DOCUMENTOS DE TRABAJO

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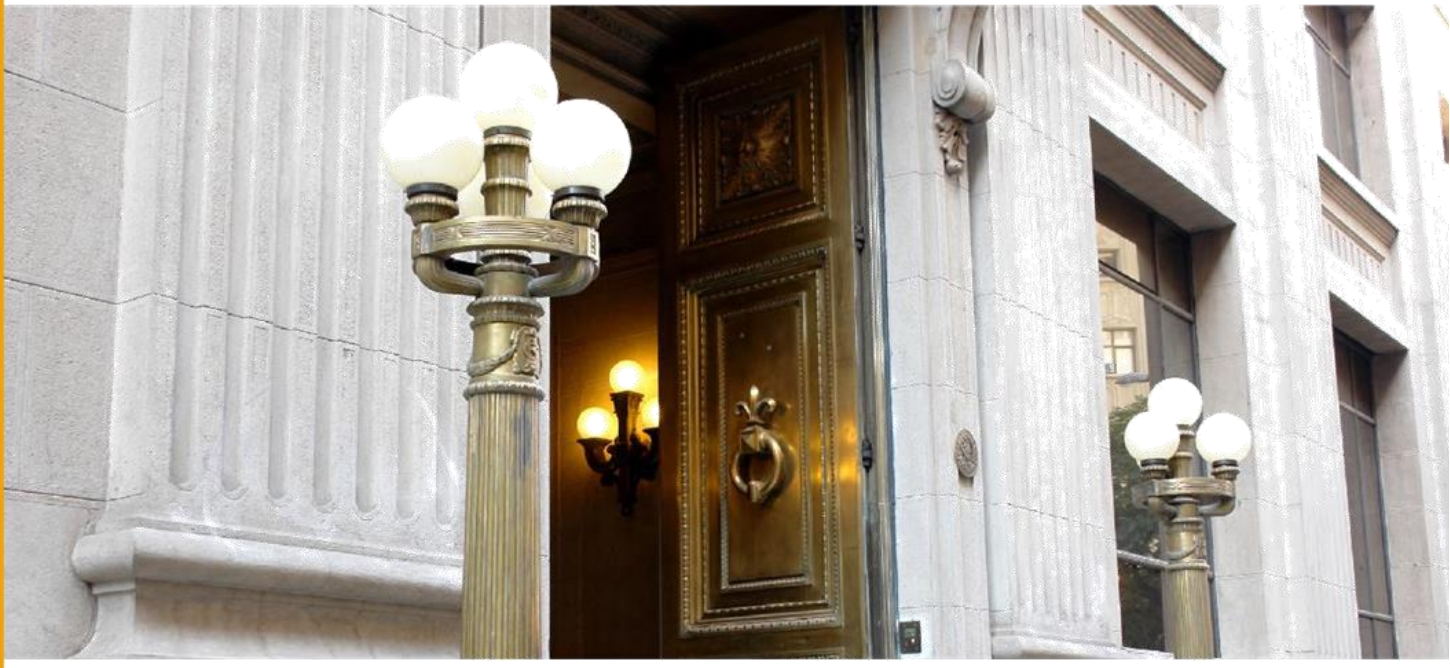
Elias Albagli
Luis Ceballos
Sebastian Claro
Damian Romero

N° 1007 Marzo 2024
BANCO CENTRAL DE CHILE



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Documentos de Trabajo del Banco Central de Chile
Working Papers of the Central Bank of Chile
Agustinas 1180, Santiago, Chile
Teléfono: (56-2) 3882475; Fax: (56-2) 38822311

Documento de Trabajo N° 1007 **Working Paper N° 1007**

UIP Deviations: Insights from Event Studies*

Elias Albagli†
Central Bank of Chile

Luis Ceballos‡
Knauss School of Business

Sebastian Claro§
Universidad de los Andes

Damian Romero¶
Central Bank of Chile

Abstract
We evaluate the behavior of the UIP relationship around monetary policy and global uncertainty shocks using event studies. We find that the covariance between exchange rate movements and changes in long-term yield differentials is conditional on the nature of shocks. A model of partial arbitrage between domestic and US bond markets predicts that tighter US monetary policy appreciates the dollar while increasing US yields relative to domestic bonds, a response that is consistent with UIP forces, while global uncertainty shocks appreciate the dollar while raising domestic yields relative to US bonds, exacerbating the widely documented UIP violation. The empirical analysis supports these mechanisms, specially for developed economies. For emerging economies, both relationships are weaker, consistent with more pervasive currency stabilization policies that mute the FX response at the expense of higher volatility in longer yields. Our results suggest a more nuanced interpretation of the unconditional failure of the UIP.

Resumen
Este trabajo evalúa el comportamiento de la paridad descubierta de tasas (UIP por sus siglas en inglés) alrededor de reuniones de política monetaria y eventos de incertidumbre global, usando una metodología de estudio de eventos. Encontramos que la covarianza entre las variaciones del tipo de cambio y variaciones en el diferencial de tasas largas depende de la naturaleza de los shocks. Presentamos un modelo de arbitraje parcial entre bonos domésticos y de EE.UU., el cual predice que una política monetaria contractiva en EE.UU. aprecia el dólar y aumenta las tasas de EE.UU. relativo a las tasas domésticas, lo cual es una respuesta consistente con la UIP. Por otro lado, shocks de incertidumbre global aprecian el dólar y aumentan las tasas domésticas relativo a las tasas de EE.UU., lo cual exacerba las violaciones de la UIP documentadas en la literatura. Nuestro análisis empírico provee evidencia de la validez de estos mecanismos, especialmente en economías desarrolladas. En economías emergentes ambas relaciones son más débiles, lo cual es consistente con políticas de estabilización cambiaria, cuyo costo es una mayor volatilidad de las tasas largas. Nuestros resultados sugieren una interpretación más benigna para las violaciones de la UIP de manera incondicional.

*The opinions and mistakes are our exclusive responsibility and do not necessarily represent the opinion of the Central Bank of Chile or its board. We thank the editor and two anonymous referees for valuable suggestions. We would also like to thank Yuriy Kitsul, Borja Larrain, Hyun-Song Shin, Ganesh Viswanath-Natraj, Xiaodong Zhu, and seminar participants at PBC School of Finance, Tsinghua University, the International Monetary Fund, the Czech National Bank, Universidad de Los Andes, Chile, 2022 Financial Management Association Conference, and the 2023 WEAI-IBEFA Summer Meeting.
†ealbagli@bcentral.cl; ‡luisceballos@sandiego.edu; §sclaro@uandes.cl; ¶dromeroc@bcentral.cl

1 Introduction

While the uncovered interest rate parity (UIP) has long been at the core of international macroeconomics, few relationships have received a starker rejection in empirical work.[1] The finding that high-interest rate currencies tend to generate higher ex-ante returns – the so-called forward premium – has been mainly interpreted as evidence of time-varying risk premia. (see, for example, Verdelham (2010), Lustig et al. (2014), Hofmann et al. (2019) and Kalemli-Özcan and Varela (2023)).[2] Supporting this explanation, several studies have found that, in response to risk-off events, the US dollar (USD) appreciates while at the same time US yields tend to compress relative to bonds denominated in other currencies, movements that are against the UIP logic.[3]

On the other hand, an extensive literature has documented that exchange rates do adjust in the direction suggested by the UIP condition in response to monetary policy innovations. For example, several studies document that the USD appreciates relative to developed economies´currencies either in response to direct measures of US MP tightening or following news that anticipates a move in such direction.[4] Some papers document an excessive reaction of exchange rates to interest rate differentials, which would imply a lower risk premium for high-interest rate currencies. This evidence, confronted by the failure of the UIP – which has been mostly interpreted as evidence of a high risk-premium in the high-interest rate currency – constitutes the essence of Engel Puzzle (Engel, 2016).

The evidence suggesting that the direction and magnitude of the forward premium are not stable over time and could be conditional on underlying shocks has motivated several further studies. Engel et al. (2022) show for major advanced economies that excess returns are highly correlated with inflation differentials, which anticipate interest rate movements. They find that as long as there is delayed FX adjustment or investors have difficulties in anticipating interest rate adjustments, the excess return increases in periods of higher inflation. Lustig et al. (2014) show evidence that the co-movement of excess returns with interest rate differentials responds to a higher risk premium by US investors of investing abroad whenever the US economy is weak. Kalemli-Özcan and Varela (2023) also documents a systematic cyclical pattern on excess returns for emerging economies, which is mostly linked to the global financial cycle and the policy responses of local central banks.

In this paper, we shed further light into these issues by explicitly evaluating the dependence of the UIP relationship to specific shocks across a broad set of emerging market economies (EMEs) and developed economies (DEVs). We provide systematic evidence of the response of (long-term) bond yield differentials and exchange rates to specific monetary policy and uncertainty shocks using

[1] See, for example, Hansen and Hodrick (1980), Meese and Rogoff (1983) and Fama (1984). For a recent survey, see Engel (2014).

[2] Another prominent explanation stresses biases of investor expectations. See for example Gourinchas and Tornell (2004) and Bachetta and van Wincoop (2010).

[3] See, for example, Caballero and Kamber (2019), Carrière-Swallow and Cespedes (2013) and Akinci et al. (2022).

[4] See Andersen et al. (2007), Clarida and Waldman (2008), Inoue and Rossi (2019) and Schmitt-Grohé and Uribe (2022).

an event study methodology. In particular, MP shocks are measured by changes in short-term rates around Monetary Policy meetings, and global uncertainty shocks are identified by sharp changes in the VIX. Our sample includes 24 countries split equally between DEVs and EMEs between November 2008 and December 2019.[5] To the best of our knowledge, the event study methodology is new in the UIP literature. Providing a common empirical methodology that systematically studies the potentially distinct joint-response of exchange rates and yield differentials to well-identified shocks is the main contribution of the paper. We also document that these responses are different between DEVs and EMEs, and further test some hypotheses about why this might be so.

To discipline our empirical strategy, we develop a model in which long-term bond markets are populated by domestic participants and US investors, where the latter react partially to deviations in the UIP condition vis–a–vis US yields. A tighter US MP raises US yields and induces capital outflows from the domestic economy. This depreciates its currency and raises long-term yields as domestic investors require lower bond prices to absorb the sales of outgoing US investors. Conditional on this shock, a contemporaneous depreciation of the domestic currency (and its ensuing appreciation) coincides with an increase in yield differentials in favor of the US, movements that are consistent with the textbook UIP relationship. In contrast, global uncertainty shocks lead to an increase in the demand for US bonds, driving down their yields relative to domestic bonds at the same time the dollar appreciates. Thus, asset prices move against the UIP logic, reflecting the impact of an increase in risk-premia.

Another important contribution of our work is the comparison of asset price responses for DEVs and EMEs using a single methodology. With a few exceptions (Bansal and Dahlquist (2000), Kalemli-Özcan (2019) Kalemli-Özcan and Varela (2023)), most studies of the UIP have been devoted to advanced economies only. Our paper not only uses the same event-study methodology for contrasting the effect of shocks on DEVs and EMEs, but it also interprets the different results in terms of specific parameters of the model. In our framework, the strength of exchange rate and bond yield responses to specific shocks depends on the degree of FX defensive policies, which dampen the effects of shocks on the exchange rate at the expense of amplifying the reaction of domestic bond yields. Another driver of the correlation between exchange rate movements and variations in yield differentials is the degree of country-specific capital flow volatility. The higher prevalence of these factors in EMEs will help to rationalize the differences with DEVs documented below.

The main advantage of the event-study strategy is its suitability for identifying the two key shocks highlighted by the model, but it does come at a certain cost. Namely, this approach is not well suited for evaluating the UIP using short-term bonds. This is because even if the rational expectations hypothesis were valid,[6] one cannot assume that the one-month expectation of the

[5] We focus on the post-GFC period as it witnessed a strong trend towards capital market integration and the presence of international investors in emerging bond markets. See Fratzscher (2012), Albagli et al. (2019), Miranda-Agrippino and Rey (2020a), Doidge et al. (2020) and Lilley et al. (2019) among others.

[6] See the survey by Engel (2014) for a discussion on the implications of deviating from rational expectations.

exchange rate at day $t-1$ – the day before the event – is well approximated by the $t+29$-day value of the exchange rate since such value will reflect the updated information set of investors after the shock. In other words, a correctly specified event study would need to observe not only spot exchange rate movements during the event window but also the change in exchange rate expectations at short maturities around each episode, which is not available.

To overcome this issue, we focus our analysis on yield differentials of long-term (10-year) bonds. Indeed, for long-maturity instruments, we show that changes in exchange rate dynamics over said maturities are well approximated by the instantaneous movement in the exchange rate –that is, the effect of shocks on expected exchange rates at longer horizons is negligible. We provide evidence to validate this methodological assumption using Consensus Forecast data, documenting that exchange rate fluctuations lead to a monotonically decaying exchange rate expectations "term structure" that approaches zero well before the 10-year maturity.[7] The vanishing response of exchange rate expectations at longer maturities holds both unconditionally as well as in response to specific shocks.

We highlight three main empirical results, all consistent with the model's key predictions. First, following US MP shocks around FOMC meetings, exchange rates, and yield differentials for long-term bonds react in the direction suggested by a UIP condition for most DEV. Specifically, country-j's currency appreciates against the USD whenever the 10-year yield differential of that country vis–a–vis the US widens around the bracketed event. An analogous result holds for domestic MP shocks in our DEV sample, albeit at a lower statistical significance. Second, as suggested by the model, exchange rates and yield differentials react in the opposite direction to the one suggested by the UIP around uncertainty shocks. Indeed, episodes of large VIX increases compress US long-term yields significantly more than elsewhere, while the USD contemporaneously appreciates against most currencies– with the notable exceptions of Switzerland and Japan, two well-known safe-haven countries.[8]

Finally, all these results are weaker for EMEs. In particular, we find a coefficient of essentially zero between exchange rate movements and changes in yield differentials around MP events (both US and domestic) and a negative but only marginally significant relationship between exchange rate adjustments and changes in yield differentials around uncertainty events. The model rationalizes these weaker responses with two possible explanations. On the one hand, the tendency of emerging markets to endogenously adjust policy to stabilize their currency, as documented by Sarno and Taylor (2001), Ghosh et al. (2016), Fratzscher et al. (2019), and Kalemli-Özcan (2019), limit the volatility of the exchange rate at the expense of more volatile capital flows and longer-term yields, flattening the equilibrium relationship between exchange rates and long yield differences. On the other hand, these results are also consistent with higher idiosyncratic volatility in EMEs

[7]Similarly, Froot and Ito (1989) document that short-term expectations overreact to exchange rate innovations relative to long-term expectations.

[8]See Ranaldo and Soderlind (2010), and Lilley et al. (2019).

that weaken the relationship between exchange rate and yield differences around the identified episodes.[9] This could reflect, among other factors, country-specific political uncertainty – including monetary, fiscal, expropriation and democratization risk considerations –, along the lines of Kalemli-Özcan and Varela (2023). Below, we provide additional empirical results suggesting that both FX interventions and domestic idiosyncratic shocks contribute to explaining the weaker results for EMEs.

Following Mueller et al. (2017), we explore an alternative empirical approach that measures the ex-post excess return of a one– and two–day investment strategy that goes long (short) on domestic (US) 10-year bonds and estimate how this excess return varies around monetary policy and uncertainty events. Consistent with our previous evidence, ex-post excess returns for DEVs do not differ significantly around US monetary events relative to other days, reflecting that yield differentials and exchange rates respond in the direction that a UIP relationship would suggest. In contrast, ex-post excess returns are significantly lower around days with uncertainty shocks relative to other days, as foreign currencies depreciate while at the same time, their yield differentials relative to US bonds increase, amplifying any pre-existing deviation of the UIP condition. For EMEs, the marginal increase in ex-post excess returns in response to US contractionary monetary policy shocks results from the more muted reaction of exchange rates in these countries. In response to uncertainty shocks, ex-post excess returns in EMEs fall even more than in DEVs, mainly as a consequence of the higher reaction of domestic longer-term yields– which is consistent with the model and the results of our baseline empirical specification.

Our work is related to several literatures. Several papers have documented how US monetary policy shocks affect global risk sentiment (usually measured using the VIX index), mostly using different vector autoregression (VAR) methodologies, as in Bekaert et al. (2013), Bruno and Shin (2015) and Miranda-Agrippino and Rey (2020b). For instance, a recent work by Kalemli-Özcan (2019) shows that US monetary policy tightening increases the UIP risk premia over time, which translates into a positive correlation between global risk sentiments and excess returns of strategies that are long in EM assets. In contrast, for a sample of developed economies, Bekaert et al. (2023) do not find evidence that US monetary policy shocks drive asset prices through a risk channel. Other papers directly study the impact of uncertainty shocks on asset prices and the UIP. Akinci et al. (2022) develop a model in which U.S. financial intermediaries, that hold foreign risky assets, operate under financial frictions. In response to a US uncertainty shock, these intermediaries deleverage and trigger a fall in risk appetite, increasing the UIP premium. Using data from Turkey, di Giovanni et al. (2022) show that the Global Financial Cycle drives the behavior of the UIP premium and that this is related to the exposure of Turkish banks to FX funding.

We complement this literature in two ways. First, we examine systematically for a group of developed and emerging economies how precisely identified monetary and uncertainty shocks affect

[9] In the empirical section we document a larger volatility of key financial variables in EME, both unconditionally and during event days.

the UIP premium. We find that, for developed countries, asset prices respond to US MP shocks in a manner consistent with the UIP logic, in contrast with the findings in the literature that these shocks amplify the failure of the UIP through a risk-taking channel. One way of rationalizing these differences is the time lag involved in the risk–premium channel documented using VAR methodologies, while our event–study approach focuses, by definition, on high-frequency responses around narrow time windows. Thus, we view these results as complementary to those in the literature. On the other hand, our results around precisely–identified events reinforce the findings of other studies that uncertainty shocks amplify the failure of the UIP in both DEVs and EMEs.

