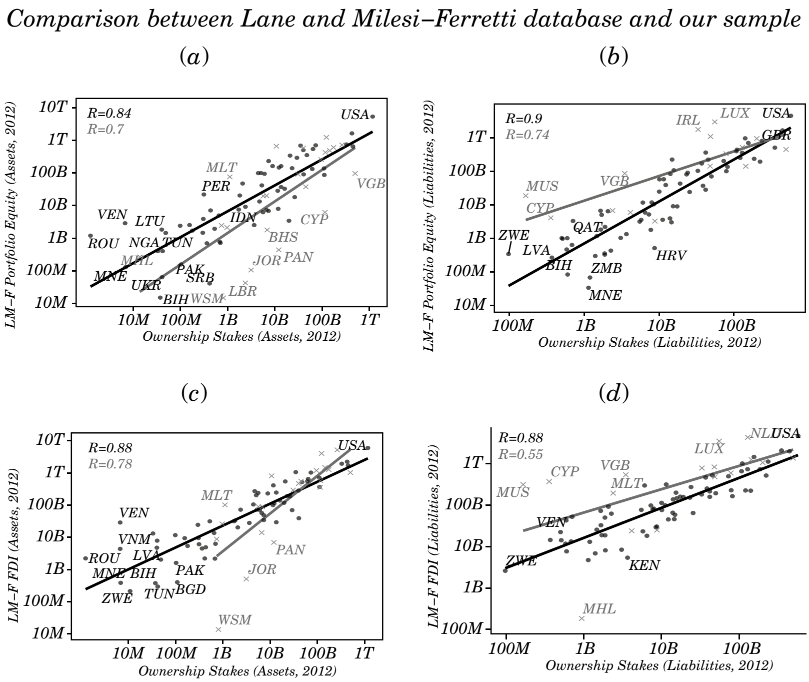
1.5 Comparison with Other Datasets

1.5.1 External Wealth of Nations, Lane and Milesi-Ferretti.

It is instructive to compare the newly complied country-aggregate stakes in international corporate ownership and control with the widely used data of Lane and Milesi-Ferretti (2007, 2018, 2021) on the external wealth of nations. Relying on multiple sources (individual countries, international organisations such as the IMF, the World Bank, and the Bank for International Settlements, and other research), Lane and Milesi-Ferretti (2007, 2018, 2021) provide annual country-level statistics of external financial assets and liabilities (based on the residence principle) for 212 economies, distinguishing between foreign direct investment (FDI, controlling equity stakes), portfolio investments (bonds and equity), financial derivatives, and foreign exchange reserves (held by the national central banks).

Figure 5 plots the cross-country correlation between the Lane and Milesi-Ferretti portfolio and FDI measures and our statistics of cross-border ownership stakes (both controlling and noncontrolling) in listed companies in 2012. Panels (a) and (b) compare with Lane and Milesi-Ferretti's portfolio equity measures, while panels (c) and (d) compare with FDI measures. Panels (a) and (c) look at foreign financial assets, taking a source country (i.e. the owner's) viewpoint in our data, while panels (b) and (d) examine the correlation between foreign liabilities and ownership stakes at the destination country (i.e. the firm's). Each panel plots the correlation across non-tax-haven jurisdictions (dark line) and tax havens (light line).

Figure 5. International Ownership of Listed Corporations vs External Wealth of Nations (Assets and Liabilities), Lane and Milesi-Ferretti (2018, updated in 2021)



Source: Authors' calculations, and Lane and Milesi-Ferretti (2018, updated in 2021).
Note: Values in current U.S. dollar; Positions < 10 M USD dropped. R indicates the correlation coefficient.

Table 1. Comparison with other Datasets

	Ownership stakes				
	Assets		Liabilities		Assets in tax havens
	Model 1	Model 2	Model 3	Model 4	Model 5
(Log) LM-F Portfolio Equity (Assets, 2012)	0.820***				
	(0.189)				
(Log) LM-F FDI (Assets, 2012)		1.269***			
		(0.187)			
(Log) LM-F Portfolio Equity (Liabilities, 2012)			0.418***		
			(0.083)		
(Log) LM-F FDI (Liabilities, 2012)				0.559***	
				(0.123)	
(Log) AJZ Total Offshore Wealth (2007)					0.102
					(0.116)
Tax haven	0.180	-1.350**	0.890+	0.494	
	(0.471)	(0.511)	(0.471)	(0.536)	
(Log) GNI per capita	-0.259	-0.408	0.224+	0.293*	1.121***
	(0.458)	(0.299)	(0.121)	(0.125)	(0.196)
(Log) Population	-0.011	-0.222+	0.382***	0.357**	0.981***
	(0.230)	(0.134)	(0.093)	(0.109)	(0.091)
Num. Obs	81	83	84	84	63
Adj. Pseudo R-Sq	0.691	0.860	0.825	0.804	0.941

Source: Authors' calculations.

Note: The table reports cross-country Poisson Pseudo-Maximum Likelihood (PPML) regressions. The PPML was chosen due to the use of a dependent variable in logs with 0 values. The coefficients should be read as elasticities. Columns (1)-(4) compare measures in our dataset with data from Lane and Milesi-Ferretti (2021), while column (5) compares with data from Alstadsæter and others (2018). In columns (1)-(2), the dependent variable is the aggregate value of ownership stakes owned by shareholders of a given country in foreign firms, i.e. assets of the country. In columns (3)-(4), the dependent variable is the aggregate value of ownership stakes owned by foreign shareholders in the public firms of a given country, i.e. liabilities of the country. In column (5), the dependent variable is the aggregate value of ownership stakes owned by shareholders of a given country in companies incorporated in tax havens. Heteroskedasticity-adjusted standard errors are reported below the estimates. + p < 0.1, * p < 0.05, ** p < 0.01, *** p < 0.001.

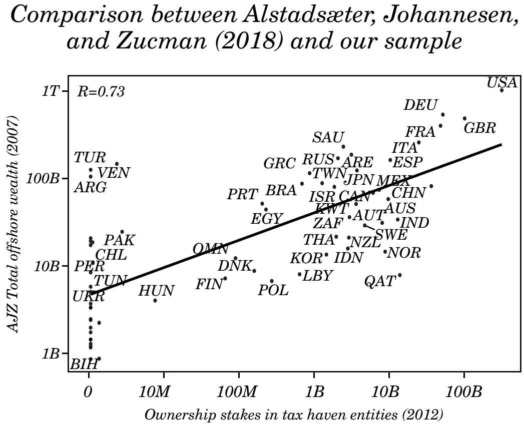
The following patterns emerge. First, the two series are strongly correlated across all measures and groups of countries, with the correlation coefficient ranging between and , when we set aside financial offshore centres. Second, the correlation is still strong (about 0.55 – 0.78) even when restricting attention to tax havens, despite evident difficulties in properly measuring ownership and the non-negligible measurement error. Third, the correlations retain their economic and statistical significance when we control for country size, (log) population, and (log) GNI per capita (table 1).

Alstadsæter and others (2018)

Data on tax havens are scant, although recently there has been increasing information (Zucman, 2013). Alstadsæter and others (2018) try to approximate countries' total wealth held in financial offshore centres by combining scattered information that has become available. In particular, they merge newly disclosed bilateral data from some prominent offshore centres with data from deposits of foreigners in Swiss banks and "errors and omissions" in aggregate country assets and liabilities to approximate the amount of wealth held offshore.

We thus explored how our estimates of ownership stakes in listed corporations in 2012 channelled via financial offshore centres (from a source-country viewpoint) correlate with their approximation of the total offshore wealth in 2007. Figure 6 plots the cross-country correlation (dropping offshore centres), while column (5) in table 1 reports Poisson Pseudo-Maximum Likelihood estimates. While the unconditional correlation is considerable, it weakens and turns statistically indistinguishable from zero once we simply condition on population and GNI (Gross National Income) per capita. There are some important differences between the two series, which future research should delve into. Our corporate ownership of listed companies' data suggests a very small use of financial offshore centres in Turkey, Venezuela, Argentina, and Pakistan; this is however not the case in the estimates of Alstadsæter and others (2018), which however mostly reflect cross-border bank holdings and deposits.

Figure 6. Comparison with Estimates of Wealth Data in Tax Havens. Alstadsæter and Others (2018)



Source: Authors' calculations, and Alstadsæter, Johannesen, and Zucman (2018).
Note: Values in current U.S. dollars indicate the correlation coefficient.

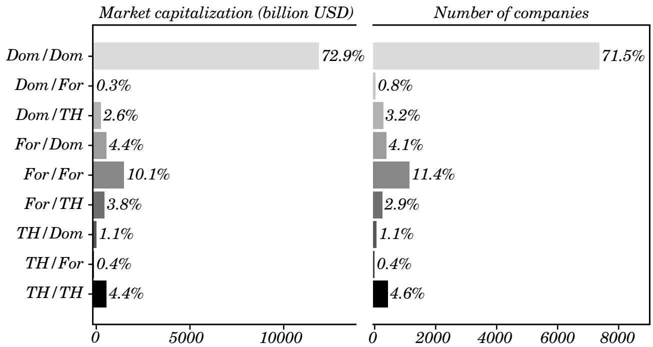
2. International Corporate Control Patterns. Cross-Border Links

This section presents the main patterns of the internationalisation of corporate control by using the newly assembled data. First, we present the main patterns of cross-border corporate control in 2012. Second, we examine differences across income and explore the role of country size.

2.1 Cross-Border Corporate Control. Main Patterns

Figure 5 plots the breakdown of controlled firms across the nationality of the ultimate shareholder and the immediate shareholding entity across the world. The controlling shareholder in the majority of firms, about 75 percent, is a domestic entity (family/individual, government, banks), telling of a strong home bias. Non-domestic entities, located in a foreign country or a tax-haven jurisdiction, control about 25 percent. The most common control chain is domestic, but there is significant control exerted through foreign entities, including tax havens. The usage of tax haven as the direct shareholder is used in the same order of magnitude by domestic and foreign controllers.

Figure 7. Share of the Different Types of Control Chains among Controlled Firms, Worldwide



Source: Authors' calculations.
Dom indicates a domestic shareholder or controlled.
For indicates a foreign non-tax haven.
TH indicates a foreign tax haven. E.g. Dom / TH indicates that the controller is domestic, and the main shareholder is from a foreign tax haven.

Figures 6 and 7 plot the share of total market capitalisation of controlled firms, by the three types of the controlling shareholder entity: domestic, foreign, and foreign tax-haven for each jurisdiction, grouped by income level. Foreign control differs considerably across the world.

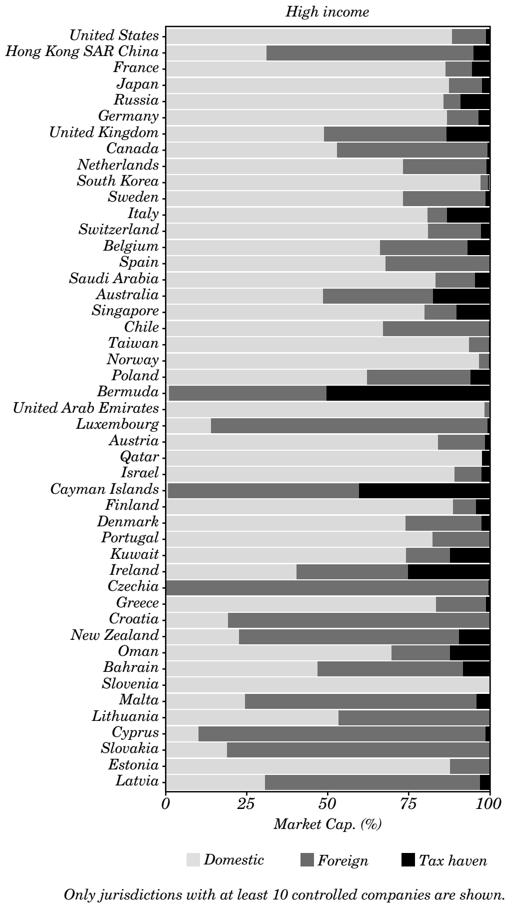
• Foreign control of listed corporations (a shareholder holding more than 20 percent voting rights) exceeds percent, sometimes significantly, in many African counties, like South Africa, Morocco, Nigeria, Kenya, Tunisia, Côte d'Ivoire, and Ghana, and parts of the former transition countries in Eastern Europe, like Czechia, Romania, Slovakia, Croatia, Serbia, Montenegro, Latvia, Bulgaria, and Ukraine.

• Control by foreign shareholding entities hovers between around percent and around percent in large emerging markets, like Brazil, Indonesia, Thailand, Malaysia, Turkey, Philippines, and Egypt, and among high-income countries in the United Kingdom, Canada, Australia, Ireland, Sweden, Spain, Chile, and Poland.

• Foreign control is low in countries across regions and income levels, such as China, Colombia, the United States, the United Arab Emirates, Qatar, South Korea, and Norway.

Appendix tables 6 and 7 provide the detailed statistics of corporate control across the 86 destination countries, distinguished by the nationality of the immediate and the controlling shareholder.

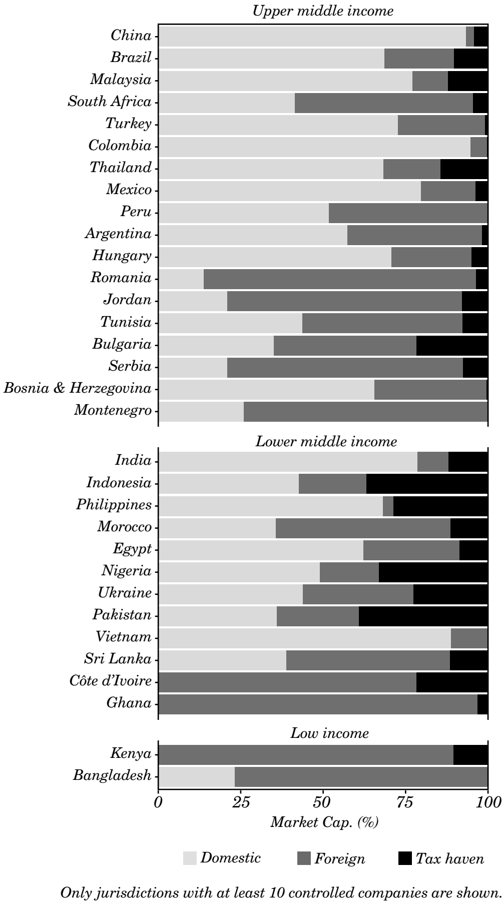
Figure 8. Nationality of Controllers in High-Income Countries



Source: Authors' calculations.

Note: Only jurisdictions with at least 10 controlled companies are shown.

Figure 9. Nationality of Controllers in Non-High-Income Countries

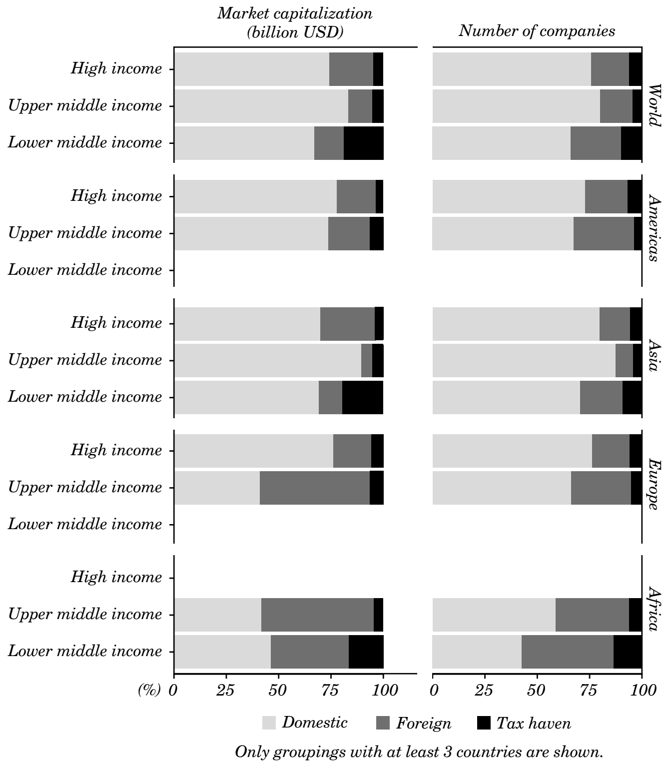


Source: Authors' calculations.
Note: Only jurisdictions with at least 10 controlled companies are shown.

2.2 Income, Population, and International Corporate Control

Figure 10 aggregates the nationality of the controller at the continent and income group levels. International control is higher in lower-middle-income countries, as compared to high and upper-middle-income nations. Foreign control is particularly frequent in middle-income countries in (Eastern) Europe and Africa.

Figure 10. Nationality of Controller across Continent and Income Levels

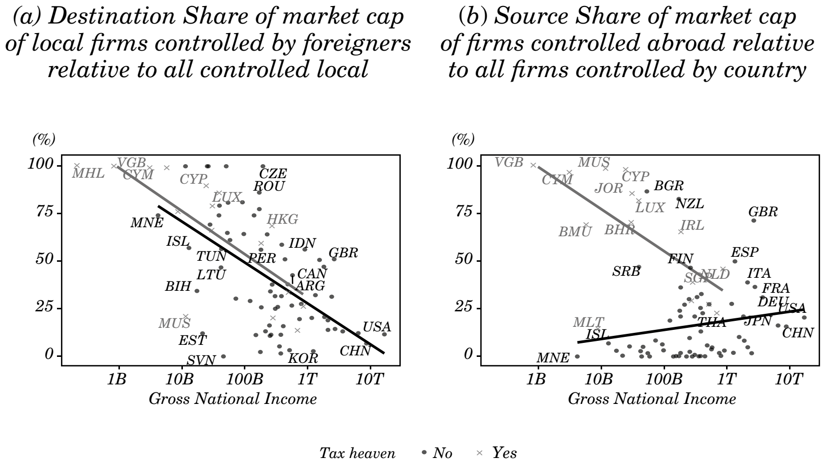


Source: Authors' calculations.
Note: Only groupings with at least three countries are shown.

To explore more formally the correlation between country control internationalisation and income, we run simple cross-country regressions linking the share of controlled firms (in terms of market capitalisation and number of companies) with development, as proxied by (log) income (per capita) and (log) population.[14]

14. Numerous studies show that size, besides trade in goods, is also related to capital flows and holdings, such as foreign direct investment and bank lending (e.g. Alfaro and others, 2008). Rose and Spiegel (2004) connect trade and asset flows, while Hau and Rey (2008) develop a risk-diversification model stressing the role of size. While we do not run country-pair regressions (as in Fonseca and others, 2022), exploiting the bilateral nature of our data, we distinguish between companies incorporated into destination countries and the positions of controlling shareholders from source countries.

Figure 11. Size and Cross-Border Corporate Control



Source: Authors' calculations.
Note: Panel A plots the share of the total market capitalisation in all controlled firms at destination controlled by foreign entities (individuals, families, banks, financial institutions, and so on) against countries' GNI. Panel B plots the share of the total market capitalisation in all firms, controlled and widely held, at destination against GNI. Square dots indicate tax-haven jurisdictions.

Table 2 panel A gives the results of linking openness in corporate control and ownership and size from a destination-country viewpoint, i.e. the jurisdiction of the listed company. Columns (1)–(2) look at the market capitalisation of controlled companies by foreigners as the share of the total market capitalisation of controlled firms, while in columns (5)–(6), the dependent variable is the share of the number companies controlled by foreign entities firms relative to the total number of listed controlled firms. Columns (3)–(4) and (7)–(8) examine the link between corporate ownership and size, looking at ownership links by foreign entities in public corporations in destination, not necessarily linked to control. Size is a strong correlate of the internationalisation of corporate control, as both (log) GNI (incl. per capita) and log (population) enter with significantly negative estimates, revealing that foreign control is more prevalent in smaller countries. Figure 9 panel A illustrates the strong inverse relation between cross-border corporate control and the size of the economy. This result echoes the inverse link between trade (exports and imports) and financial openness (capital inflows and outflows), and size, development, and population.