Our paper is also related to the literature that studies how domestic policies shape the response of the forward premium to monetary and uncertainty shocks. The sensitivity of UIP deviations to US monetary policy shocks is larger in EMEs, which Kalemli-Özcan (2019) attribute to the response of domestic monetary policy, that trying to reduce FX volatility ends up hurting the domestic economy, augmenting the effects of international risk spillovers. Kalemli-Özcan (2019) shows that when EME policymakers use their policies to limit exchange-rate volatility in response to US monetary policy shocks, the country-specific risk premium increases by tightening domestic financial conditions in such a way that EMEs need to provide additional returns to investors. An alternative perspective is developed in Itskhoki and Mukhin (2022) and Itskhoki and Mukhin (2023), who argue that UIP deviations result endogenously as risk-averse financial intermediaries require larger returns for their exposure to assets subject to high exchange-rate volatility. According to their results, financial interventions, i.e. FX interventions, can contribute to eliminate excess FX volatility, and hence contribute to the UIP stabilization.

Whether FX interventions contribute to mitigating or amplifying the deviations of the UIP is an important question. We contribute to this discussion in two dimensions. First, our model provides a simple conceptual framework to rationalize how interventions affect the response of asset prices to monetary and uncertainty innovations, showing that FX interventions mute the response of exchange rates at the expense of amplifying the reaction of domestic longer-term rates, therefore amplifying any preexisting forward premium. The second contribution is empirical. The scope of our analysis and its focus on DEVs and EMEs is novel, as it not only allows us to understand how the forward premium varies across specific events but also across countries. Our results confirm that the failure of asset prices to respond in line with the UIP forces is, at least partially, explained by a more active FX policy action by EMEs' central banks, which is in line with Kalemli-Özcan (2019).

Also, our paper contributes to a large literature in finance and international economics studying the behavior of asset prices around specific events. A growing body of evidence documents the significant impact of US MP shocks on a variety of asset prices. For example, Bernanke and Kuttner (2005) analyze the effect of US monetary policy shocks on the stock market. Hanson and Stein (2015) use the event-study methodology to analyze the effect of US monetary policy shocks on

US longer-term rates in the United States, while Albagli et al. (2019) focus also on the effect on longer-term yields in developed and emerging countries. Savor and Wilson (2014) using the event-study methodology to analyze the impact of announcement days (macro announcements as well as US monetary policy announcements) on stock returns. We contribute to this literature by extending the methodology to evaluate how US and domestic monetary policy shocks, as well as uncertainty shocks, affect the forward premium and its components –exchange rates and yield differentials.

Finally, our work is connected to recent literature analyzing the endogenous relationship between exchange rates and long-term yields (see, for example, Gourinchas et al. (2022) and Greenwood et al. (forthcoming)). In these papers –as in ours– limited bond market integration leads to movements in term premia (and as a consequence, long-term yields) and exchange rates through a mechanism driven by the UIP logic —investors' demand for bonds partially respond to perceived violations of the UIP relationship–, a framework first developed in Albagli et al. (2019). However, while the focus of Gourinchas et al. (2022) and Greenwood et al. (forthcoming) is mainly on the implications of unconventional monetary policies on yields and exchange rates, our paper evaluates the impact of conventional monetary policy shocks and uncertainty shocks on the forward premium puzzle, showing both theoretically and empirically that the behavior of the UIP puzzle fundamentally depends upon the nature of the shocks and the policy response of central banks. Hence, we see our contribution as complementary to theirs.

The rest of the paper is organized as follows. Section 2 presents the model and derives empirical predictions for the equilibrium relationship between long-term interest rates and exchange rate dynamics for different types of shocks, as well as different policy reactions. Section 3 describes the data, while sections 4 and 5 report our main empirical results around monetary policy and uncertainty events, respectively. Section 6 presents the evidence of the response of (ex-post) excess returns of different investment strategies to monetary policy and uncertainty shocks. Section 7 concludes.

2 A model of exchange rates and interest rate differentials

This section develops a model to formalize the relationship between interest rate differentials and exchange rate dynamics, conditional on different shocks and the role of endogenous policy reaction. The model is based on Blanchard et al. (2015) and expands on Albagli et al. (2019) by incorporating other sources of shocks besides US MP.

2.1 US MP and long-term US yields

US MP follows an autoregressive process, normalized at a long-run mean of zero,

$$i_t^* = m_t^* \,, \qquad \text{with} \quad m_t^* = \rho \cdot m_{t-1}^* + \varepsilon_t^* \quad \text{and} \quad \varepsilon_t^* \sim \mathcal{N}(0, \sigma_{\varepsilon^*}^2). \tag{1}$$

Besides the short-term bond that yields the MP rate i_t^*, there is a market for trading an h-year zero-coupon bond ($h = 10$ years in our empirical setup). The demand for the US h-year zero coupon bond has an endogenous component that depends positively on the yield, with elasticity β^*, and a price-inelastic term labeled $z_t^* = -i_t^* + v_t$. The first component of z_t^* depends on US MP. We refer to it as the "risk-taking channel" of US MP –the notion that a rise in the federal funds rate (FFR) is empirically associated with a risk-off movement away from long-term Treasuries, as documented by Hanson and Stein (2015). This component loads negatively on the US short-term rate (with a loading normalized to -1). The second component is a global "uncertainty shock", $v_t \sim \mathcal{N}(0, \sigma_v^2)$. In our empirical setting, these shocks correspond to swings in risk-on/risk-off sentiments, proxied by large VIX movements independent of US MP. Typically, a risk-off shock defined as $v_t > 0$ compresses US treasury yields. Normalizing bond supply to zero, we can solve for the US long-term bond yield as a function of US MP:

$$0 = \beta^* y_t^{*(h)} + z_t^* \quad \rightarrow \quad y_t^{*(h)} = -\frac{z_t^*}{\beta^*} = \frac{m_t^* - v_t}{\beta^*} \tag{2}$$

2.2 Country j block

Net capital inflows (K_t) into country j consist of foreign portfolio allocation into short-term and long-term (h-year) bonds. Each flow is proportional to the bond yield differential relative to its US equal-maturity counterpart, net of the expected depreciation rate of j's currency over the corresponding horizon. Let s_t be the (log of) US dollar value of one unit of domestic currency (an increase in s_t stands for an appreciation against the US dollar). K_t is given by:

$$K_t = \underbrace{(i_t - i_t^* - (s_t - E\left[s_{t+1}|\Omega_t\right]))}_{\text{Short-term bond}} + \underbrace{\left(y_t^{(h)} - y_t^{*(h)} - (s_t - E\left[s_{t+h}|\Omega_t\right])/h\right)}_{\text{Long-term bond: price-elastic}} + \underbrace{z_t}_{\text{Long-term bond: inelastic}} \tag{3}$$

where

$$z_t = -\lambda \cdot m_t - \kappa \cdot v_t + u_t, \tag{4}$$

is the price-inelastic capital flow into country-j's, which for simplicity we assume fully loads into the long-term bond market. We consider three components to z_t. The first allows for domestic MP to potentially trigger a risk-taking channel, with country-specific loading $\lambda \geq 0$. Second, price-inelastic capital flows are also determined by the global uncertainty shock v_t. While the shock is systematic to the international financial system, we assume a country-specific loading of $\kappa \geq 0$, so that a risk-off event ($v_t > 0$) will imply a retreat in capital inflows. The third term, $u_t \sim \mathcal{N}(0, \sigma_u^2)$, is a country-$j$

idiosyncratic shock. The introduction of a purely idiosyncratic shock serves as an orthogonal source of noise in the model, which as discussed below, may play a role in interpreting the observed correlations (or lack there-off) between yield differentials and exchange rates in an event-study setting. This idiosyncratic shock could be interpreted, for example, along the lines of Kalemli-Özcan and Varela (2023), who show that higher policy uncertainty in emerging economies (regarding monetary and fiscal policy, rule of law, or political uncertainty) is the main factor explaining the negative correlation between capital inflows and the UIP premium, even after controlling for global risk conditions. This is exactly the mechanism embedded in u_t. The vector Ω_t denotes the information set, common to all agents, which consists of all current state variables.

Country j's central bank sets the short-term interest rate according to:

$$i_t = m_t \text{ , with } m_t = \psi \cdot m_{t-1} + \varepsilon_t, \quad \text{and} \quad \varepsilon_t \sim \mathcal{N}(0, \sigma_\varepsilon^2). \tag{5}$$

Domestic MP is given by the shock m_t (a proxy for overall macroeconomic conditions, with persistence $0 < \psi < 1$). Besides traditional MP, central banks may also stabilize their currency with FX interventions (I_t), so we assume (following Blanchard et al. (2015)) an offset parameter ϕ (with $1 \geq \phi \geq 0$), such that $I_t = -\phi \cdot K_t$. Finally, dollar flows from international trade depends negatively on the domestic exchange rate, $CA_t = -\gamma \cdot s_t$. These relations lead to the following balance of payments equilibrium condition:

$$K_t + I_t + CA_t = 0. \tag{6}$$

We close the model with the domestic long-term bond market. We assume that domestic investors respond positively to long-term yields with elasticity β, irrespective of exchange rate dynamics (for example, pension funds targeting returns in domestic currency). The demand of foreign (US) investors is given by the long-term bond components of Eq. (3), and the critical assumption is that it responds positively to deviations from a version of the UIP condition for long-term yields. In other words, the difference between the domestic yield and the yield of US long-term bonds ($y_t^{(h)} - y_t^{*(h)}$), net of the expected depreciation of the domestic currency over the corresponding horizon. Assuming a zero net supply of the bond, we obtain the following market-clearing condition:

$$0 = \beta \cdot y_t^{(h)} + \underbrace{\left(y_t^{(h)} - y_t^{*(h)} - (s_t - E\left[s_{t+h}|\Omega_t\right])/h\right)}_{h\text{-period UIP deviation}} + z_t, \tag{7}$$

Eq. (7) states that an increase in the foreign demand for domestic bonds (due to a positive yield differential against the h-year US bond) must be accommodated by lower demand from domestic investors, inducing a fall in yields in equilibrium. This condition, therefore, links domestic yield movements with developments in the US long-term bond market and the other shocks of the model.

2.3 Equilibrium

Using Eq. (6), the determination of the current account, and the FX intervention rule, one can solve for exchange rate s_t as a function of the primitive shocks of the model: US MP shock (m_t^*); domestic monetary policy shock (m_t); the global uncertainty shock (v_t), and the country-specific shock (u_t). Iterating forward the time $t+h$ expectation of future exchange rates as a function of these variables yields the following expression:

$$s_t = -a_1 \cdot m_t^* + a_2 \cdot m_t - a_3 \cdot v_t + a_4 \cdot u_t, \tag{8}$$

where the coefficients are solved in closed form in the Appendix A.

While the sign of some coefficients is ambiguous ex-ante, the evidence below provides valuable guidance. The coefficient a_1, which measures the effect of US MP shocks on the exchange rate, is always positive so that an increase in US short-term rates always appreciates the US dollar. Intuitively, a contractionary US MP shock leads to capital outflows from the domestic economy into the US (an endogenous response to deviations in the UIP condition), which requires a depreciation of the domestic currency to clear the balance of payment condition and reestablish equilibrium in the longer–term bond market. The FXI parameter ϕ mitigates the effect of US MP (and other shocks) on the exchange rate by directly compensating part of the outflows or counteracting them by offering higher returns.[10]

The impact of domestic monetary policy shocks on the exchange rate is given by coefficient a_2. Consider first the case when $\lambda = 0$, implying that domestic monetary policy shocks do not trigger any investors' risk-taking response. In this case, a_2 is unambiguously positive so that a tightening of the domestic policy increases domestic longer-term rates at the same time it strengthens the domestic currency, as predicted by UIP forces. This relationship, however, might reverse if a contractionary domestic MP leads investors to anticipate a deterioration in domestic fundamentals and command considerable compensation for risk. In the model, the case $\lambda > 0$ captures this possibility, which means that a contractionary shock triggers outflows that can depreciate the domestic currency at the same time the domestic longer-term yield increases. This apparently counterintuitive reaction of exchange rates to domestic MP shocks has been documented elsewhere for EMEs and it will be confirmed in our data for this group of countries as well.[11]

Regarding an uncertainty shock, a "risk-off" episode (a positive v_t) will depreciate the domestic currency ($a_3 > 0$) as long as the loading of capital flows in response to the uncertainty shock (κ)

[10] A similar result would result if we allow the domestic monetary policy to endogenously react to exchange rate movements (for example, due to inflationary pressures or financial stability concerns). In this case, equation 5 in the text becomes $i_t = -d \cdot s_t + m_t$, where d is a country-specific parameter that measures the sensibility of domestic MP to exchange rate movements.

[11] See Kohlscheen (2013), Hnatkovska et al. (2016), and Kalemli-Özcan (2019).

is sufficiently large. Indeed, a large value for κ triggers capital outflows that depreciate the local currency, despite the compression in the US long-term yield that follows a risk-off event.[12] Finally, $a_4 > 0$, reflecting that a purely idiosyncratic capital flow shock into the local bond market will appreciate the domestic currency and lower the longer-term domestic yield.

To solve for the yield differential $y_t^{(h)} - y_t^{*(h)}$, we iterate forward Eq. (8), use US-yields from Eq. (2), and replace both in the domestic bond-market clearing Eq. (7):

$$y_t^{(h)} - y_t^{*(h)} = (1-b)\Bigg[- m_t^* \left(\beta/\beta^* + \frac{a_1(1-\rho^h)}{h} \right) + m_t \left(\lambda + \frac{a_2(1-\psi^h)}{h} \right) + v_t \left(\beta/\beta^* + \kappa - \frac{a_3}{h} \right) - u_t \left(1 - \frac{a_4}{h} \right). \Bigg] \tag{9}$$

According to Eq. (9), tighter US monetary policy (a rise in m_t^*) reduces the domestic yield differential through the direct increase in $y_t^{*(h)}$ (Eq. 2). The effect of US MP on yield differentials also depends on policy reaction, as reflected in a_1. If central banks aggressively intervene in the FX market, a_1 will be small, enhancing the upward reaction of domestic yields and closing the gap relative to the US yield. Therefore, it is always the case that $\partial(y_t^{(h)} - y_t^{*(h)})/\partial m_t^* < 0$.

The effect of domestic MP policy on long-term yield differentials will be always positive ($\partial(y_t^{(h)} - y_t^{*(h)})/\partial m_t > 0$), meaning that a tightening of the domestic monetary policy increases domestic longer yields and have no impact on US longer-term yields. This effect is amplified the larger λ is, as a higher risk-taking behavior of investors following a domestic MP shock triggers more capital outflows, leading to a larger increase in the domestic longer–term yield. The data below will generally confirm this positive relationship. Also, a risk-off event ($v_t > 0$) will always increase the domestic yield with respect to the US ($\partial(y_t^{(h)} - y_t^{*(h)})/\partial v_t > 0$), which is generally true in our sample with the exception of the safe heaven examples. Finally, an increase in the idiosyncratic capital flow ($u_t > 0$) will, by construction, increase domestic yields while leaving constant US yields, so that ($\partial(y_t^{(h)} - y_t^{*(h)})/\partial u_t > 0$).

2.4 Linking the model to the empirical strategy

A basic UIP test for horizon h in this setup can be expressed by:

$$(s_t - E\left[s_{t+h}|\Omega_t\right])/h = a_0 + b_0(y_t^{(h)} - y_t^{*(h)}) + \varepsilon_t. \tag{10}$$

[12] If κ is small enough the fall in US yields (relative to domestic yields) following risk-off shocks will dominate the pressures on capital outflows from the domestic economy, so the domestic currency will appreciate relative to the USD. This condition would be valid for "safe havens" countries, i.e., $a_3 < 0$

Our ultimate object of interest is the correlation between the exchange rate and interest rate differentials (coefficient b_0), conditional on the presence of identifiable, high-frequency shocks. This amounts to testing Eq. (10) in differences around two close dates, $t_1 < t < t_2$, or:

$$\left(\Delta s_t - \Delta E\left[s_{t+h}|\Omega_t\right]\right)/h = b_0 \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_t, \tag{11}$$

where $\Delta x_t = x_{t_2} - x_{t_1}$ for variable x, and $\nu_\tau \equiv \Delta\varepsilon_t$. Although this specification should be valid at any horizon h, its estimation over short horizons in response to specific events would require high-frequency measures of exchange rate expectations at said horizons around these events, which are not available. Furthermore, an event-study approach designed to condition on particular shocks cannot apply the rational expectations hypothesis and simply proxy $E\left[s_{t+h}|\Omega_t\right]$ by its realized value s_{t+h}. For example, the one-month expectation of the exchange rate at day $t-1$ –the day before a particular event– is obviously not well approximated by the $t+29$-day value of the exchange rate, as such value will undoubtedly incorporate the updated information set of investors *after* the shock occurs.

In our empirical strategy, we will overcome this problem by focusing on long horizons ($h = 10$ years), and assuming $\Delta E\left[s_{t+h}|\Omega_t\right] \to 0$. Appendix C presents evidence validating this assumption. Using monthly Consensus Forecast data for a broad sample of developed and emerging countries, we estimate regressions in which the change of exchange rate expectations at different horizons is a function of changes in spot exchange rates. We find that changes in spot exchange rates lead to revisions in exchange rate expectations that decline monotonically with maturity. Importantly, this exchange rate expectations "term-structure" converges to zero for forecasts more than a few years ahead.[13] As additional validation, we estimate regressions between exchange rate expectations at different horizons and our measures of monetary policy shocks, finding a similar monotonically decaying pattern that converges to zero well before 10 years.