Table 2. Size, Development, and Cross-Border Corporate Control and Ownership

	Panel (A) Destination							
	Market Cap				Number of companies			
	Share of foreign-controlled firms in all controlled firms		Share of stakes in foreign firms among all recorded stakes		Share of foreign-controlled firms in all controlled firms		Share of foreign firms in all firms with a stake	
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
Log GNI	-0.099***		-0.073***		-0.072***		-0.051+	
	(0.010)		(0.010)		(0.011)		(0.026)	
Log GNI per cap.		-0.132***		-0.111***		-0.065***		0.030
		(0.025)		(0.022)		(0.018)		(0.043)
Log Population		-0.084***		-0.062***		-0.058***		-0.058*
		(0.012)		(0.012)		(0.011)		(0.027)
Num. Obs	85	85	85	85	85	85	85	85
Adjusted R^2	0.406	0.448	0.282	0.346	0.290	0.430	0.030	0.164
Fixed Effects	Continent		Continent		Continent		Continent	

Source: Authors' calculations.
Note: The table reports cross-country OLS regressions from the perspective of the destination country, i.e. the incorporation country of a company. Columns (1)–(4) refer to measures with market capitalisation. Columns (5)–(8) refers to measures of the number of companies. The dependent variable in columns (1)–(2) is the share of the market capitalisation of foreign-controlled firms in all controlled firms in a country. The dependent variable in columns (3)–(4) is the share of the value of ownership stakes held by foreign entities in a country. The dependent variable in columns (5)–(6) is the share of the number of foreign-controlled firms in all controlled firms in a country. The dependent variable in columns (7)–(8) is the share of the number of firms that have at least one foreign entity as a shareholder. Specifications include continental fixed effects when indicated (constants not reported). Heteroskedasticity-adjusted standard errors are reported below the estimates. + $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 2 panel B reports the results taking a source-country viewpoint, i.e. the jurisdiction of the controller/shareholder. The dependent variable in columns (1)–(2) (and (5)–(6)) is the share of market capitalisation (number of) controlled companies abroad in the total of all companies controlled by entities of the source countries. Columns (3)–(4) and (7)–(8) repeat the analysis by looking at ownership links abroad (in terms of market capitalisation and the number of firms), without necessarily a controlling stake. Motivated by the pattern in panel B of figure 9, which shows strikingly different patterns for tax havens, we include a tax-haven dummy and its interaction with GNI. Overall, the size of the economy appears negatively correlated, but this is mainly driven by tax havens, and smaller tax havens in particular. Once these factors are controlled for, we see that (Log) GNI per capita enters with a significantly positive estimate showing that residents in rich countries hold relatively larger equity stakes abroad, both controlling and passive, while population is not a significant predictor.

Table 2. Size, Development, and Cross-Border Corporate Control and Ownership

	Panel (B) Source							
	Market Cap				Number of companies			
	Share of foreign-controlled firms in all controlled firms		Share of stakes in foreign firms among all recorded stakes		Share of foreign-controlled firms in all controlled firms		Share of foreign firms in all firms with a stake	
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
Log GNI	-0.049*		-0.077***		-0.031		-0.064***	
	(0.021)		(0.018)		(0.019)		(0.014)	
Log GNI per cap.		0.054*		0.082***		0.087***		0.103***
		(0.023)		(0.021)		(0.017)		(0.021)
Log Population		0.025		0.002		0.019+		0.002
		(0.015)		(0.013)		(0.011)		(0.014)
Tax haven		2.029***		1.339***		1.248**		1.037***
		(0.511)		(0.361)		(0.383)		(0.266)
Tax haven X Log GNI		-0.143***		-0.078*		-0.080*		-0.056*
		(0.040)		(0.032)		(0.032)		(0.023)
Num. Obs	78	78	89	89	78	78	89	89
Adjusted R^2	0.091	0.537	0.231	0.678	0.050	0.644	0.184	0.691
Fixed Effects				Continent				Continent

Source: Authors' calculations.
Note: The table reports cross-country OLS regressions from the perspective of the source country, i.e. the shareholder or controller of a company. Columns (1)–(4) refer to measures with market capitalisation. Columns (5)–(8) refers to measures of the number of companies. The dependent variable in columns (1)–(2) is the share of the market capitalisation of foreign-controlled firms in all controlled firms in a country. The dependent variable in columns (3)–(4) is the share of the value of ownership stakes held by foreign entities in a country. The dependent variable in columns (5)–(6) is the share of the number of foreign-controlled firms in all controlled firms in a country. The dependent variable in columns (7)–(8) is the share of the number of firms that have at least one foreign entity as a shareholder. Specifications include continental fixed effects when indicated (constants not reported). Heteroskedasticity-adjusted standard errors are reported below the estimates. + $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

3. Tax Havens in International Corporate Control

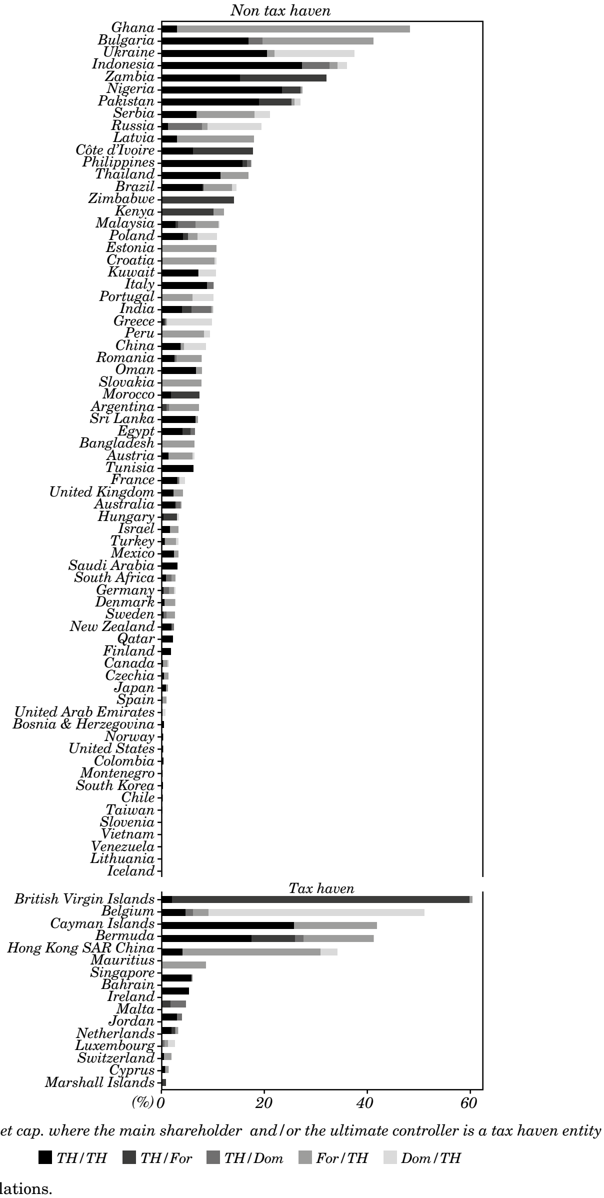
We now zoom in on the role of tax havens in international corporate control. First, we present the major patterns across all sample countries. Second, we examine differences across income group and market size.

3.1 Country Patterns on Tax-Haven Usage

Figure 10 depicts the percentage of total market capitalisation (i.e. including noncontrolled widely held listed firms) in each country where either the controlling entity or the main direct shareholder (or both) are from or incorporated in a tax-haven jurisdiction. There is wide variation in the use of tax-haven entities.

- Tax-haven use is the highest in Eastern Europe, especially in Bulgaria, Ukraine, Serbia, Latvia, and Russia.
- The use of tax-haven-incorporated intermediate vehicles is also considerable for exercising control in many African countries, mostly in Ghana, Zambia, Nigeria, Côte d'Ivoire, and Zimbabwe.
- Tax-haven jurisdiction vehicles are used widely to control listed firms in Indonesia, Pakistan, the Philippines, and other East Asian countries.
- In a few countries, domestic entities, including families and individuals, hold controlling equity stakes in firms listed in the local stock exchange by using intermediate firms incorporated in tax-haven jurisdictions. This pattern is higher in Ukraine, Russia, Greece, and Serbia, as well as in China.
- The use of intermediate firms to exercise control is smaller in countries from a wide range of regions.
- The exercise of control by or via tax-haven-incorporated vehicles appears quite low in the United States. However, while going through manual checks, we observe entities incorporated in Delaware, which has been considered a tax haven (Michel, 2021). Unfortunately, our data do not allow us to distinguish the state of incorporation.

Figure 12. Tax-Haven-Incorporated Vehicles in Corporate Control Chain across Countries

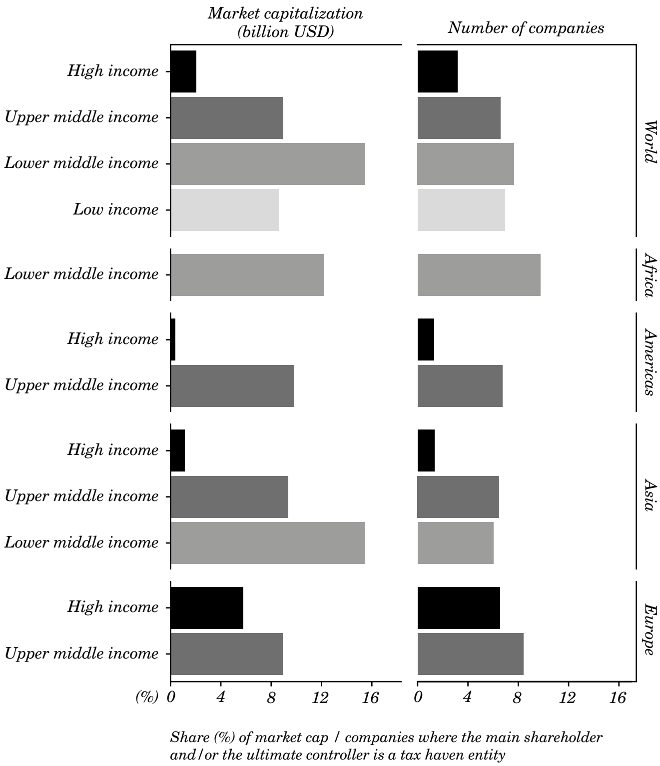


Source: Authors' calculations.
Dom indicates a domestic shareholder or controlled. For indicates a foreign non-tax haven.
TH indicates a foreign tax haven. E.g. Dom / TH indicates that the controller is domestic, and the main shareholder is from a foreign tax haven.

3.2 Differences across Income Group and Size Effects

As in the earlier section, we also examine the role of income and market size in explaining the considerable differences in the use of tax-haven entities in corporate control, either as intermediate vehicles or as ultimate owners. Figure 11 tabulates aggregations at the continent- and income-group levels, excluding public companies directly incorporated in tax havens. In general, lower-income countries have a higher percentage of tax-haven usage in the corporate control chains.

Figure 13. Share of Market Capitalisation and Number of Listed Companies Where the Main Shareholder and/or the Ultimate Controller is an Entity, Incorporated in a Financial Offshore (Tax-Haven) Jurisdiction, across Income Groups and Continents



Source: Authors' calculations.
Note: Only groupings with at least 3 countries are shown. Companies from tax-haven jurisdictions are not counted.