We then get the following equation for estimating the relationship between exchange rates and long-term yield differences in response to specific shocks:

$$\Delta s_t/h = \hat{a} + \hat{b} \cdot \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_t. \tag{12}$$

The event-study strategy assumes that we can separately identify the shocks $x = [m_t^*, m_t, v_t]$, an assumption validated by the minimal overlap between events, as explained below. In particular, the estimated coefficient $\hat{b}|_x$, conditional on shock x, is given by

$$\hat{b}|_x = \frac{Cov\left(\Delta s_t/h; \Delta(y_t^{(h)} - y_t^{*(h)})|x\right)}{Var\left(\Delta(y_t^{(h)} - y_t^{*(h)})|x\right)}. \tag{13}$$

[13] See Froot and Ito (1989) for a similar result.

The event window will nevertheless be contaminated by some idiosyncratic volatility, formally introduced in the model by the country-specific shock u_t. The presence of this shock may affect the strength and even the sign of the conditional correlation between interest rate differentials and exchange rates. Using Eqs. (8), (9), and (13) we can now compute the $\hat{b}|_x$ coefficients (solved in closed form in Appendix A).[14] The sign and dependence on model parameters of these coefficients are summarized in Proposition 1, which guides the testable predictions below:

Proposition 1 (UIP, conditional on shocks)**.** *Consider the limit where $(1-b)/h \to 0$, and $1/h^2 \to 0$,*

a) US MP shocks: i) Sign: $\hat{b}|_{m^*} > 0$ iff $\sigma^2_{m^*}(1/b + 1/\beta^*)\tilde{R}_1(\beta/\beta^*) > \sigma^2_u$, where $\tilde{R}_1 \geq 1$ depends on model parameters. ii) Effect of policy: $\partial\hat{b}|_{m^*}/\partial\phi < 0$ iff $\sigma^2_{m^*}(1/b + 1/\beta^*)\tilde{R}_1^2(\beta/\beta^*) > \sigma^2_u$. iii) Effect of noise: $\partial\hat{b}|_{m^*}/\partial\sigma^2_u < 0$.

b) Domestic MP shocks: i) Sign: $\hat{b}|_m > 0$ iff $\sigma^2_m\lambda(1/b - \lambda)\tilde{R}_2 > \sigma^2_u$, where $\tilde{R}_2 \geq 1$ depends on model parameters. ii) Effect of policy: $\partial\hat{b}|_m/\partial\phi < 0$ iff if $\sigma^2_m\lambda(1/b - \lambda)\tilde{R}_2^2 > \sigma^2_u$. iii) Effect of noise: $\partial\hat{b}|_m/\partial\sigma^2_u < 0$ iff $\lambda(\lambda(\tilde{R}_2 - 1) - 1/b\tilde{R}_2) < 0$.

c) Uncertainty Shocks: i) Sign: $\hat{b}|_v < 0$ iff $\sigma^2_v(\kappa - 1/\beta^*)(\beta/\beta^* + \kappa) + \sigma^2_u > 0$. ii) Effect of policy: $\partial\hat{b}|_v/\partial\phi > 0$ iff $\hat{b}|_v < 0$. iii) Effect of noise: $\partial\hat{b}|_v/\partial\sigma^2_u < 0$ iff $\beta/\beta^* + \kappa > 0$.

Proof. See Appendix A. □

Proposition 1 gives the key prediction of the model regarding the sign of the correlation between the exchange rate and interest rate differentials, conditional on different shocks. Part a) describes the sign of the correlation coefficient conditional on US MP shocks and its dependence on parameters. A contractionary US MP shock will increase US relative yields (lowering $y_t^{(h)} - y_t^{*(h)}$) and through the ensuing capital outflow, depreciate the domestic currency against the USD (lower s_t), thus producing a positive correlation coefficient $\hat{b}|_{m^*}$. The presence of idiosyncratic noise not controlled for within the event window may weaken and even overturn the sign of the correlation, as part a.iii) makes explicit, so that a positive coefficient $\hat{b}|_{m^*}$ requires an upper bound for σ^2_u, as stated in part a.i) of the proposition.[15]

Part a.ii) of Proposition 1 describes how the magnitude of $\hat{b}|_{m^*}$ is affected by outright FX interventions (higher ϕ), which mitigate the response of the exchange rate while amplifying the response of domestic yield $y_t^{(h)}$. The higher the degree of intervention the lower the absolute value of $\hat{b}|_{m^*}$ is. This mechanism will be important when exploring the potential causes of a negative UIP relationship for some countries in our EME sample, conditional on US MP shocks.

[14] The model-implied coefficient $\hat{a}$ is zero since all shocks and hence steady-state variables have a zero-mean. In all regressions that follow, the coefficients $\hat{a}$ are included in the empirical specifications but not reported, as they are always not statistically different from zero.

[15] This occurs since the idiosyncratic capital inflow naturally induces a negative correlation between local exchange rates (which appreciate with inflows) and yield differentials against the US (which compress).

Part b) of the proposition states analogous conditions for the sign and parameter dependence of the coefficient associated with domestic MP shocks. For a positive value $\hat{b}|_m$, a first restriction is $\lambda(1/b - \lambda) > 0$, meaning that, whenever $\lambda > 1/b$, the price-inelastic outflows are large enough to depreciate the local currency in response to a tighter local MP, at the same time as long-term yield differentials against the US are increasing. As we will see, this is a common occurrence in our EME sample. Part b.i of the proposition states that, in addition, a positive sign $\hat{b}|_m$ also requires a limited amount of noise from the idiosyncratic shock u_t. As Part b.iii) shows, that the presence of noise reduces the correlation and can even overturn it (i.e., even for $\lambda(1/b - \lambda) > 0$, large enough σ_u^2 will eventually shift the sign of $\hat{b}|_m$ to negative).[16]

Part c) of Proposition 1 analyzes uncertainty shocks. Part c.i) establishes $(\kappa - 1/\beta^*)(\beta/\beta^* + \kappa) > 0$ as a sufficient condition for $\hat{b}|_v$ to be negative, which is satisfied as long as $\kappa > 1/\beta^*$, which is the necessary condition for the local currency to depreciate against the USD after an increase in global uncertainty (a risk-off event). Part c.ii) shows that FXI policies mitigate the exchange rate effect and thus lower the (absolute value) of the coefficient $\hat{b}|_v$, while part c.iii) establishes conditions under which higher idiosyncratic noise makes $\hat{b}|_v$ more negative.

To illustrate the key predictions of the model, Figure 1 simulates the combination of local exchange rates (vertical axis) and long-term yield differentials (horizontal axis) in response to shocks, for different parameter combinations. We start with a baseline case ($\hat{b}|_{m^*} > 0$, $\hat{b}|_m > 0$, and $\hat{b}|_v < 0$) in the left plot of each panel. Then, we study the effects of a higher degree of policy intervention (central plots) and higher idiosyncratic noise (on top of higher intervention, right plot). As expected, policy intervention dampens the correlations, while higher volatility of idiosyncratic shocks enhances their tilt into negative territory. Indeed, a combination of such elements (likely to be more prevalent in many EMEs in our sample) can imply a negative sign of the correlations for MP shocks (both US and domestic).

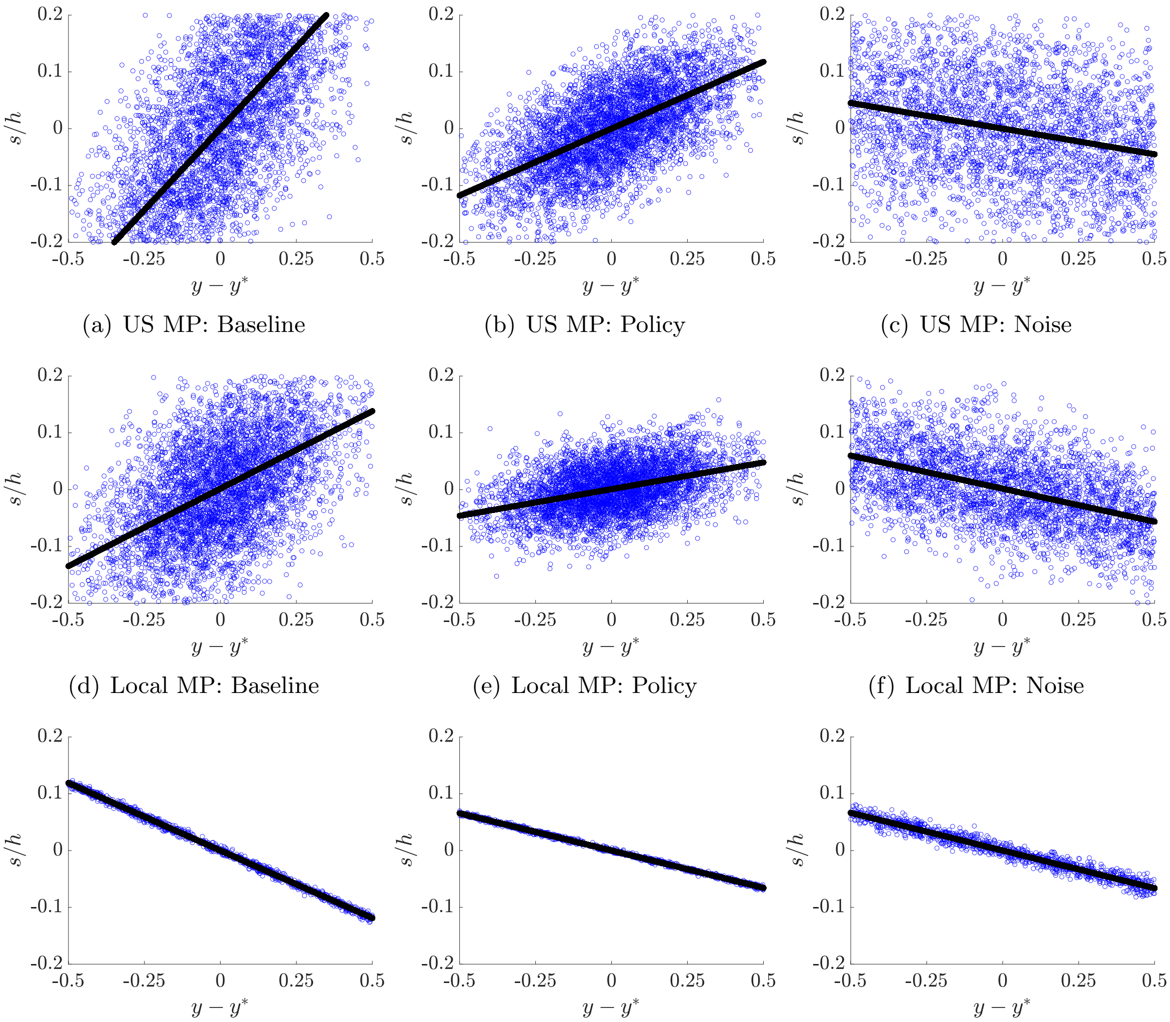
3 Data

We obtain daily data from Bloomberg[17] for 2- and 10-year yields for the US and a sample of 12 DEV and 12 EME and exchange rates (dollar per domestic currency) for November 2008–December 2019.[18] We complement this information with the dates of monetary policy meetings, also from Bloomberg, double-checked with the official dates reported by the different central banks. Appendix Table B.1 describes the countries included in the sample, their classification (DEV vs. EME), and

[16] Part b.ii) also establishes conditions regarding the sign of the effects of policies, although we do not emphasize these results as it makes less sense to think about domestic central bank interventions in response to their own MP decisions.

[17] The exceptions are Chile and Colombia, where yields are obtained directly from the central bank of each country.

[18] Specifically, our sample starts on November 24, 2008, previous to the announcement of the first quantitative easing (QE1) by the Federal Reserve.



(a) US MP: Baseline (b) US MP: Policy (c) US MP: Noise

(d) Local MP: Baseline (e) Local MP: Policy (f) Local MP: Noise

(g) Uncertainty: Baseline (h) Uncertainty: Policy (i) Uncertainty: Noise

Figure 1: Simulated exchange rates and yield differentials

Notes: This figure simulates exchange rates and yield differentials conditional on shocks. The horizontal axis corresponds to the long-term interest rate differential ($y_t - y_t^*$), while the vertical axis is the (maturity-normalized) exchange rate, s_t/h. The upper panel plots the relationship for US MP shocks, under baseline parameters (left plot), with a policy of FX interventions ($\phi = 0.9$, center plot), and with a mix of policy intervention and higher degree of idiosyncratic noise ($\phi = 0.9$ and $\sigma_u = 2$, right plot). The center and bottom panels do analogous comparative statics for local MP shocks and uncertainty shocks. Baseline simulation parameters are as follows: i) volatility parameters: $\sigma_{m^*} = \sigma_m = \sigma_v = 1$, $\sigma_u = 0.5$; ii) US block parameters: $\rho = 0.5$, $\beta^* = 5$; iii) Local price-inelastic flows parameters: $\lambda = 0.5$, $\kappa = 6$; iv) Local macroeconomic parameters: $\gamma = 0.1$, $\psi = 0.5$, $\phi = 0$, $\beta = 3$.

the number of monetary policy meetings for each country in the sample period.[19] Finally, we get daily data for the VIX from the Federal Reserve Bank of St. Louis' Database (FRED).

We define a US MP event as an FOMC meeting and a domestic MP event as each country's corresponding monetary policy meeting (MPM) date. For each event at day t, the corresponding MP shock is the change in the two–year bond yield around the event, computed as the difference between the closing values at dates $t-1$ and $t+1$. Therefore, the US MP shock is computed as the change in the two–year treasury yield between dates $t-1$ and $t+1$, and for domestic MP shocks, we compute the change in two–year government bond yields for each country.[20] Likewise, we compute the difference in 10-year yields and exchange rates corresponding to the closing values at dates $t-1$ and $t+1$. For risk premium events, we define an *uncertainty event* as a date in which the VIX has a daily variation (either positive or negative) larger than two standard deviations (computed on the daily change in the period January 2003 through December 2019). Then, we define the uncertainty shock as the 2-day differential in the VIX around those days.[21]

TABLE 1: Monetary policy meetings overlap

	FOMC	VIX	FOMC+VIX
DEV	3.88	4.91	0.17
EME	2.12	4.07	0.16

NOTES: This table presents the overlap frequency between the number of domestic monetary policy meetings, FOMC meetings in the US, and VIX events, as a fraction of domestic monetary policy meetings for each group of countries (in percentage points). The number of monetary policy meetings in developed (emerging) countries is 1161 (1228). Sample: November 24, 2008-December 31, 2019.

Table 1 documents the overlap between events, reporting the share of MPMs in DEV and EME that coincide with FOMC or VIX events. For instance, the first row in Table 1 reveals that only 3.88% of monetary policy events in DEV coincided with FOMC meetings, 4.91% with a VIX event, and 0.17% coincided with an FOMC meeting and a VIX event at the same time. The low degree of overlap in DEV and EME is a critical first validation of our empirical strategy.

We focus on the period after the global financial crisis because, as documented elsewhere, this is a period of strong capital market integration and foreign presence in EME bond markets (Fratzscher, 2012; Albagli et al., 2019; Miranda-Agrippino and Rey, 2020a; Doidge et al., 2020; Lilley et al., 2019).

[19] The list of selected countries responds to data availability. Because our focus is on the UIP relationship around events, it is critical to collect daily data for the analysis. This differentiates our work relative to previous literature, which covers a more extended period but uses lower frequency data (e.g., Hassan and Mano (2019)).

[20] For example, for the FOMC meeting held on Wednesday, December 11, 2019, the US MP shock is computed as the difference between the 2-year treasury yield at the close of Tuesday, December 10 and the close of Thursday, December 12.

[21] All results hold when we compute the standard deviation of the daily change in the VIX for the November 2008-December 2019 sample. In Appendix E we propose a series of robustness checks for our definition and identification of VIX events.

Our baseline estimations end in December 2019, excluding the year 2020, to avoid two features of the Covid-19 crisis that do not match well with the assumptions of our empirical strategy. First, during 2020 –especially during March– there is a large simultaneity of FOMC, MPM, and VIX events at the peak of the financial market turmoil. These were not random but a response to surging uncertainty and deteriorating prospects for the world economy, leading to extraordinary and coordinated MP actions. Indeed, in the 7-day window starting on March 10, 2020, 9 countries in our sample held extraordinary MPM (including the Federal Reserve). At the same time, the VIX had several of the most considerable swings in the whole sample. In particular, the coincidence of these events introduces difficulties in the identification of the shocks as discussed previously (see Appendix E.2).

Also, the uncertainty events in March 2020 led to significant increases in demand for US dollars and US short-term securities and a fall in demand even for US long-term debt and other countries' long-term securities. For example, between March 9 and March 18, 2020, the US 10-year Treasury yield increased by 64 basis points (bps). As *The Economist* put it, the priority for investors was to "liquidate holdings of risky assets, like stocks and high-yield bonds, and buy safe assets like Treasuries" but, "when their need for cash became dire, they dumped even these (Treasuries)".[22] This extraordinary market response contrasts with long-held features of financial markets and the model's basic assumption, namely, risk-off events increase demand for long-term US Treasuries. In appendix E we report the results of an extension of all the empirical analyses using the sample from November 2008 to November 2020, which overall coincide with our baseline estimations. Nevertheless, we prefer to keep the baseline specification excluding the Covid-19 episode to avoid the contamination of such extraordinary events.

Table 2 presents descriptive statistics of our data around each event for the period November 2008–December 2019.[23] We separate the analysis into three sets of columns. The first two show the mean and standard deviation of 10-year yields (Panel A) and exchange rates (Panel B) around different events for DEV and EME separately. For each group of countries, we report a test of variance that compares the volatility of yields and exchange rates around specific events (FOMC, MPM, VIX, and non-event days) with those observed unconditionally. The last set of columns presents the F statistic and the p-value of a variance test between DEV and EME, conditional on the events, where the null hypothesis is that those variances are equal.

[22] Why America's financial plumbing has seized up, *The Economist*, March 21, 2020 edition.

[23] During this period, we have 89 FOMC meetings, 1161 and 1228 domestic MPM for DEV and EME, respectively, and 139 VIX events. As shown in table 1, there is limited overlap across these events, both for DEV and EME.

Table 2: Descriptive statistics

	DEV (1)		EME (2)		Variance test (DEV/EME) (3)	
	Mean	Std. Dev.	Mean	Std. Dev.	F	p-value
Panel A: Ten-year yields						
FOMC	-1.18	6.97**	-1.07	16.00**	0.19	0.00
MPM	-0.40	7.95***	-1.28	18.98***	0.18	0.00
VIX	-1.77	9.42***	-0.50	24.72***	0.15	0.00
No event	-0.01	6.10***	-0.10	15.83***	0.15	0.00
Unconditional	-0.20	6.61	-0.26	16.82	0.15	0.00
Panel B: Nominal exchange rate						
FOMC	1.03	133.67***	2.52	119.79***	1.25	0.00
MPM	-3.02	125.39***	-1.03	101.57***	1.52	0.00
VIX	-13.35	148.93***	-22.15	148.19***	1.01	0.84
No event	1.45	82.42***	1.28	86.71***	0.90	0.00
Unconditional	-0.24	94.65	-0.78	95.88	0.97	0.02

Notes: This table presents descriptive statistics for 10-year yields and exchange rates in DEV and EME around events for the period November 2008–December 2019. Panel A (B) presents the mean and standard deviation for the 10-year yield differentials (exchange rates) around specific events in basis points. Column (1) presents the statistics for DEV, while column (2) presents the statistics for EME. Columns with standard deviations also show the statistical significance of variance tests of each event against unconditional days, where the null hypothesis is that those variances are equal. Finally, column (3) presents the F statistic and the p-value of a test of the difference in variance between DEV and EME, conditional on events, where the null hypothesis is that those variances are equal. ***, **, and * denote statistical significance at the 1, 5, and 10% levels, respectively.

In Panel A, we show statistics for variations in 10-year yields around events. To grasp the relative magnitude of the impact associated with different events, we take the unconditional volatility as the baseline and report whether volatility around other events is statistically different from such a baseline. For example, long-term yield volatility in DEV around all FOMC events is 6.97 bps, statistically higher than the unconditional volatility. A similar pattern arises around domestic MPM and VIX events, with a standard deviation of long-term rates in DEV (7.95 and 9.42 bps, respectively) that is statistically higher (at 1% confidence level) than the unconditional volatility. In contrast, volatility in days without events is statistically more negligible. In our EME sample, a similar pattern emerges, but with a much higher level of volatility compared to DEV across all events (two to three times as large), as formally reported through the F-statistics in column (3).

Panel B reports the behavior of exchange rates around these same events. For DEV, exchange rate volatility (measured in basis points) is significantly higher on event days than unconditionally. We observe a similar pattern in EME. Interestingly, compared to DEV, volatility in no-event days is higher in EME. At the same time, FX volatility during FOMC and domestic MPM events is minor (there is no statistical difference in VIX days). The crucial remarks are that long-term yields in

EME are much more volatile than in DEV during MP and VIX events, while exchange rate volatility is lower or similar around these same events. These observations suggest a more defensive policy reaction in EME, either through FXI or domestic MPM. Below, we come back to this point in the context of our main empirical results.

4 Monetary Policy Shocks

4.1 FOMC shocks

Panel A of Table 3 reports the results of Eq. (12), $\Delta s_t/h = \hat{a} + \hat{b} \cdot \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_{t,t+h}$ around US MP events.[24] To delve further into the drivers of this relationship, panels B-D report the results of US MP shocks on each component separately. In particular, each panel reports, respectively:

$$B: \ \frac{\Delta s_t}{10} = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_t; \ C: \ \Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_{t,t+h}; \ D: \ \Delta y_t^{*(h)} = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_{t,t+h}.$$

which are the empirical counterparts of expressions (8), (9) and (2) of our model, conditional on US MP shocks. As mentioned above, our preferred measure of US MP shocks, following Hanson and Stein (2015), is the change in 2-year US Treasury yield around each FOMC meeting. All variables are measured as changes between the market close at date $t+1$ and date $t-1$, where date t corresponds to the day of the event –in this case, the day of the FOMC meeting where the FFR is set. As a robustness check, we have tried different estimates of US monetary policy shocks, as discussed in Gürkaynak et al. (2005), Nakamura and Steinsson (2018), and Bu et al. (2021). None of the results of the paper changes using alternative definitions of US MP shocks (results are available upon request).

For DEV, Panel A documents a significant and positive coefficient between changes in exchange rates (an increase is an appreciation of the local currency) and the change in the 10-year yield differentials (domestic minus US yields) around FOMC meetings. In other words, US MP shocks that lower 10-year US yields (and thus increase yield differentials) also appreciate domestic currencies against the USD, consistent with the traditional UIP relationship. A 1% increase in the long-term interest rate differential leads to a highly significant dollar depreciation of 4.7% (which divided by $h = 10$ gives the 0.47 coefficient in the table, statistically significant at 1%). In contrast, the coefficient has the opposite sign for EMEs, and it is only marginally significant at 10% confidence levels.[25]

To shed more light on the results for EMEs, we analyze the response of exchange rate and yield differences to US monetary policy shocks separately. Panel B of the table shows that, in line

[24] The empirical strategy in this section and the following ones takes to the data Eq.(12) around different events. For space considerations, we do not report regressions that include all shocks simultaneously (and adequately orthogonalized). None of the results of the paper change, which is not surprising considering that the overlap of events is very small. These results are available upon request.

[25] These numbers roughly coincide with those in Zhang (2022), who show that in response to a 100 bps Fed contraction, foreign currencies depreciate on average 9.7% against the US dollar.

Table 3: UIP around US MP shocks

	$\hat{b}$	SE	R^2	N
Panel A (UIP): $\Delta s_t/h = \hat{a} + \hat{b} \cdot \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_t$				
DEV	0.47***	(0.08)	0.01	34584
EME	-0.10*	(0.06)	0.00	34584
Panel B (Exchange rates): $\Delta s_t/10 = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_t$				
DEV	-1.00***	(0.07)	0.01	34584
EME	-0.59***	(0.08)	0.00	34584
Panel C (10-year yield differentials): $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_t$				
DEV	-0.97***	(0.06)	0.02	34584
EME	-0.94***	(0.09)	0.00	34584
Panel D (US 10-year yield): $\Delta y_t^{*(h)} = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_t$				
US	1.41***	(0.17)	0.03	2882

Notes: This table presents regressions around FOMC events. s_t is the exchange rate (USD per unit of domestic currency), $y_t^{(h)}$ and $y_t^{*(h)}$ are the local and the US 10-year yields, respectively, and i_t^* is the 2-year US yield. In all specifications $\Delta x_t = x_{t+1} - x_{t-1}$, when in t there is a FOMC meeting. Each panel presents pooled regressions for the corresponding group of countries (DEV, EME, and the US), with standard errors clustered at the country level. ***, ** and * denote statistical significance at the 1, 5 and 10% levels, respectively.

with intuition, contractionary US MP depreciates domestic currencies, consistent with the model's prediction when central bank policy responses do not mute entirely the effect on exchange rate movements. Specifically, a 1% increase in 2-year US rates around FOMC events depreciates DEV currencies around 10% (significant at 1%), but only by 5.9% for EME currencies (also significant at 1%). Panel C presents the results for the 10-year yield differentials. We find that long-term interest rate differentials in DEV and EMEs exhibit a statistically significant impact of similar magnitude: a positive US MP shock of 1% triggers a reduction of the long-term interest rate differential in 0.97% and 0.94% for DEV and EME, respectively. Panel D shows that a more considerable increase in the US 10-year yield (of about 1.4%) relative to domestic yields explains such a reduction.

The evidence shows that exchange rate and interest rate differentials move in the opposite direction of what a parity condition would suggest for EME. However, this does not seem to derive from a particular violation of the signs of either of its underlying relationships: a positive US MP shock both depreciate EME currencies and increases yield differentials in favor of US treasuries. Instead, we can posit two non-competing hypotheses based on our model to explain the non-significant relationship. First, the lower coefficient of US MP on EMEs exchange rates may reflect a stronger FXI policy reaction by central banks. As shown in Proposition 1, part b), policy interventions indeed lower the regression coefficient between exchange rates and interest differentials by muting (at least partially) the exchange rate response. This interpretation is also consistent with the evidence given by Table

2, panel B). Notice that while exchange rates are significantly more volatile during non-event days in EMEs than in DEVs, they are significantly *less* volatile during FOMC events, suggesting some role for exchange rate stabilization policies around these episodes.

There is ample evidence that FX interventions are successful, at least in the short term, in reducing the volatility of exchange rates and that these policies are more prevalent in EMEs.[26] In Albagli et al. (2019), we provide further evidence about the higher degree of FX market interventions of countries in our EME sample relative to DEV using both de jure and de facto metrics of FX interventions (See table B4 of that paper). Other papers examining the volatility and composition of exchange rate reserves confirm these patterns. For example, Domanski et al. (2016) document that, after the Global Financial Crisis, the volatility of FX positions –controlling for valuation effects– has been exceptionally high in emerging markets across Latin America and Asia.

A complementary explanation rests on the higher volatility of idiosyncratic factors not identified in the event study, which in terms of our model would correspond to the noise introduced by the stochastic demand for local bonds u_t. As shown in Proposition 1, part a.iii), higher volatility of the idiosyncratic shock leads to a lower coefficient parameter around FOMC meetings as long as the loading of the risk-taking factor associated with US MP shock is not too high. The evidence reported in Panel A of Table 2, showing that the volatility of interest rate differentials in EMEs is more than twice the level exhibited by DEV –both unconditionally and around FOMC events –, suggests that such an explanation might also be contributing to lowering the significance of the relationship.

We report country-level evidence for Eq. (12) around US MP shocks in Table D.1 in Appendix D. Although there is significant heterogeneity within country categories, results are generally consistent with those from the pooled regression: We document a positive, significant correlation between exchange rate changes and long-term interest rate differentials around FOMC events for DEV, but we do not find a single country in the EME sample for which this condition holds, with most coefficients non-significant, and only a few EME exhibit a negative and statistically coefficient. When studying the underlying relationships between both exchange rates and yield differentials with US MP shocks, we find a weaker but significant negative sign associated with the exchange rate –currencies depreciate in response to tighter US MP, with the notable exception of Mexico, where the effect is non-significant. Also, we find a negative and statistically significant correlation between US MP shocks and interest rate differentials. The latter implies that when US MP tightens, US long-term yields increase more than domestic yields. Hence, it does not seem that the violation of the UIP relationship in EMEs comes from a change in sign of any of its underlying components in response to US MP shocks. Instead, the weak correlation might reflect a combination of a weaker exchange rate response due to central bank policy actions, or higher idiosyncratic volatility, as

[26] See Sarno and Taylor (2001), Ghosh et al. (2016) and Fratzscher et al. (2019).

highlighted by the model.

For robustness purposes, we run the specifications mentioned above, distinguishing between FOMC meetings in which only conventional monetary policy decisions were taken from meetings with unconventional balance sheet or forward guidance policies.[27] This distinction is useful in order to discriminate whether the results are driven by extraordinary monetary policy announcements in the US that are specific to this period. Table 4 reports the results. Overall, all the results are very similar in both subsamples. However, a few features are worthy of attention. Around FOMC meetings with unconventional announcements, the response of exchange rates is much larger than otherwise. Interestingly, while the reaction of US long-term yields to US MP shocks (Panel D) is also larger than around conventional meetings, yield differences – albeit larger – are not too different around both types of events. In our model, we have not allowed for a direct risk taking effect of US monetary policy shocks. Such a mechanism, emphasized in the literature by Bekaert et al. (2013), Bruno and Shin (2015), Kalemli-Özcan (2019) and Miranda-Agrippino and Rey (2020b), would not modify the effects of US monetary policy shocks in our model *unless* the risk taking effect is too large. In that case, the exchange rate responses to US monetary shocks are magnified, but the large response of domestic longer–term yields dampen the reaction of interest rate differentials. These results suggest that this mechanism might be present around unconventional monetary policy events.

Finally, Table E.1 in appendix E reports results using an extended sample until November 2020, both for the full November 2008-November 2020 sample, and another one excluding March 2020 due to the concerns discussed in Section 3. The results, including the Covid crisis, are similar to those using the baseline sample. In particular, for DEV, we find that a 1% increase in the long-term interest rate differential coincides with a US dollar depreciation of 3.9% (divided by $h = 10$ gives the 0.39 coefficient in the table, statistically significant at 1%). For EMEs, we document a coefficient with the opposite sign (marginally significant at 10%), similar to our baseline results. We conclude that neither set of results depends upon excluding the Covid-19 crisis.

The weaker results for EMEs across the different specifications suggest that either FX interventions or idiosyncratic shocks play a significant role in weakening the relationship between exchange rates and yield differences around US MP events. To obtain more direct evidence on the role of FX polices, we evaluate whether the response of asset prices to FOMC shocks differ in periods of high and low FX interventions in EMEs. For that, we rely on a data set put together by Adler et al. (2021), who provide data on FX interventions at monthly frequency for a set of countries included in our sample for the period 2000-2020.[28]

[27] Dates for meetings with balance sheet policies taken from https://www.federalreserve.gov/monetarypolicy/timeline-balance-sheet-policies.htm. Dates for meetings with forward guidance policies taken from https://www.federalreserve.gov/monetarypolicy/timeline-forward-guidance-about-the-federal-funds-rate.htm.

[28] In particular, these authors define FXI as any transaction that comprise the following four elements: (i) Active transactions by the Central Bank (excluding valuation changes or interests from income on reserves); (ii) the Central Bank is the main entity conducting these transactions; (iii) Transactions affecting the Central Bank's foreign currency

Table 4: UIP around US MP shocks–Conventional vs Unconventional Monetary Policy Meetings

	All FOMC Meetings				Conventional Policies				Only Unconventional Policies			
	(1)				(2)				(3)			
	$\hat{b}$	SE	R^2	N	$\hat{b}$	SE	R^2	N	$\hat{b}$	SE	R^2	N
Panel A (UIP): $\Delta s_t/h = \hat{a} + \hat{b} \cdot \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_t$												
DEV	0.47***	(0.08)	0.01	34584	0.29***	(0.07)	0.00	34584	0.73***	(0.11)	0.01	34584
EME	-0.10*	(0.06)	0.00	34584	-0.14**	(0.06)	0.00	34584	-0.07	(0.07)	0.00	34584
Panel B (Exchange rates): $\Delta s_t/10 = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_t$												
DEV	-1.00***	(0.07)	0.01	34584	-0.69***	(0.06)	0.00	34584	-1.54***	(0.11)	0.01	34584
EME	-0.59***	(0.08)	0.00	34584	-0.45***	(0.06)	0.00	34584	-0.84***	(0.14)	0.00	34584
Panel C (10-year yield differentials): $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_t$												
DEV	-0.97***	(0.06)	0.02	34584	-0.94***	(0.06)	0.01	34584	-1.01***	(0.09)	0.01	34584
EME	-0.94***	(0.09)	0.00	34584	-0.84***	(0.09)	0.00	34584	-1.10***	(0.17)	0.00	34584
Panel D (US 10-year yield): $\Delta y_t^{*(h)} = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_t$												
US	1.41***	(0.17)	0.03	2882	1.28***	(0.18)	0.02	2882	1.62***	(0.31)	0.02	2882

Notes: This table presents regressions around FOMC events. s_t is the exchange rate (USD per unit of domestic currency), $y_t^{(h)}$ and $y_t^{*(h)}$ are the local and the US 10-year yields, respectively, and i_t^* is the 2-year US yield. In all specifications $\Delta x_t = x_{t+1} - x_{t-1}$, when in t there is a FOMC meeting. Each panel presents pooled regressions for the corresponding group of countries (DEV, EME, and the US), with standard errors clustered at the country level. ***, ** and * denote statistical significance at the 1, 5 and 10% levels, respectively. Column (1) presents results with all FOMC meetings in the period November 24, 2008-December 31, 2019 (total of 89 meetings). Column (2) includes FOMC meeting with conventional policies only, excluding meetings with either balance sheet or forward guidance policies (excludes 31 meetings). Column (3) considers only FOMC meetings that include either balance sheet or forward guidance policies (total of 31 meetings).

Because the definition of FXI denotes changes in the foreign currency position of central banks (as a share of GDP), which are noisy in spite of the efforts to clean for valuation effects, dividends, and interest rate income, we focus on *extreme* periods to evaluate whether our results and mechanisms differ during months of high-FXI interventions vs. other periods. To operationalize this definition, for each EME country, we define periods of high FXI as those months below/above $x\%$ of the distribution of monthly changes in international reserves as a fraction of GDP, for $x = \{1, 5\}$. For example, with $x = 5$, high-FXI months are defined as those in which the changes in the reserve/GDP ratio are within the bottom 5% or the top 95% of the distribution of months.[29]

Table 5 presents the results. The first column (Baseline) replicates panels A to C for EMEs of Table 3. The next two columns report similar regressions distinguishing between periods of low– and high–FXI, for each threshold. The results in panel A reveal that the reaction of asset

position, encompassing any exchange of foreign and domestic currency assets and; (iv) All operations encompassing the above criteria, irrespective of the stated intent.