While our focus is not delving into the drivers of tax-haven use, we estimated simple cross-country specifications to further understand the role of market size. Table 1 shows cross-country regression results, associating the use of tax-haven-incorporated firms in the control chain to log population and log GNI per capita. As there are evident regional differences, the specifications include continental constants. For these results, we drop countries classified as tax havens to focus on the usage of offshores in non-tax-haven countries. In columns (1)–(4) we take a 'destination'-country viewpoint, i.e. the country of the public company. The dependent variable in (1)–(2) is the share of domestic market capitalisation and, in (3)–(4), of the listed firms where control passes via companies incorporated in tax havens (the categories shown in figure 10) to the total market capitalisation and number of controlled firms in the local stock market. The estimate on log GNI per capita is negative and highly significant, while the coefficient on log population is both small and statistically indistinguishable from zero. In line with the income-group tabulations, there is some evidence that corporate control in relatively low-income countries operates often by or via entities incorporated in tax havens. The dependent variables in (5)–(8) take a 'source'-country perspective, i.e. what the share of tax-haven usage is in the companies controlled by entities from the source country. The estimates for GNI per capita and population are not precise enough to conclude that there is a strong relation with the use of tax havens to control firms.

Table 3. Size (Population and Income) and the Use of Tax Havens in International Corporate Control

	Destination				Source			
	Market Cap.		Num. Companies		Market Cap.		Num. Companies	
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
Log GNI	-0.008		-0.004		0.004		-0.001	
	(0.007)		(0.004)		(0.008)		(0.006)	
Log GNI per cap.		-0.041**		-0.024**		-0.025		-0.025
		(0.012)		(0.008)		(0.024)		(0.021)
Log Population		0.006		0.005		0.012		0.005
		(0.008)		(0.005)		(0.007)		(0.006)
Num. Obs	66	66	66	66	46	46	46	46
Adjusted R^2	0.043	0.241	0.053	0.220	-0.017	0.078	-0.055	0.048
Fixed Effects	Continent		Continent		Continent		Continent	

Source: Authors' calculations.
Note: The table reports cross-country OLS regressions. The dependent variable in columns (1)–(2) and (3)–(4) is the share of controlled firms at destination where control is exercised by or via firms incorporated in financial offshore (tax-haven) jurisdictions. The dependent variable in columns (5)–(6) and (7)–(8) is the share of controlled firms at source country where control is exercised via firms incorporated in financial offshore (tax-haven) jurisdictions. All specifications include continental fixed effects (constants not reported). Heteroskedasticity-adjusted standard errors are reported below the estimates. + $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

4. Conclusion

Drawing on our parallel work (Fonseca and others, 2022), and the extension, update, and cleaning of the ORBIS data on corporate ownership in Aminadav and Papaioannou (2020), we provide an anatomy of corporate control across more than 25,000 public companies in 2012. Our global mapping of corporate control distinguishes between three nationality types of the immediate shareholder and ultimate controlling entities (domestic, foreign, and foreign tax-haven), and the various types of entities in ownership structures.

The first part of our descriptive analysis reveals considerable differences in cross-border corporate control across countries of company incorporation on one hand, and listed traded exchange (destination) and sizable variation across the main shareholder's countries (source), on the other. International corporate control is relatively high in Eastern Europe and Africa, where foreigners control the majority of listed companies and market capitalisation, but lower in Latin America and East Asia. There are also non-negligible differences even across nearby countries. Control by foreign entities is less significant in larger economies, mirroring the international trade and capital flow patterns. In addition, shareholder entities from wealthier jurisdictions own and control a larger share of holdings abroad.

In the second part of our analysis, we zoom in on financial offshore centres, whose role has come into scrutiny given the recent policy efforts to tax international investors and enhance transparency. We document the importance of shareholder entities in offshore financial centres as conduits of international control. We discuss the wide heterogeneity in the usage of tax havens across and within continents. In some instances, domestic residents use tax-haven-incorporated shells to channel their controlling stakes in domestic listed companies. The use of tax-haven-incorporated vehicles is larger in lower-income economies.

Our mapping of cross-border corporate control raises questions that our ongoing research (Fonseca and others, 2022) examines. First, updating the data backward and forward will allow examining the dynamics of cross-border corporate control and the use of tax-haven-incorporated conduits. Second, a thorough analysis of the drivers of cross-border control is needed, looking at the role of taxation, political institutions, investor protection, and more. Third, by exploring the country-pair structure, we examine the role of cultural, political, and economic ties, the impact of bilateral investment, and trade treaties, also distinguishing by investor type.

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Appendix

Additional Tables

Table A.1 Matrix of Ownership Stakes between Countries from the G7, BRICS, and MINT
(in billion U.S. dollars)

Located/Owned	BRA	CAN	CHN	FRA	DEU	IND	IDN	ITA	JPN	MEX	NGA	RUS	ZAF	TUR	GBR	USA
Brazil	320	1		4	1			1	1	2					7	39
Canada	0	229	3	1	0	3	0	0		0	0	1	0	0	17	58
China		0	1,377	7	2	0	0	0	1				0		11	47
France	0	0	2	624	35	0		10	5				0	0	7	62
Germany		1	3	21	349			4	0			0	0	0	6	69
India		0	1	1	10	477	0	0	3				0		42	23
Indonesia		0		0	5	0	106		1	0				0	11	0
Italy		0	0	7	2			170	0			0	0	0	4	18
Japan			0	25	6		0	0	717						6	65
Mexico		0						0		89					1	21
Nigeria		0		1	0			0			12		0		2	0
Russia		0		16	4				0			400			6	55
South Africa		1	4	0	1	0							120		47	26
Turkey				3	2			6	1					119	5	12
United Kingdom		2	0	39	22	0	0	1	0	13	0	6	5		486	223
United States	1	47	2	200	159	0		2	6	3	0	1	2		33	4,340

Source: Authors' calculations.

Table A.2 Matrix of Controlled Market Capitalisation between Countries from the G7, BRICS, and MINT
(in billion U.S. dollars)

Located/Owned	BRA	CAN	CHN	FRA	DEU	IND	IDN	ITA	JPN	MEX	NGA	RUS	ZAF	TUR	GBR	USA
Brazil	379	1	3	26	0			4		0					26	9
Canada	0	203	8	1	0	11	0	0	2	0	0	4	0	0	20	127
China			2,300	12				1	3			1			24	1
France		0		712	45			5							3	2
Germany			1	10	419	0		9	3						5	10
India				2	31	645		0	12						11	20
Indonesia				1	7	1	116		1					0	2	27
Italy				0				252	0			0				4
Japan			0	53	5	0	0	0	695						2	17
Mexico		0								119						2
Nigeria			0	0	0						12		0		4	0
Russia				5	5			2				594			0	5
South Africa		0	1		1	3							75		66	27
Turkey	0			3	2			13				4		127	1	6
United Kingdom		3	0	72	1	5	1	0	1	16		6	1		198	31
United States	0	36	1	15	17			1	100	27		1			31	1,959

Source: Authors' calculations.

Table A.3 Market Capitalisation of Firms from each Jurisdiction According to the Nationality of the Control Chain

	Market Capitalisation (USD bn)												
	Nationality of control chain (ultimate controller & main shareholder)												
	1	2	3		4 5 6 7 8 9 10 11 12 13								
	Total	Noncontrolled	Controlled	DD	DF	DT	FD	FN	FT	TD	TF	TT	FA
United States	14,453	12,195	2,257	1,955	1	1	132	90	13	2	5	16	42
China	**2,933**	**310**	**2,623**	**2,154**	**0**	**120**	**19**	**30**	**14**	**4**	**0**	**97**	**186**
Japan	2,573	1,661	912	668	6	0	10	60	8	0	1	17	142
United Kingdom	2,373	1,965	409	195	1	0	38	73	42	2	0	51	7
France	1,448	555	893	696	0	15	16	50	1	4	2	39	70
Canada	1,373	981	392	199	0	4	8	158	11	1	0	2	9
Hong Kong SAR China	1,249	233	1,015	266	5	41	188	121	331	0	0	49	14
Switzerland	1,182	753	430	248	0	1	35	9	6	1	1	6	122
Germany	1,152	667	486	413	0	3	18	18	10	13	0	3	7
Australia	1,067	840	228	106	0	0	23	49	2	10	2	27	10
India	**1,011**	**188**	**822**	**591**	**2**	**1**	**3**	**67**	**2**	**36**	**18**	**36**	**67**
South Korea	847	490	357	340	4	0	3	5	0	0	0	1	3
Russia	798	106	692	489	9	82	8	19	8	53	0	8	16
Brazil	**713**	**152**	**561**	**364**	**3**	**6**	**59**	**18**	**38**	**0**	**1**	**55**	**17**
Netherlands	640	267	373	263	0	8	15	76	5	3	0	0	2
Spain	487	257	230	156	0	0	23	47	4	0	0	0	0
Sweden	458	139	320	233	0	0	2	71	7	3	0	1	3
South Africa	**422**	**240**	**182**	**75**	**0**	**0**	**2**	**93**	**3**	**5**	**0**	**3**	**2**

Source: Authors' calculations.

Note: Rows in bold text indicate non-high-income countries according to the World Bank classification in 2012.

Table A.3 Market Capitalisation of Firms from each Jurisdiction According to the Nationality of the Control Chain (continued)

	Market Capitalisation (USD bn)												
	Nationality of control chain (ultimate controller & main shareholder)												
	1	2	3	4	5	6	7	8	9	10	11	12	13
	Total	Noncontrolled	Controlled	DD	DF	DT	FD	FN	FT	TD	TF	TT	FA
Taiwan	421	237	185	171	0	0	1	11	0	0	0	0	2
Italy	407	96	311	249	1	0	14	5	0	6	0	35	3
Singapore	383	157	225	145	19	0	2	18	1	0	0	21	20
Malaysia	**368**	**164**	**204**	**150**	**1**	**0**	**1**	**5**	**16**	**13**	**2**	**9**	**9**
Saudi Arabia	342	116	227	182	0	0	3	24	0	0	0	10	8
Indonesia	**310**	**34**	**276**	**104**	**1**	**5**	**30**	**18**	**4**	**16**	**0**	**79**	**16**
Belgium	305	32	273	53	0	127	4	60	9	5	0	14	2
Mexico	**272**	**92**	**180**	**119**	**0**	**0**	**0**	**22**	**2**	**0**	**0**	**6**	**30**
Chile	270	62	208	129	1	0	5	58	0	0	0	0	16
Turkey	**236**	**55**	**181**	**126**	**0**	**1**	**4**	**37**	**5**	**0**	**0**	**1**	**7**
Norway	225	76	148	143	0	0	0	4	0	0	0	0	0
Thailand	**205**	**43**	**162**	**110**	**0**	**0**	**0**	**17**	**11**	**0**	**0**	**23**	**1**
Colombia	**195**	**25**	**170**	**161**	**0**	**0**	**0**	**9**	**0**	**0**	**0**	**0**	**0**
Denmark	176	138	38	28	0	0	4	2	4	0	0	1	0
Bermuda	173	77	96	1	0	0	0	22	23	3	14	30	2
Ireland	145	116	29	11	0	0	0	9	0	4	2	0	3
Poland	144	18	126	72	0	5	1	37	3	0	2	6	0
Finland	126	73	53	43	1	0	0	3	0	0	0	2	3

Source: Authors' calculations.