[29] FXI can occur either in the spot and derivatives market. For space considerations, we do not report the results distinguishing between both types of interventions, but they are very similar to those in Table 5.

prices in EMEs to US monetary policy shocks does not vary across different intervention states. However, panels B and C show that the mechanisms discussed in the model are confirmed in the data, meaning that the response of the domestic currency to FOMC shocks is significantly lower in periods of high FX intervention. Likewise, the response of yield differentials differs between high- and low–intervention periods, so that when FXI is high, the domestic–to–US 10–yield difference hardly changes, revealing that FX interventions that mute the effect on exchange rates amplify the impact on domestic longer–term rates. In periods of low intervention, the FX response is more pronounced, and the reaction of domestic longer-yields is much smaller, therefore amplifying the reaction of $(y_t^{(h)} - y_t^{*(h)})$. These results support the idea that FX defensive policies do alter the transmission mechanism of US MP shocks to exchange rates and domestic long-term yields, as predicted by the model. The results in panel A however provide evidence that these policies are not sufficient to explain the weaker results for EMEs, meaning that idiosyncratic volatility must play a role in such a result. These results are very much in line with Kalemli-Özcan and Varela (2023).

TABLE 5: UIP around US MP shocks–FXI (sales and purchases)

	Baseline	1%	5%
Panel A (UIP): $\Delta s_t/h = \hat{a} + \hat{b} \cdot \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_t$			
Base	-0.10*		
Low FXI		-0.10	-0.13*
High FXI		-0.30	-0.03
Panel B (Exchange rates): $\Delta s_t/10 = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_t$			
Base	-0.59***		
Low FXI		-0.60***	-0.61***
High FXI		-0.22	-0.42
Panel C (10-year yield differentials): $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_t$			
Base	-0.93***		
Low FXI		-0.97***	-0.89***
High FXI		-0.02	-1.22*
Panel D: Fraction of FXI events in daily sample			
%		2.95	10.77

4.2 Domestic MP shocks

Panel A in Table 6 reports results of Eq. (12) around domestic MP events. As before, all variables are measured as changes between the market close at date $t+1$ and date $t-1$, where date t corresponds to the day of the event –in this case, an MPM where the domestic MP rate is set. Similar to the US MP events, we document that around MPM, domestic currencies in DEV appreciate relative to the USD as long-term yield differentials $(y_t - y_t^*)$ increase. On average, across countries, a 1% widening in long-term yield differentials is associated with a domestic currency appreciation of 2.7%, although the significance is marginal, at 10% confidence. In contrast, for EMEs, we document no significant relationship between exchange rate changes and interest rate differentials in domestic MP events.

As above, we further study the response of exchange rates and long-term yield differentials separately around domestic MPM in panels B and C. For DEV, we document a domestic currency appreciation of 4.0% in response to a 1% increase in domestic MP shocks (weakly significant at 10% confidence). For EME, we find a response with the opposite sign, so that a 1% contractionary domestic MP shock *depreciates* domestic currencies in about 1.5%, which is highly statistically significant at 1%. Panel C of table 6 reports the impact on long-term yield differentials of domestic MP shocks. For both groups of countries, a contractionary MP shock of 1% leads to an increase in long-term rates that widens the yield differential in favor of domestic bonds of about 0.4% and 0.9% for DEV and EME respectively, both highly significant at 1% confidence. As expected, Panel D shows that domestic MP shocks have a negligible effect on 10-year US yields in the case of EME and are only marginally significant in the case of DEV. The implication is that the reaction of long-term domestic rates drives the bulk of the action in yield differentials. Moreover, these results confirm that there are spillovers from US monetary policy to domestic longer–term yields (as implicit in panels C and D of table 3), but not vice versa.

TABLE 6: Around domestic MPM events

	$\hat{b}$	SE	R^2	N
Panel A (UIP): $\Delta s_t/h = \hat{a} + \hat{b} \cdot \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_t$				
DEV	0.27*	(0.12)	0.00	28820
EME	-0.10	(0.05)	0.00	34584
Panel B (Exchange rates): $\Delta s_t/10 = \hat{a} + \hat{b} \cdot \Delta i_t + \nu_t$				
DEV	0.40*	(0.20)	0.00	28820
EME	-0.15***	(0.05)	0.00	34584
Panel C (10-year yield differentials): $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b} \cdot \Delta i_t + \nu_t$				
DEV	0.39***	(0.07)	0.01	28820
EME	0.90***	(0.15)	0.01	34584
Panel D (US 10-year yield): $\Delta y_t^{*(h)} = \hat{a} + \hat{b} \cdot \Delta i_t + \nu_t$				
DEV	0.20*	(0.11)	0.00	28820
EME	0.02	(0.01)	0.00	34584

NOTES: This table presents regressions around MPM events. s_t is the exchange rate (USD per unit of domestic currency), $y_t^{(h)}$ and $y_t^{*(h)}$ are the local and the US 10-year yields, respectively, and i_t^* is the 2-year US yield. In all specifications $\Delta x_t = x_{t+1} - x_{t-1}$, when in t there is a MPM meeting. Each panel presents pooled regressions for the corresponding group of countries (DEV – excluding Japan and Switzerland –, EME, and the US), with standard errors clustered at the country level. ***, ** and * denote statistical significance at the 1, 5 and 10% levels, respectively.

Finally, the results using the extended sample (reported in Table E.2 of appendix E) are, overall, consistent with the baseline sample. It is still the case that exchange rate and yield differences behave in line with a UIP-type relationship for DEV, although the size of the coefficient is smaller relative to our baseline results. In contrast, for EME, the relationship between long-term yield differentials and exchange rates is slightly negative. The largest difference is that the exchange rate response to domestic monetary policy shocks is less significant in the full sample for both country groups. However, all results of the baseline sample hold if we exclude March 2020.

These findings are informative of the nature of the failure of the UIP relationship in EMEs, where the reaction of the exchange rate to domestic MP policy shocks is inconsistent with a UIP-type relationship. In particular, domestic currencies depreciate relative to the USD when there is a contractionary domestic MP shock. While perhaps counterintuitive, a similar finding has been documented by Kohlscheen (2013) and Hnatkovska et al. (2016). In our model, such an effect can result for two reasons: a large enough risk-taking channel of domestic MP shocks ($\lambda > 1/b$) or high volatility of the idiosyncratic shock u_t that cannot be controlled for at a daily frequency.

Several mechanisms have been discussed in the literature to rationalize this evidence. For example, Kalemli-Özcan (2019) argues that risk premium considerations are key drivers of the monetary policy decision-making process in developing countries. In particular, the sensitivity of capital flows to risk innovations is high EMEs, meaning that risk-off events generate important capital outflows and weaken domestic currencies. In such a context, central banks try to defend their currencies by increasing their short-term interest rate, leading to a positive correlation between domestic interest rates and the exchange rate. Moreover, Kalemli-Özcan (2019) also concludes that large domestic monetary policy actions designed to limit exchange rate volatility can be counterproductive by increasing credit costs, spreads, and country risk premiums. Hnatkovska et al. (2016) explore a somewhat similar mechanism. According to their paper, if the domestic interest rate hike is a genuine response to a domestic inflationary environment, we would expect an appreciation of the domestic currency, unless the output or fiscal consequences of such a monetary policy tightening are strong enough when the domestic currency ends up depreciating relative to the USD even when domestic yields are increasing.

The identification of a specific mechanism through which contractionary domestic monetary policy shocks can depreciate the domestic currency is quite challenging in our event-study setting, as there is no daily data to control for the rationale behind the monetary policy reactions by central banks. To deal with this issue, we evaluate the reaction of exchange rates to domestic monetary policy shocks, controlling for the state of the economy. In particular, we distinguish periods of high domestic inflation (in absolute levels and relative to the US) and periods of high capital outflows from each country, and run the following empirical specification:

$$\frac{\Delta s_t}{10} = \hat{a} + \hat{b} \cdot \Delta i_t + \hat{c} \cdot \text{Event}_{j\tau} + \hat{d} \cdot \text{Event}_{j\tau} \cdot \Delta i_t + \nu_t, \tag{14}$$

where $\text{Event}_{j\tau}$ is a dummy variable taking the value of 1 in month τ in country j in either of the following three scenarios: (i) high domestic inflation, (ii) high domestic inflation differential with respect to the US or (iii) high capital outflows from the domestic economy. The value of $\text{Event}_{j\tau}$ is zero otherwise. Monthly inflation data comes from Ha et al. (2021), while capital flows come from Emerging Portfolio Fund Research (EPFR) dataset. We focus our attention in coefficients $\hat{b}$ and $(\hat{b} + \hat{d})$ in Eq. (14). The former captures the impact of a contractionary domestic MP shock on the exchange rate during "Normal times" (Event = 0), while the latter captures the total effect of a domestic MP shock in periods of "Distress" (Event = 1). The intuition for distinguishing between "Normal times" and "Distressed times" is that, presumably, in periods of high domestic inflation or high capital outflows, the domestic central bank might require a more contractionary policy to attain its domestic inflation and FX objectives, as in Kalemli-Özcan (2019).

Columns 1 and 2 in Table 7 present the results, where each panel has a different definition of Event. In Panel A, the Event takes the value of 1 in months in which the yearly inflation rate of the domestic economy is in the top 25% of the country-months observations. In panel B, Event = 1 if the 12-month inflation differential between the domestic economy and the US is also in the top 25% of the country pair-month observations. Finally, in panel C, the period of distress (Event = 1) is defined as a month in which capital outflows from the domestic economy (as a share of GDP) are in the top 25% of the distribution across months. The negative and significant coefficient $(\hat{b} + \hat{d})$ in all specifications reveals that the downward pressure in the value of the domestic currency that a contractionary domestic monetary policy exerts is relevant only in periods of "Distress". In terms of our model, this can be rationalized as assuming that the risk-taking effect of domestic MP shocks λ becomes active in periods of macroeconomic stress, which could induce capital outflows.

Table 7: Around domestic MPM events–Controlling for the state of the economy

	(1)	(2)
	Normal	Distress
Panel A: Headline Inflation		
	-0.12	-0.22**
Panel B: Inflation Differential		
	-0.10	-0.25**
Panel C: Capital Outflows		
	-0.09	-0.34***

Notes: This table presents regressions on the response of the NER to domestic monetary policy shocks, separating between normal periods from periods of distress for headline inflation (panel A), inflation differentials with respect to the US (panel B) and capital outflows (panel C). Periods of distress are defined (country-by-country) as months in which headline inflation, inflation differentials, and capital outflows are at the top 25% of the distribution at the country level. Normal times are defined as the remaining periods.

5 Uncertainty Shocks

Table 8 reports the results of regression (12) around global uncertainty events. To maintain coherence with the definition of the event window, all variables are measured as changes between the market close at day $t+1$ and day $t-1$, where day t corresponds to the day of the event. In our baseline specification, such an event is triggered at day t whenever the VIX movement between of market closing of t vs. $t-1$ is larger than two standard deviations of daily changes throughout our sample.

Panel A shows that coefficient $\hat{b}|_v$ is negative for both DEV and EME. In particular, there is, on average, a dollar appreciation of 1.9% (significant at 1%) for each percentage point of compression in US long-term interest rates relative to those in DEV. Similarly, we document a dollar appreciation of 1.7% (significant at 10%) relative to EME currencies for each 1% broadening of domestic long-term yield differentials around VIX events. Although the point estimates are similar for DEV and EME, it is less significant for the latter group.

Table 8: Around VIX events

	$\hat{b}$	SE	R^2	N
Panel A (UIP): $\Delta s_t/h = \hat{a} + \hat{b} \cdot \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_t$				
DEV	-0.19***	(0.05)	0.00	28,820
EME	-0.17*	(0.10)	0.01	34,584
Panel B (Exchange rates): $\Delta s_t/10 = \hat{a} + \hat{b} \cdot \Delta \text{VIX}_t + \nu_t$				
DEV	-5.92***	(0.98)	0.02	28,820
EME	-6.29***	(1.00)	0.02	34,584
Panel C (10-year yield differentials): $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b} \cdot \Delta \text{VIX}_t + \nu_t$				
DEV	3.49***	(0.68)	0.01	28,820
EME	8.38***	(1.58)	0.01	34,584
Panel D (US 10-year yield): $\Delta y_t^{*(h)} = \hat{a} + \hat{b} \cdot \Delta \text{VIX}_t + \nu_t$				
US	-5.92***	(1.15)	0.03	2,882

Notes: This table presents regressions around MPM events. s_t is the exchange rate (USD per unit of domestic currency), $y_t^{(h)}$ and $y_t^{*(h)}$ are the local and the US 10-year yields, respectively, and i_t^* is the 2-year US yield. In all specifications $\Delta x_t = x_{t+1} - x_{t-1}$, when in t there is an uncertainty event (days in which the daily change in the VIX is above or below two standard deviations). Panels B through D use the VIX normalized by its standard deviation during the corresponding events, which for the period 2008-2019 reaches 647 bp. Each panel presents pooled regressions for the corresponding group of countries (DEV –excluding Japan and Switzerland–, EME, and the US). Standard errors clustered at the country level. ***, ** and * denote statistical significance at the 1, 5 and 10% levels, respectively.

Panels B and C of Table 8 report the response of exchange rates and long-term yield differentials to changes in the VIX. Given the large jumps in the VIX around these events –the standard deviation of the 2-day VIX change between day $t+1$ and $t-1$ is 647 bps– we normalize the change in VIX by said standard deviation to facilitate the interpretation of the regression coefficients.[30] Regarding exchange rates, a one-standard deviation increase in the VIX leads to domestic currency depreciation of 0.59% and 0.63% for DEV and EME (significant at 1%), respectively. Regarding yields, a one standard deviation increase in VIX around the event window compresses the US 10-year yield close to 6 bps, leading to a positive long-term yield differential of 3.5 bps in DEV and 8.4 in EME (both significant at 1%). The implication is that DEV yields also compress around episodes of large VIX increases, albeit less than US yields, while EME long-term rates rise, on average, which is consistent with the findings in Kalemli-Özcan and Varela (2023) that the higher co-movement of global risk aversion and the UIP premium is EMEs is explained by the correlation between the VIX and interest rate differentials. Together with a quantitatively similar response in exchange rates, the larger response in yield differentials in EME is consistent with the weaker negative coefficient of the UIP relationship in panel A of the table. Also, these results are consistent with Bhattarai et al. (2020) who show how US uncertainty shocks lead to

[30] We exclude both Japan and Switzerland from the sample of DEV. As discussed in more detail below, there is a significant appreciation of the Swiss Franc and the Japanese Yen relative to the US dollar in risk-off events. See Appendix Table D.3.

currency depreciation and capital outflows in EMEs. Maybe more interestingly, they also find that the reversal in capital flows is larger in those countries that pay less attention to smoothing capital flows, i.e., have more active FX policies to mitigate the effect of uncertainty shocks on exchange rates.

Country-level results are reported in Appendix D. Similar to our baseline results, we find that $\hat{b}|_v$ has a negative sign for most countries around uncertainty events. Panel A in Table D.3 reports the results for the DEV sample. US long-term yields compress significantly more than that of most countries around days of large increases in the VIX, yet at the same time, the dollar exhibits a strong appreciation against most DEV currencies. Notable exceptions are Japan and Switzerland, which exhibit the opposite effect with a domestic currency appreciation of 0.53% and 0.37%, respectively. These findings are consistent with the evidence that both countries serve as safe-haven currencies during risk-off events (Ranaldo and Soderlind, 2010; Habib and Stracca, 2012). Panel B of Table D.3 presents the results for individual countries in the EME sample. Consistent with Table 8, the key feature is the large widening of yield differentials around uncertainty events.

Because the precise definition and measurement of uncertainty events are somewhat arbitrary, we perform a series of robustness checks in appendix D. A first set of robustness checks is reported in Table E.3, where we extend the sample period until November 2020. We present three alternative specifications: (i) we define uncertainty events as days with a large change in the VIX that do not fully reverse on the following day, (ii) we exclude March 2020 – which is the month of highest VIX volatility in the sample, and (iii) we consider only days of increases (rather than changes) in VIX beyond the two standard deviation threshold. Overall, the results are similar to those in the baseline sample in all specifications, except for the November 2008-November 2020 sample. As explained in Section 3, the large swings in VIX, frequent overlap between events of different types, and the reversal of long-held relationships between financial variables during March 2020 lead us to discard this specification as the preferred one.[31]

Also, we follow Bekaert et al. (2022) to extract from the VIX component the genuine uncertainty shock, in contrast to shocks to risk aversion. The VIX index combines a measure of the price of risk (risk aversion) and a conditional variance term (the quantity of risk), which is the relevant concept in our model. Table E.4 in appendix D reports the results. Overall, the results are very similar to those in the main text, regardless of the specification.

All in all, the evidence suggests a relevant role to risk-premia around episodes of strong VIX increases, as captured by the parameter κ in (4). The reason is that such events generally lead to an increase in domestic long-term yield differentials against US Treasuries while strengthening the USD

[31] We have also tried alternative thresholds for defining the VIX shock. In particular, instead of an uncertainty shock being defined out of a two standard deviation variation in the VIX, we also try one– and three–standard deviations. The results – not reported to conserve space, are very similar to the main specification, and they are available upon request.

relative to most currencies. Therefore, these results are consistent with the risk-premia hypothesis underlying most work addressing the failures of the UIP condition, but in our case, identified around specific events. The weaker results in EMEs suggest that, in line with Propositions 1.c.ii) and 1.c.iii), FX defensive policies and idiosyncratic noise may be playing some role in explaining this evidence, as the more muted response of exchange rates to uncertainty shocks is compensated with higher volatility in longer–term yields.

6 Evidence on currency excess returns

So far, our results shed light on whether exchange rates and long-term interest rate differentials respond to monetary policy and uncertainty shocks in the direction that an interest-rate parity condition suggests. An alternative approach to tackle this question is to analyze the behavior of excess returns of a strategy that goes short on USD-denominated bonds and long on domestic bonds around these events. In particular, following Mueller et al. (2017) we can evaluate the extent to which the ex-post excess return of such a strategy around monetary policy and risk events differs from non-event days. In contrast to Mueller et al. (2017), which evaluates one-day horizon strategies using a linear approximation of one-month risk-free interest rate differentials, we evaluate the returns of one- and two-day strategies using 10-year yield differentials. This approach allows us to compare the results with those in the previous section which are, by themselves, novel results in the literature.

We construct an investment strategy that goes short on US bonds and long on domestic bonds at day $t-1$, and that is held up today $t+k$, and compute ex-post currency excess returns using h-period bonds as:

$$rx_{t+k}^{(h)} = r_{t+k}^{(h)} - {r^*_{t+k}}^{(h)} + \Delta s_{t+k}, \tag{15}$$

where $r_{t+k} = p_{t+k}^{(h)} - p_{t-1}^{(h)}$ and ${r^*_{t+k}}^{(h)} = {p^*_{t+k}}^{(h)} - {p^*_{t-1}}^{(h)}$ denote the domestic and foreign k-period log-bond return for an instrument maturing h-periods ahead, and $\Delta s_{t+k} = s_{t+k} - s_{t-1}$ is the percentage variation in the exchange rate (dollars per domestic currency). This is the same approach taken by Mueller et al. (2017), but generalized to h-period bonds and for strategies lasting k days. Equation (15) allows us to compute daily excess returns of such strategy, and hence compare how it varies around monetary policy and risk events relative to unconditional days. After computing such excess returns, we run the following regression:

$$rx_{t+k}^{(h)} = \alpha_{t+k}^{(h)} + \beta_{t+k}^{(h)} \cdot \text{Event}_t + \epsilon_{t+k}^{(h)}, \tag{16}$$

where "Event$_t$" is a discrete variable that takes a value of 1 (-1) if there is an FOMC event on the day t and if the two-year yield change in the US is positive (negative), and 0 otherwise. Note that the variable "Event$_t$" considers the "direction" of the shock (i.e., if there are increases or decreases in the monetary policy rate). This fact is crucial for evaluating whether asset prices–exchange rates and interest rate differentials– amplify or dampen any pre-existing failure of the UIP condition.

We run this regression for the overall excess return in (15), as well as for each of its components (domestic and US interest rate returns as well as exchange rate returns). Panel A of Table 9 presents our results of two strategies using returns for ten-year bonds around FOMC events. Columns (1) and (2) present results for returns spanning the time windows $[t-1, t]$ and $[t-1, t+1]$, respectively, for DEV, meaning that in column (1) the strategy is activated in $t-1$ and closed in t (the day of the FOMC), while in column (2) we report the (cumulative) excess return of the same strategy when closed in $t+1$. The former time window allows us to compare the results with those in Mueller et al. (2017), while the latter allows us to contrast the evidence with our results in Section 4. Both columns show no systematic difference in excess returns around FOMC days relative to non-FOMC days.

The decomposition of ex-post excess returns for DEV reveals that a rise in US interest rates benefits such an investment strategy due to a relative increase in long-term US interest rates. However, the strengthening of the US dollar fully compensates for such an effect by lowering the return of a strategy that is long in domestic currencies. The results are entirely consistent with our findings in Section 4. Essentially, around FOMC meetings, there is no significant change in excess returns for DEV, suggesting that exchange rates and interest rate differentials neither amplify nor dampen any pre-existing deviation from the UIP condition.

TABLE 9: Currency excess returns in ten-year bonds around US MP shocks and VIX events

	DEV		EME	
	(1)	(2)	(3)	(4)
	$[t-1,t]$	$[t-1,t+1]$	$[t-1,t]$	$[t-1,t+1]$
Panel A: FOMC events				
Excess return	-1.16	-1.38	17.23**	11.01*
Interest rate	30.22***	32.65***	33.55***	35.83***
Domestic	-2.11	-18.44***	1.22	-15.25***
US	-32.33***	-51.09***	-32.33***	-51.09***
Exchange rate	-31.39***	-34.02***	-16.32**	-24.83***
Panel B: VIX events				
Excess return	-98.87***	-80.47***	-123.87***	-136.93***
Interest rate	-39.96***	-25.12***	-70.59***	-74.91***
Domestic	17.81**	27.65***	-12.81**	-22.14*
US	57.77***	52.77***	57.77***	52.77***
Exchange rate	-58.90***	-55.35***	-53.28***	-62.02***

NOTES: This table presents regressions for excess returns and their components around events as in Eq. (16). Numbers correspond to basis points. Excess returns and components are constructed as in Eq. (15) and bracketed intervals denote the period of each strategy. Panel A centers returns on FOMC events. Panel B center returns around VIX events. Each panel presents pooled regressions for the corresponding group of countries (DEV, EME), with standard errors clustered at the country level. ***, **, and * denote statistical significance at the 1, 5, and 10% levels, respectively.

The compensating effect of longer-term bond returns and currency responses around monetary policy meetings is related to recent work that has documented a downward-sloping term structure of carry trade risk premia. Lustig et al. (2019) find that excess returns of strategies involving long-term bonds are zero, driven by the compensatory effect between domestic bond returns and dollar appreciation, while excess returns of strategies involving short-term maturity bonds are significantly different from zero. In untabulated results,[32] we replicate the investment strategy presented in Table 9 focusing on short-term bonds (2-year yields) and document positive excess returns driven by USD appreciation, in line with Mueller et al. (2017).[33]

The evidence for EME is presented in columns (3) and (4) of Panel A in Table 9. There is some (weak) evidence that excess returns increase between 11 and 17 bps around FOMC events relative to non-FOMC days. This marginal increase in ex-post excess returns for EME comes from the more muted response of exchange rate returns to monetary policy shocks, which is indeed the evidence presented in Table 3 that shows that exchange rate responses in EMEs to US monetary policy shocks

[32] Available upon request.

[33] Mueller et al. (2017) analyze the excess return from a strategy that (i) goes from $t-1$ to t (hence $k =$ in our setup) and (ii) uses short-term interest rates (30 days). They find that, around the average FOMC meeting in 1994-2013, excess returns of a one-day long strategy are 10.28 basis points larger than in non-FOMC days. In our setting, we replicate our investment strategy focusing on short-term bonds (2-year bond yields) during the 1994-2013 and 2008-2019 periods, finding very similar results to theirs (10.63 and 15.60 bps in excess returns, respectively).

are weaker than in DEV.

Following our analysis in Section 5, we evaluate the response of excess returns of the strategy mentioned above around VIX events. For this, we re-define the variable "Event" in Eq. (16) to take a value of 1 (-1) if there is an event and if the VIX increases (decreases), and 0 otherwise. We present our results in Panel B of Table 9. For DEV and EME, we find significant decreases in excess returns in response to uncertainty shocks, which result from the fall in the US longer–term yield (relative to the domestic yield) as well as the domestic currency depreciation. Both adjustments lower the ex-post excess return of the strategy, consistent with the estimates in Section 5. The large effect on excess returns of a VIX shock in strategies exposed to EME currencies is due to a much larger increase in EME long-term interest rates, a phenomenon that is also documented in Section 5.

Table F.1 in the Appendix evaluates the economic significance of these results, reporting some unconditional descriptive statistics of excess returns for the different groups of countries and strategies. For example, for the two-day strategy covering $[t-1, t+1]$, the unconditional standard deviation of excess returns in DEV is 113 bps, while in EME is 213 bps. This means that, as reported in Panel B in Table 9, excess returns in EME around VIX events are 0.6 (=137/213) standard deviations larger than otherwise, and 0.7 times (=80/113) in DEV.

7 Conclusion

Traditional empirical strategies lead to a stark rejection of the UIP condition, a feature most commonly attributed to movements in risk-premia. This paper provides an event-study identification strategy, novel in this context, to evaluate how long-term interest rate differentials and exchange rates respond to monetary and uncertainty shocks.

Consistent with a simple model of partial arbitrage between domestic and US bonds, we find that asset prices in DEV respond in line with the UIP relationship around days of US and domestic MP shocks: around these events, ex-post excess returns of strategies that are long in domestic bonds and short in US bonds neither magnify nor dampen any pre-existing deviation of the UIP, as changes in yield differentials are largely compensated by offsetting movements in exchange rates. In contrast, we find a negative relationship between yield differentials and exchange rates around uncertainty events: while US long-term interest rates fall relative to domestic yields, the USD appreciates. This response is consistent with the attribution of the UIP violation to risk-premia. We find weak evidence of any relationship between these variables for emerging economies, which results from the extensive use of defensive policies to stabilize exchange rates – which has been widely documented for this group of countries –, as well as a higher prevalence of idiosyncratic shocks in EMEs.

Three corollaries seem to derive from our results. First, it appears that the rejection of the UIP condition in standard tests, which do not condition on specific shocks, may reflect a higher incidence of events that affect risk-premia compared to those related to changes in MP stances. Second, finding a UIP-consistent correlation between exchange rates and interest rates differential around MP shocks for developed economies provides some validation to the core relationship on which most international macroeconomics models rely, which we believe is a valuable policy insight. Third, our results suggest that, while emerging countries may succeed in partially containing exchange rate volatility through defensive policy actions, they do so at the cost of enhanced longer-term interest rate volatility –perhaps a less appreciated trade-off. Taken together, these results suggest a more nuanced interpretation of the UIP: Beyond its documented unconditional failure, the uncovered interest rate parity condition can still provide a valuable framework for understanding movements in financial markets in response to specific events.

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A Model: equilibrium characterization and proofs

A.1 Equilibrium characterization

The key equilibrium element of the model is the expression for the exchange rate, (8). To solve for the coefficients, we begin writing $K_t = -\gamma S_t/(1-\phi)$ from expression (6), and then use expressions (3), (4), and (2) to obtain a relationship between the exogenous shocks (m_t^*, m_t, v_t and u_t), the contemporaneous exchange rate S_t, and its one and h-period ahead expectations, $E\left[s_{t+1}|\Omega_t\right]$, and $E\left[s_{t+h}|\Omega_t\right]$, respectively. We can then apply the conjecture (8) on the one and h-period ahead expectations of the exchange rate, or $E\left[s_{t+1}|\Omega_t\right] = -a_1\rho m_t^* + a_2\psi m_t$ and $E\left[s_{t+h}|\Omega_t\right] = -a_1\rho^h m_t^* + a_2\psi^h m_t$, which leads to

$$s_t = -m_t^*\left[\frac{(1-\phi)\left(1+b/\beta^*+a_1(\rho+\rho^h b/h)\right)}{\gamma+(1-\phi)(1+b/h)}\right] + m_t\left[\frac{(1-\phi)\left(1-b\lambda+a_2(\psi+\psi^h b/h)\right)}{\gamma+(1-\phi)(1+b/h)}\right]$$
$$-v_t\left[\frac{(1-\phi)\left(b\kappa-b/\beta^*\right)}{\gamma+(1-\phi)(1+b/h)}\right] + u_t\left[\frac{(1-\phi)b}{\gamma+(1-\phi)(1+b/h)}\right].$$

We then apply the method of undetermined coefficients to solve for the equilibrium values of a_1 through a_4, which yields:

$$\begin{aligned} a_1 &= \frac{(1-\phi)\left(1+b/\beta^*)\right)}{\gamma+(1-\phi)\left(1-\rho+(1-\rho^h)b/h\right)},\ a_2 = \frac{(1-\phi)\left(1-\lambda b\right)}{\gamma+(1-\phi)\left(1-\psi+(1-\psi^h)b/h\right)}, \\ a_3 &= \frac{(1-\phi)b\left(\kappa-1/\beta^*\right)}{\gamma+(1-\phi)\left(1+b/h\right)}, \qquad \text{and} \qquad a_4 = \frac{(1-\phi)b}{\gamma+(1-\phi)\left(1+b/h\right)}, \end{aligned} \tag{A.1}$$

where $b \equiv \beta/(1+\beta) < 1$.

Using Eqs. (8), (9) and (13) we obtain the $\hat{b}|_x$ coefficients:

$$\begin{aligned} \hat{b}|_{m^*} &= \frac{\sigma_{m^*}^2\left(\beta/\beta^*\frac{a_1}{h}(1-\rho^h)\right)\frac{a_1}{h} - \sigma_u^2\left(1-\frac{a_4}{h}\right)\frac{a_4}{h}}{(1-b)\left[\sigma_{m^*}^2\left(\beta/\beta^*+\frac{a_1}{h}(1-\rho^h)\right)^2+\sigma_u^2\left(1-\frac{a_4}{h}\right)^2\right]}, \\ \hat{b}|_m &= \frac{\sigma_m^2\left(\frac{a_2(1-\psi^h)}{h}+\lambda\right)\frac{a_2}{h} - \sigma_u^2\left(1-\frac{a_4}{h}\right)\frac{a_4}{h}}{(1-b)\left[\sigma_m^2\left(\frac{a_2(1-\psi^h)}{h}+\lambda\right)^2+\sigma_u^2\left(1-\frac{a_4}{h}\right)^2\right]}, \\ \hat{b}|_v &= \frac{-\sigma_v^2\left(\beta/\beta^*+\kappa-\frac{a_3}{h}\right)\frac{a_3}{h} - \sigma_u^2\left(1-\frac{a_4}{h}\right)\frac{a_4}{h}}{(1-b)\left[\sigma_v^2\left(\beta/\beta^*+\kappa-\frac{a_3}{h}\right)^2+\sigma_u^2\left(1-\frac{a_4}{h}\right)^2\right]}. \end{aligned} \tag{A.2}$$

A.2 Proof of Proposition 1

a.i) US MP shocks: sign

From expression (A.2),

$$\hat{b}|_{m^*} \propto \sigma^2_{m^*} \left(\beta/\beta^* - \delta + \frac{a_1(1-\rho^h)}{h} \right) \frac{a_1}{h} - \sigma^2_u \left(1 - \frac{a_4}{h} \right) \frac{a_4}{h}$$

Taking the limit as $1/h^2 \to 0$, we get $\hat{b}|_{m^*} > 0$ iff $\frac{a_1}{a_4} \left(\beta/\beta^* - \delta \right) \sigma^2_{m^*} > \sigma^2_u$. Computing the ratio a_1/a_4 from (A.1) gives the condition in the proposition, where $\tilde{R}_1 \equiv \frac{\gamma+(1-\phi)(1+b/h)}{\gamma+(1-\phi)(1-\rho+b(1-\rho^h)/h)} \geq 1$.

a.ii) US MP shocks: effect of policy

We begin by noting that the derivative of the denominator of $\hat{b}|_{m^*}$ in expression (A.2) with respect to both policies is proportional to $(1-b)/h$, and can thus be ignored in the proposed limit. It follows that

$$\frac{\partial \hat{b}|_{m^*}}{\partial(1-\phi)} \propto \sigma^2_{m^*} \left(\beta/\beta^* - \delta \right) \frac{\partial a_1}{\partial(1-\phi)} - \sigma^2_u \frac{\partial a_4}{\partial(1-\phi)}.$$

Taking the ratio of the coefficient derivatives, we get that $\partial \hat{b}|_{m^*}/\partial(1-\phi) > 0$ (and thus $\partial \hat{b}|_{m^*}/\partial \phi < 0$) iff the condition of the proposition holds.

a.iii) US MP shocks: effect of noise

Taking the derivative of $\hat{b}|_{m^*}$ in expression (A.2) with respect to the variance of the idiosyncratic shock, and taking the proposed limit, yields:

$$\frac{\partial \hat{b}|_{m^*}}{\partial \sigma^2_u} \propto -\left(\beta/\beta^* - \delta \right) \left(\beta/\beta^* + \delta(\tilde{R}_1 - 1) + (1/b + 1/\beta^*)\tilde{R}_1 \right),$$

which is negative iff $\beta/\beta^* > \delta$ since $\tilde{R}_1 \geq 1$.

b.i) Domestic MP shocks: sign

From expression (A.2),

$$\hat{b}|_m \propto \sigma^2_m \frac{a_2}{h} \left(\frac{a_2(1-\psi^h)}{h} + \lambda \right) - \sigma^2_u \frac{a_4}{h} \left(1 - \frac{a_4}{h} \right).$$

Taking the limit as $1/h^2 \to 0$, and computing the ratio a_2/a_4 from (A.1) gives the condition in the proposition, where $\tilde{R}_2 \equiv \frac{\gamma+(1-\phi)(1+b/h)}{\gamma+(1-\phi)(1-\psi+b(1-\psi^h)/h)} \geq 1$.

b.ii) Domestic MP shocks: effect of policy

We begin by noting that the derivative of the denominator of $\hat{b}|_m$ in expression (A.2) with respect to both policies is proportional to $(1-b)/h$, and can thus be ignored in the proposed limit. It follows that

$$\frac{\partial \hat{b}|_m}{\partial(1-\phi)} \propto \sigma^2_m \lambda \frac{\partial a_2}{\partial(1-\phi)} - \sigma^2_u \frac{\partial a_4}{\partial(1-\phi)}.$$

Taking the ratio of the coefficient derivatives, we get that $\partial\hat{b}|_m/\partial(1-\phi) > 0$ (and thus $\partial\hat{b}|_m/\partial\phi < 0$) iff the condition of the proposition holds.

b.iii) Domestic MP shocks: effect of noise

Taking the derivative of $\hat{b}|_m$ in expression (A.2) with respect to the variance of the idiosyncratic shock, and taking the proposed limit, yields:

$$\frac{\partial\hat{b}|_m}{\partial\sigma_u^2} \propto \lambda(\lambda(\tilde{R}_2 - 1) - 1/b\tilde{R}_2),$$

which gives the stated condition.

c.i) Uncertainty shocks: sign

From expression (A.2),

$$\hat{b}|_v \propto -\sigma_v^2 \frac{a_3}{h}\left(\beta/\beta^* + \kappa - \frac{a_3}{h}\right) - \sigma_u^2 \frac{a_4}{h}\left(1 - \frac{a_4}{h}\right).$$

Taking the limit as $1/h^2 \to 0$, and computing the ratio a_3/a_4 from (A.1) gives the condition in the proposition.

c.ii) Uncertainty shocks: effect of policy

We begin by noting that the derivative of the denominator of $\hat{b}|_v$ in expression (A.2) with respect to both policies is proportional to $(1-b)/h$, and can thus be ignored in the proposed limit. It follows that

$$\frac{\partial\hat{b}|_v}{\partial(1-\phi)} \propto -\sigma_v^2(\kappa - 1/\beta^*)\frac{\partial a_3}{\partial(1-\phi)} - \sigma_u^2 \frac{\partial a_4}{\partial(1-\phi)}.$$

Taking the ratio of the coefficient derivatives, we get that $\partial\hat{b}|_v/\partial(1-\phi) > 0$ (and thus $\partial\hat{b}|_m/\partial\phi < 0$) iff $\left[\sigma_v^2(\kappa - 1/\beta^*)(\beta/\beta^* + \kappa) + \sigma_u^2\right] > 0$. That is, the sign of the derivative of $\hat{b}|_v$ with respect to $(1-\phi)$ is the exact opposite of the sign of $\hat{b}|_v$.

c.iii) Uncertainty shocks: effect of noise

Taking the derivative of $\hat{b}|_v$ in expression (A.2) with respect to the variance of the idiosyncratic shock, and taking the proposed limit, yields:

$$\frac{\partial\hat{b}|_v}{\partial\sigma_u^2} \propto -(\beta/\beta^* + \kappa)\left(\frac{1+\beta}{\beta^*}\right),$$

which gives the stated condition.