Note: Rows in bold text indicate non-high-income countries according to the World Bank classification in 2012.

Table A.3 Market Capitalisation of Firms from each Jurisdiction According to the Nationality of the Control Chain (continued)

	Market Capitalisation (USD bn)												
	Nationality of control chain (ultimate controller & main shareholder)												
	1	2	3	4	5	6	7	8	9	10	11	12	13
	Total	Noncontrolled	Controlled	DD	DF	DT	FD	FN	FT	TD	TF	TT	FA
Philippines	**123**	**48**	**74**	**50**	**0**	**0**	**0**	**2**	**0**	**1**	**1**	**19**	**0**
Luxembourg	111	24	87	12	0	0	4	69	2	0	0	1	0
United Arab Emirates	110	19	92	90	0	1	0	1	0	0	0	0	0
Israel	109	41	69	60	0	0	0	3	2	0	0	2	1
Cayman Islands	101	32	69	0	0	0	3	16	14	0	0	22	14
Austria	93	13	80	65	0	0	6	1	4	0	0	1	2
Peru	**89**	**21**	**69**	**33**	**0**	**1**	**4**	**21**	**7**	**0**	**0**	**0**	**2**
Qatar	87	9	78	75	0	0	0	0	0	0	0	2	2
Morocco	**56**	**20**	**36**	**13**	**0**	**0**	**5**	**14**	**0**	**0**	**3**	**1**	**0**
Kuwait	55	23	32	22	0	2	0	4	0	0	0	4	0
Portugal	51	8	43	25	0	2	0	3	2	0	0	0	11
Egypt	**41**	**10**	**30**	**19**	**0**	**0**	**3**	**5**	**0**	**0**	**1**	**2**	**0**
New Zealand	39	30	10	2	1	0	1	5	0	0	0	1	0
Argentina	**32**	**6**	**26**	**13**	**0**	**0**	**2**	**6**	**2**	**0**	**0**	**0**	**3**
Nigeria	**30**	**6**	**25**	**12**	**0**	**0**	**0**	**4**	**0**	**0**	**1**	**7**	**0**
Greece	29	10	19	12	1	3	0	3	0	0	0	0	0
Ukraine	**23**	**2**	**21**	**5**	**0**	**3**	**0**	**6**	**0**	**0**	**0**	**4**	**2**
Jordan	**22**	**15**	**8**	**2**	**0**	**0**	**2**	**3**	**0**	**0**	**0**	**0**	**0**

Source: Authors' calculations.

Note: Rows in bold text indicate non-high-income countries according to the World Bank classification in 2012.

Table A.3 Market Capitalisation of Firms from each Jurisdiction According to the Nationality of the Control Chain (continued)

	Market Capitalisation (USD bn)												
	Nationality of control chain (ultimate controller & main shareholder)												
	1	2	3	4	5	6	7	8	9	10	11	12	13
	Total	Noncontrolled	Controlled	DD	DF	DT	FD	FN	FT	TD	TF	TT	FA
Czechia	22	0	22	0	0	0	0	21	0	0	0	0	1
Vietnam	**21**	**12**	**9**	**8**	**0**	**0**	**0**	**1**	**0**	**0**	**0**	**0**	**0**
Croatia	20	2	18	3	0	0	7	6	2	0	0	0	0
Hungary	**19**	**8**	**11**	**8**	**0**	**0**	**0**	**3**	**0**	**0**	**1**	**0**	**0**
Pakistan	**18**	**6**	**12**	**4**	**0**	**0**	**0**	**3**	**0**	**0**	**1**	**3**	**2**
Oman	17	8	9	7	0	0	0	1	0	0	0	1	0
Bahrain	15	4	11	4	0	0	0	4	0	0	0	1	3
British Virgin Islands	14	5	9	0	0	0	0	0	0	0	8	0	0
Romania	**14**	**3**	**11**	**1**	**0**	**0**	**0**	**8**	**1**	**0**	**0**	**0**	**0**
Bangladesh	**12**	**3**	**9**	**1**	**0**	**0**	**0**	**4**	**1**	**0**	**0**	**0**	**3**
Venezuela	**11**	**2**	**8**	**7**	**0**	**0**	**0**	**1**	**0**	**0**	**0**	**0**	**0**
Kenya	**8**	**0**	**7**	**0**	**0**	**0**	**0**	**6**	**0**	**0**	**1**	**0**	**0**
Sri Lanka	**7**	**3**	**4**	**2**	**0**	**0**	**0**	**2**	**0**	**0**	**0**	**0**	**0**
Tunisia	**7**	**1**	**5**	**2**	**0**	**0**	**0**	**2**	**0**	**0**	**0**	**0**	**0**
Slovenia	6	1	5	5	0	0	0	0	0	0	0	0	0
Iceland	6	1	4	2	0	0	2	0	0	0	0	0	0
Cyprus	5	2	4	0	0	0	3	0	0	0	0	0	0
Lithuania	4	1	4	2	0	0	0	2	0	0	0	0	0

Source: Authors' calculations.

Note: Rows in bold text indicate non-high-income countries according to the World Bank classification in 2012.

Table A.3 Market Capitalisation of Firms from each Jurisdiction According to the Nationality of the Control Chain (continued)

	Market Capitalisation (USD bn)												
	Nationality of control chain (ultimate controller & main shareholder)												
	1	2	3	4 5 6 7 8 9 10 11 12 13									
	Total	Noncontrolled	Controlled	DD	DF	DT	FD	FN	FT	TD	TF	TT	FA
Bulgaria	**4**	**0**	**4**	**1**	**0**	**0**	**0**	**1**	**1**	**0**	**0**	**1**	**0**
Marshall Islands	**4**	**3**	**1**	**0**	**0**	**0**	**0**	**1**	**0**	**0**	**0**	**0**	**0**
Malta	4	0	4	1	0	0	0	2	0	0	0	0	0
Slovakia	4	1	3	1	0	0	0	2	0	0	0	0	0
Serbia	**3**	**0**	**3**	**0**	**0**	**0**	**0**	**1**	**0**	**0**	**0**	**0**	**0**
Bosnia & Herzegovina	**3**	**0**	**2**	**1**	**0**	**0**	**0**	**1**	**0**	**0**	**0**	**0**	**0**
Mauritius	**2**	**1**	**1**	**1**	**0**	**0**	**0**	**0**	**0**	**0**	**0**	**0**	**0**
Cote d'Ivoire	**2**	**0**	**2**	**0**	**0**	**0**	**0**	**1**	**0**	**0**	**0**	**0**	**0**
Zambia	**2**	**0**	**2**	**0**	**0**	**0**	**0**	**1**	**0**	**0**	**0**	**0**	**0**
Montenegro	**2**	**0**	**2**	**0**	**0**	**0**	**0**	**1**	**0**	**0**	**0**	**0**	**0**
Estonia	2	0	2	1	0	0	0	0	0	0	0	0	0
Ghana	**1**	**0**	**1**	**0**	**0**	**0**	**0**	**0**	**0**	**0**	**0**	**0**	**0**
Latvia	1	0	1	0	0	0	0	0	0	0	0	0	0
Zimbabwe	**0**	**0**	**0**	**0**	**0**	**0**	**0**	**0**	**0**	**0**	**0**	**0**	**0**
Total	**41,345**	**24,016**	**17,329**	**11,978**	**57**	**434**	**716**	**1,657**	**623**	**184**	**69**	**723**	**889**

Source: Authors' calculations.

Note: Rows in bold text indicate non-high-income countries according to the World Bank classification in 2012.

Table A.4 Number of Firms from each Jurisdiction According to the Nationality of the Control Chain

	Number of companies												
	Nationality of control chain (ultimate controller & main shareholder)												
	1	2	3	4 5 6 7 8 9 10 11 12 13									
	Total	Noncontrolled	Controlled	DD	DF	DT	FD	FN	FT	TD	TF	TT	FA
United States	4,246	3,095	1,151	795	4	6	17	71	7	2	3	15	231
Canada	1,862	1,388	474	208	5	7	16	105	17	1	0	16	99
China	**1,659**	**450**	**1,209**	**797**	**0**	**57**	**8**	**19**	**6**	**4**	**0**	**23**	**294**
Japan	1,446	760	686	564	5	0	13	27	6	1	1	5	64
India	**1,317**	**669**	**648**	**332**	**6**	**4**	**11**	**62**	**11**	**10**	**4**	**15**	**192**
Australia	1,314	1,009	305	144	2	3	28	51	12	3	1	16	45
United Kingdom	1,252	995	257	108	2	3	17	47	11	1	0	11	57
Taiwan	940	796	144	120	0	0	3	10	0	0	0	3	8
South Korea	798	507	291	250	1	0	4	17	0	0	0	4	15
France	772	247	525	333	3	10	14	16	8	7	2	28	104
Germany	681	208	473	300	2	7	33	47	20	18	1	20	25
Hong Kong SAR China	671	270	401	65	2	61	22	46	51	2	1	66	85
Poland	622	230	392	262	4	44	7	35	11	1	1	19	8
Malaysia	**516**	**227**	**289**	**220**	**4**	**2**	**3**	**11**	**5**	**7**	**1**	**13**	**23**
Singapore	464	218	246	89	3	2	9	43	4	0	0	14	82
Israel	435	160	275	252	0	2	2	4	2	1	0	3	9
Russia	432	90	342	262	5	35	2	6	4	8	0	7	13
Sweden	325	184	141	94	1	1	5	22	1	2	0	2	13

Source: Authors' calculations.

Note: Rows in bold text indicate non-high-income countries according to the World Bank classification in 2012.