B Data coverage

Table B.1: Sample coverage

Code	Country	Classification	Number of MPM
AUD	Australia	DEV	122
CAD	Canada	DEV	89
CHK	Czech Republic	DEV	89
FR	France	DEV	113
GER	Germany	DEV	113
ITA	Italy	DEV	113
JPN	Japan	DEV	119
NOR	Norway	DEV	77
NZ	New Zealand	DEV	86
SW	Sweden	DEV	70
SZ	Switzerland	DEV	49
UK	United Kingdom	DEV	121
US	United States	DEV	89
CL	Chile	EME	123
COL	Colombia	EME	129
HUN	Hungary	EME	135
IND	India	EME	73
INDO	Indonesia	EME	128
ISR	Israel	EME	116
KOR	Korea	EME	121
MX	Mexico	EME	96
POL	Poland	EME	123
SG	Singapore	EME	23
SOA	South Africa	EME	70
THA	Thailand	EME	91

Notes: This table presents the sample coverage of the data. The last column reports the number of monetary policy meetings between November 24, 2008, and December 31, 2019. The number of VIX events in the sample is 139.

C Auxiliary evidence on expectations and exchange rate shocks

In this appendix we study the response of exchange rates forecasts at different horizons to changes in the spot exchange rate as well as to monetary policy shocks. Our first specification is the following:

$$\Delta E\left[s_{t,t+h}\right] = \alpha^h + \beta^h \Delta s_t + \epsilon_{t,t+h}. \tag{C.1}$$

Equation (C.1) corresponds to a regression on the monthly log-change between the expected exchange h periods ahead against the monthly log-change in the spot exchange rate, computed using the average daily exchange rate for each month. We have to run this specification with monthly data due to the lack of data on expectations at a higher frequency. We obtain information on expected exchange rates at different horizons (1, 3, and 12 months ahead, and for years $t+1$ to $t+6$, being t the current year) for November 2008-December 2019 from Consensus Forecast. This data reports market participants' expectations about the exchange rate evolution at horizons between 1 and 72 months ahead. It is available for seven developed economies: Canada, Japan, UK, Australia, Sweden, Switzerland, Czech Republic, and the Euro Zone, as well as for seven emerging economies: Mexico, Poland, Hungary, Chile, India, Korea, and Singapore. We run this regression country-by-country and also pooling data for each group of developed and emerging economies.

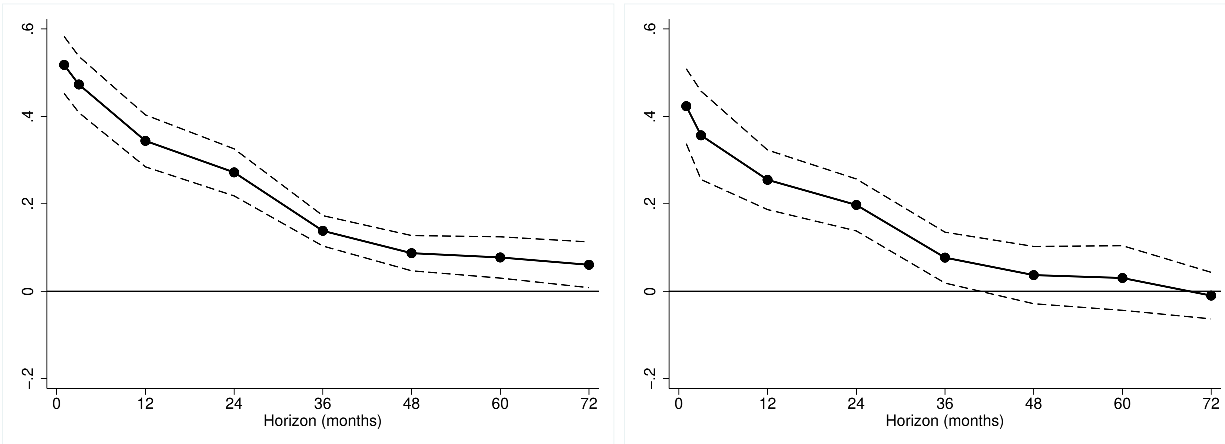
Consensus Forecast fixes the horizon of the expectation only in the case of 1, 3, and 12 months ahead. For $t+1$ through $t+6$, such horizon is mobile, depending on the month of the survey. For example, consider the expectations reported in January 2015. Since the expectation for the year $t+1$ corresponds to the average expected exchange rate between January 2016 and December 2016, this observation corresponds to an expected exchange rate between 12 and 24 months ahead. However, the same $t+1$ expectation observed in December 2015, which also refers to the period January 2016-December 2016, now corresponds to a shorter horizon of between 1 and 12 months ahead.

To address this issue, we interpolate expectations to fix the same horizon for every month in the sample by transforming the data on expectations for years $t+1$ to $t+6$. In particular, indexing every survey month by j, such that January takes value $j=1$ and December takes value $j=12$, we compute a weighted average between expectations for years $t+k$ and $t+k+1$, with $k=1,\ldots,5$, where the weights are $\omega_{t+k}=(12-j)/11$ and $\omega_{t+k+1}=(j-1)/11$, respectively. These weights allow us to extract the information on expectations for months to come, considering the moment of the year in which we obtain the data and the forecasting horizon of the survey respondents. Coming back to our example, note that in January 2005 the expectations for $t+1$ and $t+2$ will give us the average expectation between 12 and 24 months (because $\omega_{t+1}=1$ and $\omega_{t+2}=0$), which is the same information that we will obtain in December 2005 (because $\omega_{t+1}=0$ and $\omega_{t+2}=1$). For simplicity in the notation, we index these weighted average expectations as expectations $\{24,36,48,60,72\}$ months ahead.[34]

Figure C.1 plots the β^h coefficients for the pooled regression in each group of countries.[35]

[34] We find similar results using the raw data from Consensus Forecast.

[35] We find similar results when covering a more extended sample period starting from January 2003.



(a) DEV

(b) EME

Figure C.1: Exchange rate expectations and nominal exchange rate

Notes: This figure shows the β^h coefficient of the pooled OLS regression $\Delta E\left[s_{t,t+h}\right] = \alpha^h + \beta^h \Delta s_t + \epsilon_{t,t+h}$, where $E\left[s_{t,t+h}\right]$ is the log of expected exchange rate h months ahead, and s_t is the spot exchange rate, computed as a monthly average. The regression considers monthly data between November 2008 and December 2019. Dashed lines are the 95% confidence interval computed with standard errors clustered at the country level.

Figure C.1 is suggestive that, to first-order, short-term expectations overreact to (unconditional) exchange rate innovations, while this is not the case over longer horizons.[36] Ideally, we would like to evaluate how exchange rate expectations react to the shocks identified by our event studies. To delve deeper into this issue, we propose the following two-step specification:

$$\begin{aligned} \Delta s_t &= \alpha + \sum_i \beta_i^h \times \text{Shock}_{it-1} + \nu_t \\ \Delta E[s_{t,t+h}] &= \alpha^h + \sum_i \gamma_i^h \times \text{Shock}_{it-1} + \gamma_\nu^h \times \widehat{\nu}_t + \epsilon_{t,t+h}. \end{aligned} \tag{C.2}$$

In the first step, we jointly evaluate the impact of the US and domestic monetary policy shocks, as described in the text, on contemporaneous exchange rates. Because we are working with monthly data, the coefficients measure the effect on monthly exchange rate movements of US and domestic monetary policy shocks. The term ν_t captures all other determinants of monthly exchange rate movements, including uncertainty shocks. We do not include uncertainty shocks directly in the regression because their monthly equivalent is challenging to define, as these shocks tend to be short-lived. However, none of the results vary if we include the monthly change in the VIX as an independent variable. The second stage in Eq. C.2 evaluates the effect of monetary policy shocks, as well as other orthogonal determinants of FX movements, on FX expectations at different horizons, as in Eq. C.1.

[36]See Froot and Ito (1989) for a similar result.

The results, reported in Table C.1, show that both FOMC and domestic MP shocks have the expected impact on exchange rate expectations and that these effects vanish over time. In particular, as Table 3 shows, contractionary monetary policy shocks in the US systematically appreciate the US dollar vis–a–vis DEV and EME currencies. Therefore, the negative sign in the FOMC column in Table C.1 reveals that a contractionary monetary policy shock in the US also appreciates the expected value of the US dollar over the short term, and this effect disappears after two years. Likewise, domestic monetary policy shocks have transitory effects on exchange rate expectations, and the different sign for DEV and EME is coherent with the results in Table 6. Indeed, a contractionary monetary policy shock in developed economies appreciates their currencies vis–a–vis the US dollar, and appreciates the expected value of the exchange rate over time, as revealed by the positive coefficient in the MPM column. In contrast, contractionary MP shocks in EME depreciate their currencies, as shown in Table 6. The negative sign in Table C.1 shows that there is also a persistent (up to three to five years) depreciation of the expected exchange rate of EME currencies. Finally, the regressions also show that movements in the exchange rate orthogonal to FOMC and domestic MP shocks explain the bulk of monthly exchange rate movements, and these changes do have a high and significant effect on exchange rate expectations on the short run.

TABLE C.1: Exchange rate expectations and monetary policy shocks

Horizon	DEV				EME			
	Unconditional	Shocks			Unconditional	Shocks		
(months)		FOMC	MPM	NER ($\widehat{\nu}$)		FOMC	MPM	NER ($\widehat{\nu}$)
1	0.52***	-0.10***	0.04***	0.49***	0.42***	-0.09***	-0.02**	0.40***
3	0.47***	-0.09***	0.04**	0.45***	0.36***	-0.08**	-0.02**	0.34***
12	0.34***	-0.05**	0.02	0.33***	0.25***	-0.05*	-0.01**	0.25***
24	0.27***	-0.06*	0.01	0.26***	0.20***	-0.05*	-0.02*	0.18***
36	0.14***	-0.02	0.00	0.14**	0.08**	-0.03	-0.02**	0.06
48	0.09***	0.01	-0.01	0.10**	0.04	0.00	-0.02	0.04
60	0.08***	0.00	-0.01	0.08*	0.03	0.01	-0.02*	0.03
72	0.06**	0.01	-0.01	0.07	-0.01	0.03	-0.01	0.00

NOTES: This table presents results on the regression $\Delta E[s_{t,t+h}] = \alpha^h + \sum_i \gamma_i^h \times \text{Shock}_{it-1} + \gamma_\nu^h \times \widehat{\nu}_t + \epsilon_{t,t+h}$, where $\widehat{\nu}$ is the residual of the specification that regress Consensus Forecast exchange rate expectations on monetary policy shocks as follows: $\Delta s_t = \alpha + \sum_i \beta_i^h \times \text{Shock}_{it-1} + \nu_t$. Regressions use monthly data between November 2008 and December 2019. Unconditional refers to the results presented in Figure C.1. Shocks columns report the estimates for FOMC and MPM (domestic) monetary policy shocks, and the impact of changes in exchange rate orthogonal to shocks (NER ($\widehat{\nu}$)). Standard errors (not reported) are double clustered at the country and month level. ***, ** and * denotes statistical significance at the 1, 5 and 10% level, respectively.

Overall, these results show that monetary policy innovations significantly affect spot exchange rates and their expected values over short periods. Over long horizons, we do not find evidence that exchange rate expectations respond to monetary policy shocks or, more generally, to changes in spot exchange rates. Therefore, we assume $\Delta E\left[s_{t,t+h}\right] \to 0$ for h sufficiently high, i.e. 10 years, which leads to equation (12) in the text.

D Country-specific results

Table D.1: Around FOMC events: Country-by-country regressions

			Components of UIP			
	UIP (1)		Exchange rate (2)		10-year differential (3)	
Country	$\hat{b}$	SE	$\hat{b}$	SE	$\hat{b}$	SE
Panel A: DEV						
Canada	0.31	(0.29)	-0.51**	(0.21)	-0.52***	(0.11)
Japan	0.78***	(0.11)	-1.24***	(0.20)	-1.28***	(0.17)
United Kingdom	0.30	(0.24)	-0.58**	(0.23)	-0.71***	(0.22)
Germany	0.81***	(0.20)	-0.98***	(0.25)	-0.80***	(0.17)
Italy	0.18	(0.14)	-0.98***	(0.25)	-1.15***	(0.25)
France	0.67***	(0.18)	-0.98***	(0.25)	-0.92***	(0.20)
Australia	0.11	(0.16)	-0.95***	(0.25)	-0.92***	(0.22)
New Zealand	0.41	(0.26)	-1.27***	(0.27)	-1.00***	(0.20)
Norway	0.64**	(0.25)	-1.34***	(0.35)	-1.10***	(0.20)
Sweden	0.47*	(0.28)	-0.92***	(0.32)	-0.92***	(0.19)
Switzerland	0.84***	(0.29)	-1.25***	(0.26)	-1.02***	(0.18)
Czech Republic	0.40**	(0.16)	-0.93***	(0.22)	-1.30***	(0.23)
Pooled	0.47***	(0.08)	-1.00***	(0.07)	-0.97***	(0.06)
Panel B: EME						
Chile	-0.04	(0.11)	-0.54**	(0.24)	-1.19***	(0.28)
Colombia	-0.28**	(0.11)	-0.63**	(0.30)	-0.83**	(0.35)
Hungary	-0.05	(0.12)	-1.11***	(0.24)	-1.17	(0.74)
India	0.07	(0.11)	-0.43***	(0.14)	-1.39***	(0.24)
Indonesia	-0.18***	(0.04)	-0.23**	(0.11)	-0.41	(0.42)
Israel	0.00	(0.02)	-0.54***	(0.18)	-1.40***	(0.29)
Korea	0.13	(0.14)	-0.58***	(0.19)	-1.02***	(0.14)
Mexico	-0.62***	(0.16)	-0.17	(0.24)	-0.60**	(0.26)
Poland	-0.12	(0.25)	-0.92***	(0.23)	-1.04***	(0.24)
Singapore	0.06	(0.18)	-0.57***	(0.13)	-0.78***	(0.20)
South Africa	-0.56**	(0.24)	-0.91**	(0.38)	-0.90**	(0.37)
Thailand	-0.03	(0.07)	-0.44***	(0.09)	-0.48*	(0.26)
Pooled	-0.10*	(0.06)	-0.59***	(0.08)	-0.94***	(0.09)

Notes: This table presents country-by-country regressions around FOMC events. Panel A and B presents results for DEV and EME countries, respectively. Column (1) runs the regression $\frac{\Delta s_t}{10} = \hat{a} + \hat{b}\Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_{t,t+h}$, where s_t is the exchange rate (USD per unit of domestic currency), and $y_t^{(h)}$ and $y_t^{*(h)}$ are the domestic and US 10-year yield, respectively. Column (2) runs the regression $\frac{\Delta s_t}{10} = \hat{a} + \hat{b}\Delta i_t^* + \nu_t$, where i_t^* is the 2-year US yield. Finally, column (3) runs the regression $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b}\Delta i_t^* + \nu_{t,t+h}$. In all specifications $\Delta x_t = x_t - x_{t-2}$, when in $t-1$ there is a FOMC meeting. On each column, $\hat{b}$ denotes the coefficient and SE its standard error computed with Newey-West, except for pooled regressions with standard errors clustered at the country level. ***, ** and * denote statistical significance at the 1, 5 and 10% levels, respectively.