Table A.4 Number of Firms from each Jurisdiction According to the Nationality of the Control Chain (continued)

	Number of companies												
	Nationality of control chain (ultimate controller & main shareholder)												
	1	2	3	4 5 6 7 8 9 10 11 12 13									
	Total	Noncontrolled	Controlled	DD	DF	DT	FD	FN	FT	TD	TF	TT	FA
Turkey	**294**	**69**	**225**	**153**	**0**	**2**	**8**	**19**	**4**	**1**	**1**	**3**	**34**
Brazil	**273**	**78**	**195**	**118**	**2**	**1**	**24**	**15**	**5**	**0**	**2**	**4**	**24**
Switzerland	269	132	137	87	1	3	8	19	4	1	1	5	8
Italy	260	79	181	148	4	6	4	4	1	4	0	5	5
Indonesia	**244**	**50**	**194**	**103**	**2**	**5**	**7**	**22**	**5**	**7**	**0**	**14**	**29**
Greece	224	52	172	138	4	7	2	6	2	2	1	1	9
South Africa	**202**	**123**	**79**	**41**	**0**	**0**	**5**	**16**	**2**	**1**	**0**	**3**	**11**
Norway	197	117	80	60	0	1	2	9	1	0	0	4	3
Chile	179	59	120	72	1	0	12	13	0	1	0	1	20
Spain	178	89	89	66	0	1	3	13	2	0	0	1	3
Croatia	167	65	102	64	1	1	13	15	4	0	0	1	3
Bermuda	158	73	85	2	0	1	2	15	11	2	2	35	15
Belgium	156	58	98	49	0	7	6	13	2	3	0	6	11
Kuwait	153	72	81	67	0	7	0	3	0	1	1	1	1
Denmark	151	98	53	33	0	1	6	5	1	0	1	2	4
Romania	**144**	**33**	**111**	**74**	**1**	**3**	**7**	**10**	**8**	**3**	**0**	**3**	**2**
Netherlands	133	71	62	31	0	2	4	13	2	3	1	3	3
Cayman Islands	127	57	70	1	0	0	1	6	16	0	0	22	24

Source: Authors' calculations.

Note: Rows in bold text indicate non-high-income countries according to the World Bank classification in 2012.

Table A.4 Number of Firms from each Jurisdiction According to the Nationality of the Control Chain (continued)

	Number of companies												
	Nationality of control chain (ultimate controller & main shareholder)												
	1	2	3	4 5 6 7 8 9 10 11 12 13									
	Total	Noncontrolled	Controlled	DD	DF	DT	FD	FN	FT	TD	TF	TT	FA
Thailand	**126**	**45**	**81**	**40**	**0**	**0**	**1**	**24**	**4**	**0**	**0**	**6**	**6**
Peru	**119**	**30**	**89**	**44**	**2**	**6**	**6**	**14**	**6**	**1**	**0**	**2**	**8**
Saudi Arabia	119	61	58	42	0	0	3	6	0	0	0	6	1
Jordan	**116**	**59**	**57**	**32**	**1**	**1**	**2**	**10**	**3**	**2**	**1**	**2**	**3**
Finland	112	71	41	28	1	0	1	8	0	0	0	1	2
New Zealand	97	69	28	10	2	0	2	6	0	2	0	4	2
Ukraine	**97**	**25**	**72**	**20**	**0**	**17**	**3**	**6**	**5**	**0**	**0**	**7**	**14**
Austria	95	17	78	48	0	3	13	5	2	0	0	3	4
United Arab Emirates	94	30	64	57	0	1	0	4	0	0	0	1	1
Pakistan	**89**	**29**	**60**	**22**	**2**	**1**	**0**	**19**	**1**	**0**	**2**	**9**	**4**
Egypt	**85**	**31**	**54**	**31**	**0**	**0**	**8**	**8**	**0**	**1**	**1**	**3**	**2**
Serbia	**77**	**33**	**44**	**22**	**0**	**1**	**4**	**5**	**4**	**0**	**0**	**1**	**7**
Argentina	**76**	**5**	**71**	**25**	**0**	**0**	**10**	**13**	**1**	**1**	**2**	**0**	**19**
Bosnia & Herzegovina	**74**	**36**	**38**	**20**	**0**	**0**	**2**	**13**	**0**	**0**	**0**	**2**	**1**
Sri Lanka	**73**	**18**	**55**	**37**	**0**	**0**	**2**	**8**	**2**	**0**	**0**	**1**	**5**
Montenegro	**72**	**25**	**47**	**26**	**0**	**0**	**5**	**10**	**1**	**0**	**0**	**0**	**5**
Ireland	65	48	17	5	0	0	1	5	1	2	1	0	2
Bulgaria	**60**	**9**	**51**	**32**	**1**	**0**	**4**	**3**	**3**	**2**	**0**	**2**	**4**

Source: Authors' calculations.

Note: Rows in bold text indicate non-high-income countries according to the World Bank classification in 2012.

Table A.4 Number of Firms from each Jurisdiction According to the Nationality of the Control Chain (continued)

	Number of companies												
	Nationality of control chain (ultimate controller & main shareholder)												
	1	2	3	4 5 6 7 8 9 10 11 12 13									
	Total	Noncontrolled	Controlled	DD	DF	DT	FD	FN	FT	TD	TF	TT	FA
Cyprus	57	33	24	15	0	0	4	2	0	0	1	0	2
Morocco	**57**	**13**	**44**	**24**	**0**	**0**	**8**	**5**	**0**	**0**	**1**	**1**	**5**
Philippines	**57**	**22**	**35**	**17**	**0**	**2**	**0**	**6**	**0**	**2**	**1**	**5**	**2**
Portugal	51	15	36	26	0	2	0	2	2	0	0	0	4
Mexico	**50**	**9**	**41**	**28**	**0**	**0**	**1**	**6**	**1**	**0**	**0**	**1**	**4**
Slovenia	46	16	30	26	0	1	0	0	0	0	0	0	3
Vietnam	**46**	**34**	**12**	**10**	**0**	**0**	**0**	**2**	**0**	**0**	**0**	**0**	**0**
Bangladesh	**45**	**17**	**28**	**9**	**0**	**0**	**0**	**7**	**1**	**0**	**0**	**0**	**11**
Luxembourg	42	16	26	4	0	0	4	14	3	0	0	1	0
Oman	41	14	27	14	0	0	3	6	1	0	0	2	1
Bahrain	40	19	21	9	1	0	0	9	0	0	0	1	1
Hungary	**37**	**15**	**22**	**11**	**0**	**1**	**0**	**4**	**1**	**0**	**1**	**1**	**3**
Nigeria	**37**	**12**	**25**	**4**	**1**	**0**	**0**	**11**	**1**	**0**	**2**	**4**	**2**
Lithuania	34	8	26	19	1	0	1	5	0	0	0	0	0
Colombia	**33**	**11**	**22**	**16**	**0**	**0**	**0**	**4**	**0**	**0**	**0**	**1**	**1**
Tunisia	**31**	**7**	**24**	**12**	**1**	**0**	**3**	**6**	**0**	**0**	**0**	**2**	**0**
Slovakia	28	7	21	12	0	0	0	5	2	0	0	0	2
British Virgin Islands	27	19	8	1	0	0	0	2	1	0	1	2	1

Source: Authors' calculations.

Note: Rows in bold text indicate non-high-income countries according to the World Bank classification in 2012.

Table A.4 Number of Firms from each Jurisdiction According to the Nationality of the Control Chain (continued)

	Number of companies												
	Nationality of control chain (ultimate controller & main shareholder)												
	1	2	3		4 5 6 7 8 9 10 11 12 13								
	Total	Noncontrolled	Controlled	DD	DF	DT	FD	FN	FT	TD	TF	TT	FA
Qatar	27	9	18	15	0	0	0	0	0	0	0	1	2
Latvia	24	6	18	13	0	0	1	2	1	0	0	1	0
Czechia	19	1	18	0	0	0	0	13	2	0	0	1	2
Iceland	17	8	9	8	0	0	1	0	0	0	0	0	0
Kenya	**16**	**3**	**13**	**0**	**0**	**0**	**0**	**10**	**0**	**0**	**1**	**0**	**1**
Malta	16	1	15	5	0	0	2	2	1	1	0	2	2
Estonia	15	4	11	10	0	0	0	0	1	0	0	0	0
Cote d'Ivoire	**14**	**1**	**13**	**0**	**0**	**0**	**0**	**9**	**0**	**0**	**2**	**1**	**1**
Mauritius	**14**	**5**	**9**	**4**	**0**	**0**	**0**	**2**	**1**	**0**	**0**	**0**	**2**
Ghana	**12**	**1**	**11**	**0**	**0**	**0**	**0**	**5**	**1**	**0**	**0**	**1**	**4**
Venezuela	**12**	**3**	**9**	**7**	**0**	**0**	**0**	**1**	**0**	**0**	**0**	**0**	**1**
Marshall Islands	**11**	**9**	**2**	**0**	**0**	**0**	**0**	**1**	**0**	**0**	**0**	**0**	**1**
Zimbabwe	**11**	**5**	**6**	**0**	**0**	**0**	**0**	**4**	**0**	**0**	**1**	**0**	**1**
Zambia	**10**	**2**	**8**	**0**	**0**	**0**	**0**	**6**	**0**	**0**	**1**	**1**	**0**
Total	**25,884**	**13,864**	**12,020**	**7,362**	**78**	**328**	**423**	**1,174**	**300**	**111**	**44**	**473**	**1,725**

Source: Authors' calculations.

Note: Rows in bold text indicate non-high-income countries according to the World Bank classification in 2012.

The Reversal Problem: Development Going Backwards

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The Covid-19 pandemic triggered the most synchronous economic downturn in more than a century. Ninety percent of countries posted a decline in real per-capita GDP in 2020, a share that surpassed any other year since 1900, which includes two world wars and the Great Depression of the 1930s.[1] The health crisis pushed an estimated 90 million people into extreme poverty.[2] We document that, for emerging and developing economies (EMDEs) as a group, the setback in their development markers did not start with the pandemic. Covid-19 deepened and accelerated a troubling trend of economic backsliding that had appeared around half a decade earlier. We call this the Reversal Problem.

In the 21st century, several benign trends emerged in EMDEs: progress on poverty reduction, a narrowing of the gap between rich and poor nations, higher growth rates in per-capita incomes, and a much lower incidence of economic crises. Gender equality gained a stronger footing, and many social indicators improved. Yet a turning point around 2015, marked by the largest decline in commodity prices since the early 1980s, began to chip away at these gains in development. The setback in economic and social conditions was already broadly synchronized when Covid-19 hit the global economy.

1. Holston and Reinhart (2022).

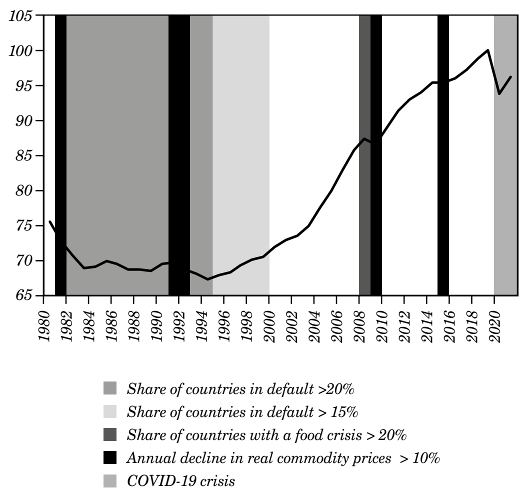
2. Mahler,Daniel Gerszon & Yonzan,Nishant & Lakner,Christoph, 2022. "The Impact of Covid-19 on Global Inequality and Poverty," Policy Research Working Paper Series 10198, The World Bank.

Credibility of Emerging Markets, Foreign Investors' Risk Perceptions, and Capital Flows edited by Álvaro Aguirre, Andrés Fernández, and Şebnem Kalemli-Özcan, Santiago, Chile.

The Reversal Problem is widespread in terms of geography, though more acute in Africa and other low-income regions. It is also encompassing in that it not only affects economic and social indicators but is potentially politically destabilizing internally and across national borders, as these trends appear to coincide with a setback in democratic values.

We take a step back and review the bonanza years, which roughly span 2000–2015 and followed the calamitous 1980s (and somewhat better 1990s), which marked an earlier *reversal in development* in scores of low- and middle-income countries (figure 1, table 1). We document these trends with an encompassing array of indicators. The commonality in these economic, social, and political indicators is that they follow a distinctive U-shaped pattern, from worse to better and back to worse, indicating that, across many dimensions, EMDEs as a group have been moving from progress to stagnation to reversal.

Figure 1. The Big Picture: Real Per-capita GDP EMDEs Uunweighted Average, 1980–2021 Index 2019 = 100



Sources: World Economic Outlook, IMF; Farah and others (2022).

Table 1. Real per-capita GDP

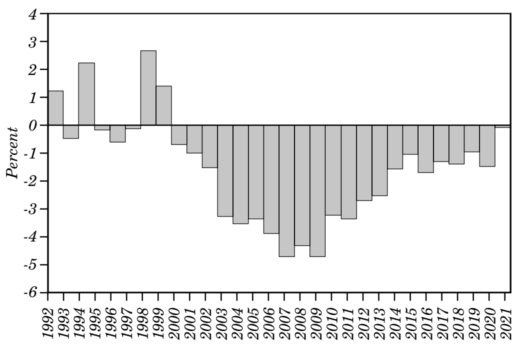
Period	*Change*	*Period*	*Change*
1980–1994	–0.8	2010–2014	2.0
1995–2000	1.2	2015–2019	1.0
2001–2007	2.8	2020–2021	–1.9
2008–2009	0.5		

Source: Authors' calculations.

1. The Bonanza Years

The EMDE debt overhang of the 1980s was ultimately addressed in the early 1990s for middle-income countries by the Brady Plan and, some years later, for the low-income ones through the Heavily Indebted Poor Countries (HIPC) debt relief initiative. The resolution of protracted debt crises paved the way for recovery. For EMDEs, the first decade of the 21st century was one of high growth by historic standards, as growth in per-capita GDP outstripped that of advanced economies (AEs) by a widening margin (figure 2). The share of EMDEs with negative 5-year real GDP growth declined from 40 percent in the early 1990s to less than 10 percent in 2008. Social conditions improved significantly as compared to the previous two decades, with considerable declines in poverty rates and a more equitable income distribution.

Figure 2. Real Per-capita GDP Growth Differential: AEs Less EMDEs, 1992–2021
(in percent)



Source: Authors' calculations and World Economic Outlook, IMF.

Robust economic growth allowed EMDEs to improve their fiscal and external positions and strengthen their financial-sector policy frameworks. The incidence of high inflation and debt and financial crises, which dealt devastating setbacks to growth in many EMDEs from the 1970s to 1990s, saw a significant reduction.[3] Subdued inflation allowed central banks to maintain comparatively low policy rates, and rising international reserves strengthened external buffers. EMDE current-account deficits narrowed, on average, from 3.5 percent of GDP in 2001 to 1.2 percent of GDP in 2007. About 70 percent of EMDEs increased their international reserves-to-external-debt ratio by more than 10 percentage points, while one-quarter posted increases of more than 50 percentage points. Many EMDEs faced the 2009 global recession with greater resilience than in the past.

Sustained growth helped reduce global poverty and inequality. The number of countries classified as 'low-income' declined from 66 in 2001 to 31 in 2015. Extreme poverty fell from 30 percent of the global population in 2000 to 10 percent in 2015. Between-country inequality improved, halving the global Gini index between 2000 and 2015. In most EMDEs, within-country inequality also declined. In Latin America, for instance, from 2000 to 2015, more than 50 million people were lifted out of poverty and the middle class swelled to more than a third of the population. Educational gaps were reduced. For example, between 2000 and 2017, in India and Nigeria, rates of secondary attainment increased from 10.9 to 37.2 percent and from 11.5 to 45 percent, respectively.

In sum, the first 15 years of this century is the only period since at least 1950 that has delivered a sustained EMDE outperformance. However, the new prosperity possibly owed less to a lasting shift in EMDE's fundamentals than to sustained favorable external conditions. An important factor was China's rapid investment-led growth, which lifted global commodity prices and fueled the longest commodity boom since the early 19th century. Another engine of growth was the surge in world trade, which as a share of world GDP, rose from less than 40 percent in 1990 to over 60 percent in 2011. Importantly, globalized finance and low interest rates in financial centers incentivized risk-taking; investors sought opportunities in emerging markets. Gross capital inflows to EMDEs excluding foreign direct investment swelled nearly seven-fold (from 1 percent of GDP in 2001 to 6.5 percent in 2007).

3. See Reinhart and Rogoff (2009).

2. The Onset of the Reversal Problem

During 2015–2019, real per-capita GDP growth in EMDEs slowed markedly, averaging one percent or one half the growth rate of the previous five years, and about a third of the 2000–2007 rate (table 1). The economic slowdown was followed by the Covid-19 outbreak. While 2020 was an abysmal year for most countries, preliminary data indicate that 2021 delivered a regressive and uneven rebound. As table 2 highlights, 2021 marked a new peak in per-capita income in more than a third of AEs. By contrast, for middle- and low-income EMDEs, the share of countries at peak per-capita GDP in 2021 is much lower, about 21 percent and 12 percent, respectively. Put differently, between 80 and 90 percent of EMDEs were below their prior per-capita income levels, let alone the levels that would have prevailed had output followed its pre-crisis trend. Mounting evidence suggests that the crisis may have lasting effects on growth through its impact on human capital and income inequality.

The scale and breadth of the economic setback are evident in other macroeconomic variables. The re-emergence of debt problems, particularly in low-income countries, figures prominently in the Reversal Problem. Levels of public indebtedness also follow a distinctive U-shaped pattern (figure 3).

EMDEs managed to weather the 2008–2009 crisis thanks to fiscal buffers and sustained deleveraging during the years leading up to that crisis, but since then, debt and budget deficits have risen steadily, thus denting resilience. The fiscal health of commodity-exporting EMDEs deteriorated sharply since the end of the commodity super-cycle in 2013–2014. The setback in public finances has been mirrored in a reverse-U in the sovereign credit ratings of EMDEs, as the steady improvement in ratings during the high-growth era was replaced by recurrent downgrades (figure 4). While public debt loads were mounting in most EMDEs well before 2019, the pandemic-induced crash in economic activity and government revenues pushed several 'graduates' of the Brady and HIPC plans back into debt distress territory.

The reversal in inflation is of a more recent vintage and common to AEs. In a manner reminiscent of 2008, although on a lesser scale (to date), food price inflation, with its regressive impacts, has also reemerged. The rekindling of inflationary pressures, even prior to the Russia-Ukraine war, was proving to be more persistent than anticipated by many central banks and will represent a challenge to the relatively new-found central-bank independence in EMDEs.

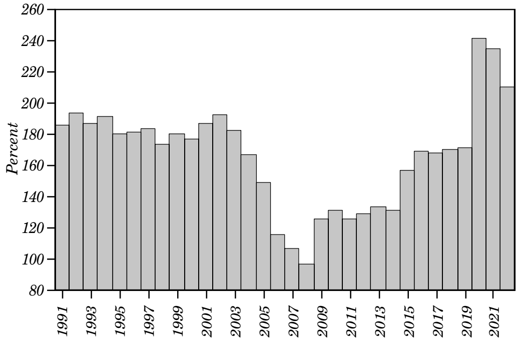
Table 2. Per-capita GDP, Peak Versus 2021: 1980–2021, 194 Countries

	Advanced economies		*Middle-income*		*Low-income*	
	Number of countries	Share (%)	Number of countries	Share (%)	Number of countries	Share (%)
2021 = peak	13	35.1	27	20.6	3	11.5
2021 < peak	24	64.9	104	79.4	23	88.5
Total	**37**	**100**	**131**	**100**	**26**	**100**

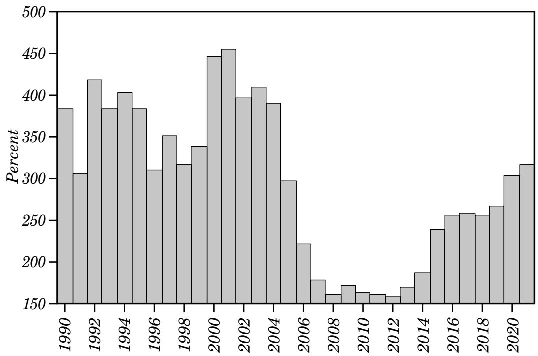
Sources: World Economic Outlook, IMF; and Holston and Reinhart (2022).

Figure 3. General Government Debt-to-Revenue Ratio: EMDEs, 1990–2021
(in percent)

80 low-and middle-low-income countries

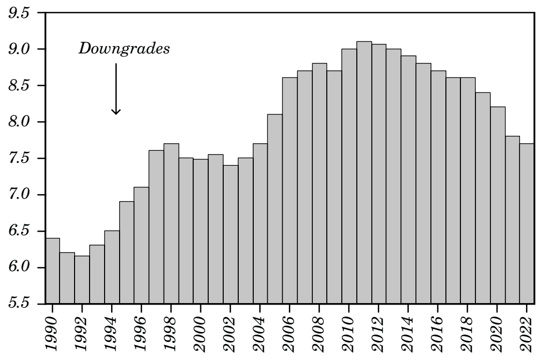


75 middle-upper and high-income EMDEs



Source: Authors' calculations and World Economic Outlook, IMF.

Figure 4. The Reverse U: Fitch, Institutional Investors, Moody's, and S&P Sovereign Ratings: Average Numeric Score: EMDEs, 1990.1–2022.3



Sources: Authors' calculations based on data from Fitch, Institutional Investors, Moody's, and S&P Sovereign Ratings.

3. The Reversal in Shared Prosperity

Following two decades of progress in terms of shared prosperity, reductions in poverty reduction have also suffered a marked reversal. The world made extraordinary progress in reducing poverty at the turn of the century. However, the sustained decline in extreme poverty that began in the 1990s was already stalling in the pre-Covid-19 decade. As the health crisis persisted, stalling morphed into outright reversal. According to the World Bank, from 1990 to 2015, extreme global poverty dropped by an average of about 1 percentage point per year, but this pace slowed from 2013 to 2015 to 0.6 percentage point per year before slowing further in the years before Covid-19.

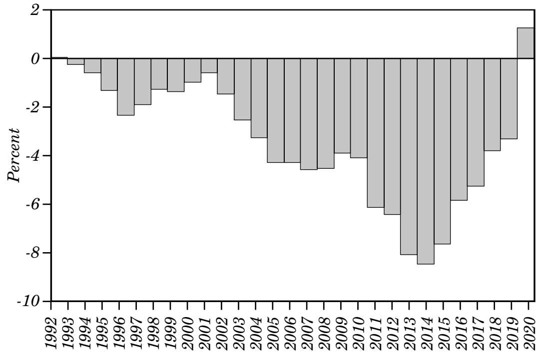
The slowdown in global poverty reduction is partly accounted for by the increasing concentration of extreme poverty in two regions: Sub-Saharan Africa, which has experienced a slower reduction in poverty than any other region, and poverty spikes in the Middle East and North Africa, where extreme-poverty rates nearly doubled between 2015 and 2018, from 3.8 percent to 7.2 percent, spurred by the ongoing conflicts.[4]

The poverty reversal can be gleaned from the share of countries that have higher levels of extreme poverty than the average during the prior 5 years (figure 5). From 1995 to 2008, the share of worsening cases dropped from 42 percent of EMDEs to less than 10, but progress

4. World Bank (2022).

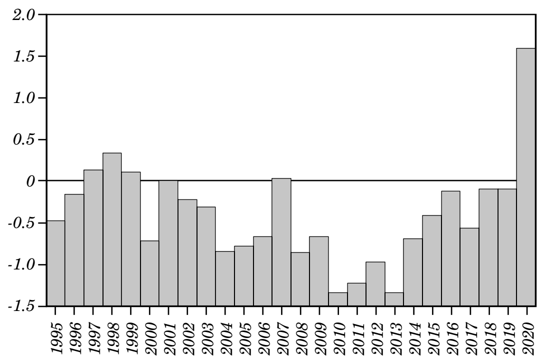
stalled, and the share climbed back to 18 percent in 2019 before surging to almost 42 percent in 2020 amid the pandemic. After the initial Covid-19 impact, extreme poverty has continued to increase. World Bank estimates suggest that the Covid-19-induced extreme poverty is set to rise by 0.9 percentage points and even more in Sub-Saharan Africa.[5]

Figure 5. Annual Change in the Global Poverty Rate, 1992–2020
(moving 5-year average, in percent)



Sources: Authors' calculations based on data from Mahler and others (2022).
Notes: This chart shows the 5-year rolling average of the annual percent change in the global poverty rate ($1.90 poverty line), using the global distributions in the Poverty and Inequality Portal (PIP) for the historical series and the simulations by Mahler and others (2022).

Figure 6. Income Inequality: Global Gini, 1995–2020
Rolling 5-year Average of Annual Change



Source: Authors' calculations and Mahler and others (2022).
Note: This chart shows the annual percentage change in the global Gini index (5-year moving average), using the global distributions in PIP for the historical series and the simulations conducted by Mahler and others (2022).

5. Cojocaru and others (2022).

The pandemic has been particularly harsh on the urban poor and, notably, women. In many cases, the Covid-19 rise in inequality reinforced trends already in place and clawed back earlier progress. Faster economic growth during 2000–2010 facilitated a decline in within-country income inequality globally, albeit from historically high levels in the 1990s and early 2000s. Income inequality declined in more than 60 percent of EMDEs, as measured by the Gini Index. As with poverty reduction, improvement had stalled in the years prior to the pandemic, when the reversal materialized in earnest. In stark contrast to the Global Financial Crisis of 2008–10, the lagging economic recovery in EMDEs compared with AEs has raised between-country income inequality (table 2 and figure 6). According to the Gini and Theil indices, between-country inequality reverted to the levels of the early 2010s. Within-country income inequality in EMDEs has also increased in 2020 (by about 0.3 points), reversing the steady decline since the 2000s.[6]

4. The human capital reversal

The rise in income inequality is already significant and it is poised to climb further over the medium term. The accumulation of human capital has been severely disrupted in most EMDEs, but the pandemic has had a disproportionate impact on children from low-income households. The long-term consequences of students missing many days of school and some, notably girls, not returning to the classroom cannot be underestimated as a fault line in the development process.

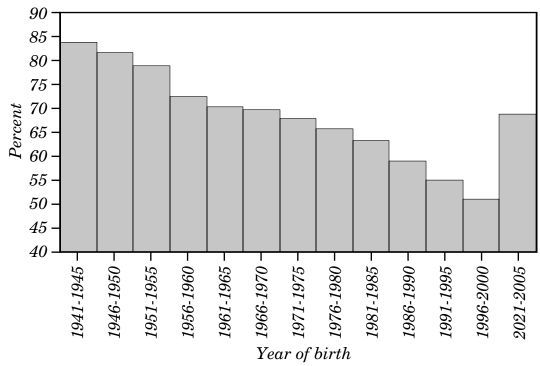
School closures affected 1.6 billion children across the globe,[7] but lower- and middle-income countries have closed schools for far longer than their higher-income counterparts. In parts of South Asia, Latin America, and Africa, schools have been closed for over 80 weeks. Uganda, which reopened schools in January 2022, topped the charts with 82 weeks of partial or full closure. Recent research suggests that, in Latin America, the likelihood that today's students will complete secondary education has dropped from a regional average of 61 percent to 46 percent (figure 7), but averages mask the striking differences across socioeconomic groups. Although school closures affected all students, their ability to continue learning depended on their parents' income and educational level. Children in disadvantaged households found it difficult, if not impossible, to continue their education at

6. Adarov (2022).
7. World Bank (2021a).

home due to a lack of adequate equipment, connectivity, and one-on-one coaching by well-educated parents.[8] For these children, the odds of completing secondary school may be closer to 32 percent, a level of educational attainment that was last reported for cohorts born in the 1960s.

Covid-19 has also set back the provision of health services in EMDEs. Evidence compiled by the World Health Organization and the World Bank shows that the Covid-19 pandemic is likely to halt two decades of global progress toward Universal Health Coverage.[9] Many EMDEs had made progress on service coverage and, by 2019, 68 percent of the world's population was covered by essential health services, such as pre- and post-natal care, immunization services, treatment for diseases like HIV, TB, and malaria, and services to diagnose and treat noncommunicable diseases like cancer and diabetes. The pandemic disrupted regular health services and stretched countries' health systems. As a result, for instance, immunization coverage dropped for the first time in ten years, and the incidence of deaths from TB and malaria has increased.

Figure 7. Likelihood that Disadvantaged Children do not Complete Secondary Education: Latin America, 1941–2005



Sources: Neidhöfer and others (2021) based on data from Latinobarometro.
Notes: This chart shows the likelihood that children, whose parents have less than secondary education, will complete secondary education. Unweighted average.

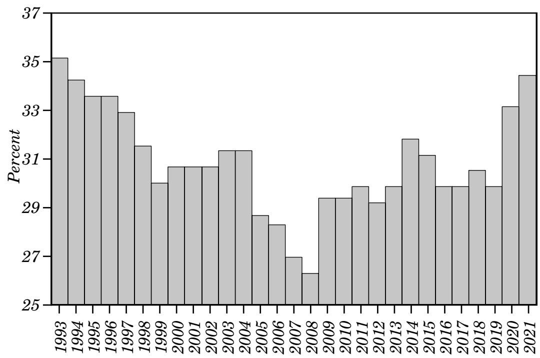
8. World Bank (2021b).
9. World Bank (2021c).

5. Political reversals

Slower growth, increasing poverty and inequality, and dissatisfaction with the quality of public services, particularly evident in the (mis) management of the pandemic in many EMDEs, are contributing to dissatisfaction with governments and democratic values. According to data collected by *Freedom House*, the share of countries that are 'not free' has followed a U-shaped pattern that mimics what we observe in other economic and social indicators.[10] Societies around the world are experiencing radical political fractures that were deemed long gone. These manifestations preceded the pandemic but have spiked in the past two years.

Violent conflicts have increased[11] to the highest levels observed over the past three decades, thus affecting both low- and middle-income countries, and there were some 80 million forcibly displaced people worldwide as of the end of 2019, the highest number recorded. Russia's invasion of Ukraine is adding to those numbers daily (Figure 9).

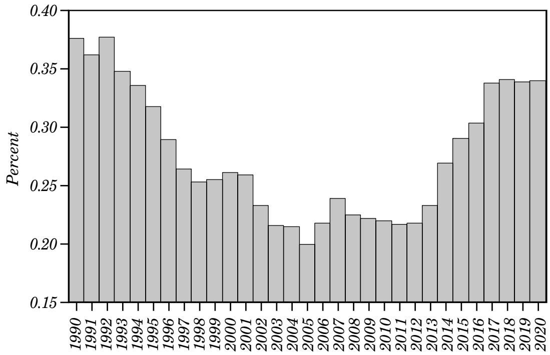
Figure 8. Share of Countries that are *Not Free*, 1993–2021



Source: Freedom House.
Note: This chart shows the percentage of 191 countries that have been classified as Not Free by Freedom in the World. For each country and territory, Freedom in the World analyzes the electoral process, political pluralism and participation, the functioning of the government, freedom of expression and belief, associational and organizational rights, the rule of law, and personal autonomy and individual rights.

10. See notes to figure 8.
11. World Bank (2020).

Figure 9. Refugees as a Percentage of World Population, 1990–2021



Source: World Bank World Development Indicators (2022).

6. Final Thoughts

The Covid-19 crisis has set back many of the development markers where significant progress had been achieved, especially so in the many countries where the economic and social fundamentals had already started to backslide prior to the pandemic. While the setbacks are more acute in low-income countries, the trend is far more encompassing among EMDEs. If the past is any guide, policy reversals often follow closely on the heels of 'bad times'. Policy reversals, in turn, set the stage for self-reinforcing economic and social setbacks. The international community and the AEs must come to terms with the heightened risks of delaying coordinated assistance replaying the calamitous 1980s. While the extent of the development Reversal Problem described here is not *yet* on the scale development reversals of 40 years ago, there are added new risks posed by climate change that will disproportionally harm EMDEs.

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