Table D.2: Around domestic MPM events: Country-by-country regressions

	UIP (1)		Components of UIP: Exchange rate (2)		Components of UIP: 10-year differential (3)	
Country	$\hat{b}$	SE	$\hat{b}$	SE	$\hat{b}$	SE
Panel A: DEV						
Canada	0.79***	(0.20)	0.71***	(0.09)	0.24**	(0.10)
Japan	0.79***	(0.13)	-0.96	(0.82)	-0.33	(0.64)
United Kingdom	0.36**	(0.16)	0.27	(0.26)	0.12	(0.16)
Germany	0.91***	(0.19)	0.64***	(0.20)	0.09	(0.10)
Italy	-0.07	(0.09)	-0.12	(0.08)	0.50***	(0.12)
France	0.48***	(0.18)	0.59***	(0.19)	0.10	(0.12)
Australia	0.21	(0.14)	0.71***	(0.08)	0.42***	(0.07)
New Zealand	0.29	(0.22)	1.16***	(0.18)	1.03***	(0.20)
Norway	0.48*	(0.26)	0.78***	(0.16)	0.59***	(0.10)
Sweden	0.55**	(0.22)	0.64***	(0.16)	0.29**	(0.15)
Switzerland	-0.53	(1.22)	-2.32	(1.78)	0.57***	(0.10)
Czech Republic	0.24	(0.18)	0.27	(0.19)	0.33***	(0.11)
Pooled	0.28**	(0.12)	0.33	(0.19)	0.40***	(0.07)
Pooled (ex JPN/SZ)	0.27*	(0.12)	0.40*	(0.20)	0.39***	(0.07)
Panel B: Emerging economies						
Chile	0.00	(0.09)	-0.02	(0.13)	0.54***	(0.20)
Colombia	-0.32***	(0.11)	-0.06	(0.13)	0.57***	(0.13)
Hungary	-0.29***	(0.07)	-0.31***	(0.09)	0.83***	(0.13)
India	0.02	(0.07)	0.00	(0.08)	0.61***	(0.11)
Indonesia	-0.20***	(0.02)	-0.22***	(0.07)	0.94***	(0.27)
Israel	-0.01	(0.01)	-0.02	(0.06)	2.48***	(0.70)
Korea	-0.11	(0.09)	-0.07	(0.09)	0.72***	(0.08)
Mexico	-0.23***	(0.08)	0.01	(0.12)	0.75***	(0.17)
Poland	-0.42***	(0.11)	-0.41**	(0.19)	0.94***	(0.14)
Singapore	-0.26	(0.26)	-0.68**	(0.33)	0.60	(0.46)
South Africa	-0.09	(0.20)	-0.24	(0.15)	0.64***	(0.12)
Thailand	-0.02	(0.05)	0.05	(0.04)	0.62***	(0.17)
Pooled	-0.10	(0.05)	-0.15***	(0.05)	0.90***	(0.15)

Notes: This table presents country-by-country regressions around domestic MPM events. Panel A and B present results for DEV and EME countries, respectively. Column (1) runs the regression $\frac{\Delta s_t}{10} = \hat{a} + \hat{b}\Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_{t,t+h}$, where s_t is the exchange rate (USD per unit of domestic currency), and $y_t^{(h)}$ and $y_t^{*(h)}$ are the domestic and US 10-year yield, respectively. Column (2) runs the regression $\frac{\Delta s_t}{10} = \hat{a} + \hat{b}\Delta i_t + \nu_t$, where i_t is the domestic 2-year yield. Finally, column (3) runs the regression $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b}\Delta i_t + \nu_{t,t+h}$. In all specifications $\Delta x_t = x_t - x_{t-2}$, when in $t-1$ there is a domestic meeting. On each column, $\hat{b}$ denotes the coefficient and SE its standard error computed with Newey-West, except for pooled regressions with standard errors clustered at the country level. ***, ** and * denote statistical significance at the 1, 5 and 10% levels, respectively.

Table D.3: Around VIX events: Country-by-country regressions

	Components of UIP					
	UIP (1)		Exchange rate (2)		10-year differential (3)	
Country	$\hat{b}$	SE	$\hat{b}$	SE	$\hat{b}$	SE
Panel A: DEV						
Canada	-0.47**	(0.23)	-8.36***	(1.32)	1.57***	(0.38)
Japan	0.53***	(0.17)	5.49***	(1.80)	5.23***	(1.01)
United Kingdom	0.11	(0.22)	-5.14**	(2.15)	2.55***	(0.50)
Germany	0.10	(0.18)	-2.37	(2.81)	2.42***	(0.57)
Italy	-0.16**	(0.08)	-2.37	(2.81)	8.35***	(1.96)
France	-0.09	(0.16)	-2.37	(2.81)	4.16***	(0.62)
Australia	-0.24	(0.15)	-10.77***	(2.01)	1.61	(1.12)
New Zealand	-0.24	(0.16)	-9.40***	(2.09)	3.76***	(1.07)
Norway	-0.28*	(0.15)	-7.80***	(2.17)	2.31***	(0.80)
Sweden	-0.05	(0.20)	-6.40**	(3.13)	2.45***	(0.70)
Switzerland	0.37	(0.27)	0.73	(2.33)	3.97***	(0.67)
Czech Republic	-0.40**	(0.16)	-4.21	(3.47)	5.74***	(0.89)
Pooled	-0.09	(0.08)	-4.41***	(1.33)	3.68***	(0.58)
Pooled (ex JPN/SZ)	-0.19***	(0.05)	-5.92***	(0.98)	3.49***	(0.68)
Panel B: EME						
Chile	-0.06	(0.11)	-6.39***	(1.36)	5.60***	(1.16)
Colombia	-0.53***	(0.12)	-8.97***	(1.63)	11.41***	(1.50)
Hungary	-0.32***	(0.08)	-7.06*	(3.76)	17.10***	(3.01)
India	-0.11*	(0.06)	-3.75***	(0.73)	6.38***	(0.93)
Indonesia	-0.17***	(0.02)	-3.09***	(0.66)	16.20***	(2.67)
Israel	-0.02***	(0.01)	-2.96**	(1.50)	-2.72	(6.73)
Korea	-0.39**	(0.16)	-6.78***	(1.24)	5.44***	(1.17)
Mexico	-0.70***	(0.15)	-10.60***	(1.70)	10.36***	(2.09)
Poland	-0.66***	(0.14)	-7.38**	(3.05)	9.30***	(1.48)
Singapore	-0.15	(0.10)	-4.38***	(1.06)	3.71***	(0.70)
South Africa	-0.79***	(0.08)	-13.03***	(2.82)	11.59***	(1.88)
Thailand	-0.03	(0.02)	-1.12**	(0.52)	6.19***	(1.28)
Pooled	-0.17*	(0.10)	-6.29***	(1.00)	8.38***	(1.58)

Notes: This table presents country-by-country regressions around risk-premium events. Panels A and B presents results for DEV and EME countries, respectively. Column (1) runs the regression $\frac{\Delta s_t}{10} = \hat{a} + \hat{b}\Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_{t,t+h}$, where s_t is the exchange rate (USD per unit of domestic currency), and $y_t^{(h)}$ and $y_t^{*(h)}$ are the domestic and US 10-year yield, respectively. Column (2) runs the regression $\frac{\Delta s_t}{10} = \hat{a} + \hat{b}\Delta\text{VIX}_t + \nu_t$. Finally, Column (3) runs the regression $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b}\Delta\text{VIX}_t + \nu_{t,t+h}$. In all specifications $\Delta x_t = x_t - x_{t-2}$, when in $t-1$ there is a risk-taking event. The specifications in columns (2) and (3) use the VIX normalized by its standard deviations during the corresponding events, which for the period 2008-2019 reaches 647 bp. On each column, $\hat{b}$ denotes the coefficient and SE its standard error computed with Newey-West, except for pooled regressions with standard errors clustered at the country level. ***, ** and * denote statistical significance at the 1, 5 and 10% levels, respectively.

E Robustness

This appendix reports the results of robustness checks for the regressions described in Tables 3, 6 and 8, along two main dimensions. First, we extend the sample period until November 2020 to include the Covid-19 period. Second, we present the results for uncertainty shocks using alternative specifications for how these events are defined.

E.1 Sample extension: November 2008–November 2020

Extending the sample period until November 2020 incorporates the Covid-19 crisis, which generated extraordinary volatility in asset prices, in particular during March. As a result, out of the ten days with the largest daily increases in the absolute value of the VIX index in the period November 2008-November 2020, two correspond to March 2020 (March 12th and 16th), and out of the ten days with the largest daily fall in the VIX index, four occurred in March 2020. Besides, the policy reactions of Central Banks with extraordinary meetings during those days, together with the high demand for USD liquidity in that period, pose a challenge to our identification strategy due to the large degree of overlap between events in such a short window. Therefore, we present two sets of regressions, one for the total sample November 2008-November 2020 and another one that includes an interaction with a dummy variable for March 2020 (we report the coefficient associated with the main regressor and thus, exclude the March 2020 effect).

TABLE E.1: UIP around US MP shocks (extended sample)

	Baseline		Ex March 2020	
	(1)		(2)	
	$\hat{b}$	SE	$\hat{b}$	SE
Panel A (UIP): $\Delta s_t/h = \hat{a} + \hat{b} \cdot \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_t$				
DEV	0.39***	(0.09)	0.43***	(0.08)
EME	-0.12*	(0.06)	-0.13*	(0.06)
Panel B (Exchange rates): $\Delta s_t/10 = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_t$				
DEV	-0.85***	(0.06)	-0.99***	(0.07)
EME	-0.52***	(0.07)	-0.58***	(0.08)
Panel C (10-year yield differentials): $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_t$				
DEV	-0.89***	(0.06)	-0.96***	(0.06)
EME	-0.76***	(0.11)	-0.93***	(0.10)
Panel D (US 10-year yield): $\Delta y_t^{*(h)} = \hat{a} + \hat{b} \cdot \Delta i_t^* + \nu_t$				
US	1.27***	(0.19)	1.41***	(0.17)

NOTES: This table presents regressions around FOMC events. s_t is the exchange rate (USD per unit of domestic currency), $y_t^{(h)}$ and $y_t^{*(h)}$ are the local and the US 10-year yields, respectively, and i_t^* is the 2-year US yield. In all specifications $\Delta x_t = x_{t+1} - x_{t-1}$, when in t there is a FOMC meeting. Column (1), "Baseline", presents the results for the extended sample. Column (2), "ex March 2020", extends the specification of column 1 by including a dummy variable taking value one for March 2020 (and zero otherwise) and the interaction between the dummy and the regressor of interest. We report the coefficient associated to the main regressor and thus, excludes March 2020 effect. Each panel presents pooled regressions for the corresponding group of countries (DEV, EME, and the US), with standard errors clustered at the country level. ***, ** and * denote statistical significance at the 1, 5 and 10% levels, respectively.

Table E.2: UIP around domestic MPM events (extended sample)

	Baseline		Ex March 2020	
	(1)		(2)	
	$\hat{b}$	SE	$\hat{b}$	SE
Panel A (UIP): $\Delta s_t/h = \hat{a} + \hat{b} \cdot \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_t$				
DEV	0.20*	(0.10)	0.25*	(0.12)
EME	-0.10*	(0.06)	-0.09*	(0.05)
Panel B (Exchange rates): $\Delta s_t/10 = \hat{a} + \hat{b} \cdot \Delta i_t + \nu_t$				
DEV	0.36	(0.20)	0.39*	(0.20)
EME	-0.11	(0.06)	-0.12**	(0.05)
Panel C (10-year yield differentials): $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b} \cdot \Delta i_t + \nu_t$				
DEV	0.39***	(0.08)	0.40***	(0.07)
EME	0.86***	(0.14)	0.88***	(0.15)
Panel D (US 10-year yield): $\Delta y_t^{*(h)} = \hat{a} + \hat{b} \cdot \Delta i_t + \nu_t$				
DEV	0.21*	(0.11)	0.20*	(0.11)
EME	0.03	(0.02)	0.03	(0.02)

Notes: This table presents regressions around MPM events. s_t is the exchange rate (USD per unit of domestic currency), $y_t^{(h)}$ and $y_t^{*(h)}$ are the local and the US 10-year yields, respectively, and i_t^* is the 2-year US yield. In all specifications $\Delta x_t = x_{t+1} - x_{t-1}$, when in t there is a MPM meeting. Column (1), "Baseline", presents the results for the extended sample. Column (2), "ex March 2020", extends the specification of column 1 by including a dummy variable taking value one for March 2020 (and zero otherwise) and the interaction between the dummy and the regressor of interest. We report the coefficient associated to the main regressor and thus, excludes March 2020 effect. Each panel presents pooled regressions for the corresponding group of countries (DEV – excluding Japan and Switzerland –, EME, and the US), with standard errors clustered at the country level. ***, ** and * denote statistical significance at the 1, 5 and 10% levels, respectively.

E.2 Robustness on uncertainty events

This appendix presents alternative definitions of uncertainty events. Under our standard definition, incorporating January 2020 through November 2020 increases the number of VIX events by 39, from 139 to 178. This situation is particularly relevant during March 2020, where three days with the largest daily increases in the VIX index were followed by three of the largest daily falls in the VIX, from November 2008 to November 2020. We present alternative results that incorporate only VIX shocks that do not reverse during the next day. That is, the alternative definition considers only events where we observe in day t a change in the VIX two or more standard deviations either above or below its mean, *and* a two–day change in the index (between the end of $t-1$ and the end of $t+1$) also larger than two standard deviations. Additionally, we present results restricting events to days of VIX increases only, and we also estimate the regressions excluding March 2020 events.

Table E.3: UIP around VIX events (extended sample)

	Baseline		No reversion		ex March 2020		Only up	
	(1)		(2)		(3)		(4)	
	$\hat{b}$	SE	$\hat{b}$	SE	$\hat{b}$	SE	$\hat{b}$	SE
Panel A (UIP): $\Delta s_t/h = \hat{a} + \hat{b} \cdot \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_t$								
DEV	-0.10*	(0.05)	-0.15**	(0.05)	-0.17***	(0.04)	-0.21**	(0.07)
EME	-0.18**	(0.08)	-0.20**	(0.08)	-0.18*	(0.09)	-0.19*	(0.10)
Panel B (Exchange rates): $\Delta s_t/10 = \hat{a} + \hat{b} \cdot \Delta \mathrm{VIX}_t + \nu_t$								
DEV	-5.04***	(0.93)	-5.06***	(0.90)	-5.45***	(0.97)	-5.70***	(1.10)
EME	-5.39***	(0.83)	-5.31***	(0.83)	-5.88***	(0.92)	-6.56***	(0.91)
Panel C (10-year yield differentials): $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b} \cdot \Delta \mathrm{VIX}_t + \nu_t$								
DEV	3.15***	(0.74)	3.14***	(0.73)	3.32***	(0.65)	2.94**	(0.93)
EME	7.04***	(1.38)	6.94***	(1.38)	7.73***	(1.36)	7.69***	(1.85)
Panel D (US 10-year yield): $\Delta y_t^{*(h)} = \hat{a} + \hat{b} \cdot \Delta \mathrm{VIX}_t + \nu_t$								
US	-4.13***	(1.30)	-4.01***	(1.28)	-5.64***	(0.99)	-4.83***	(1.45)

Notes: This table presents regressions around uncertainty events, where s_t is the exchange rate (USD per unit of domestic currency), $y_t^{(h)}$ and $y_t^{*(h)}$ are the domestic and US 10-year yield, respectively. In all specifications $\Delta x_t = x_{t+1} - x_{t-1}$, when in t there is an uncertainty event (days in which the daily change in the VIX is above or below two standard deviations). Column (1), "Baseline", presents the results for the extended sample using the baseline definition for uncertainty events (178 events). Column (2), "No Reversion", presents the results using a definition of events in which the daily change of the VIX is larger than two standard deviations *and* the two-days change is also larger than two standard deviations (130 events). Column (3), "ex March 2020", extends the specification of column 1 by including a dummy variable taking value one for March 2020 (and zero otherwise) and the interaction between the dummy and the regressor of interest. We report the coefficient associated to the main regressor and thus, excludes March 2020 effect. Finally, column (4), "Only up", presents results using a definition of events in which there are only increases in the VIX (102 events). In panels A–C, the regression for DEV excludes Japan and Switzerland. Panels B through D use the VIX normalized by its standard deviation during the corresponding events, which for the period 2008-2019 reaches 647 bp. Each panel presents pooled regressions for the corresponding group of countries, with standard errors clustered at the country level. ***, ** and * denote statistical significance at the 1, 5 and 10% levels, respectively.

Another robustness exercise is related to the nature of the uncertainty shock. We follow Bekaert et al. (2022) to decompose the VIX into its variance risk premium risk term and a conditional variance term. We follow two different approaches to define the uncertainty shock:

1. We assume that fluctuations in the VIX determine the event day (meaning that the event day does not change relative to the baseline specification), and we study the impact of either VIX or uncertainty shocks around these days.

2. We define uncertainty events as those days in which uncertainty – as defined by Bekaert et al. (2022) – changes significantly around a two-day window. Hence, this specification not only modifies the measure of uncertainty shock but also the days in which these shocks take place, relative to the baseline regressions.

Table E.4 presents the results. In the first two columns, the event day is defined according to the VIX shock, while in the last two columns, the event day is defined by spikes in the uncertainty component of the VIX. (The first column – VIX-VIX – reports the results of the main specification in the paper.) Panels A-D are directly comparable with those in Table 6 in the paper, while panel E presents the number of events identified in each case, and the overlap between uncertainty and VIX events (in percentage points), measured as the fraction of VIX events for which there is also an uncertainty event.[37] Overall, the results are very similar to those in the main text, regardless of the specification.

Table E.4: VIX and uncertainty events

Variable defining event	VIX		Uncertainty	
Shock	VIX	Uncertainty	VIX	Uncertainty
Panel A (UIP): $\Delta s_t/h = \hat{a} + \hat{b} \cdot \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_t$				
DEV	-0.19***	-0.19***	-0.19**	-0.19**
EME	-0.17*	-0.17*	-0.17*	-0.17*
Panel B (Exchange rates): $\Delta s_t/10 = \hat{a} + \hat{b} \cdot \Delta\text{VIX}_t + \nu_t$				
DEV	-5.92***	-5.97***	-6.58***	-4.88***
EME	-6.29***	-6.09***	-6.28***	-4.80***
Panel C (10-year yield differentials): $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b} \cdot \Delta\text{VIX}_t + \nu_t$				
DEV	3.49***	4.05***	4.36***	4.52***
EME	8.38***	8.69***	10.07***	8.74***
Panel D (US 10-year yield): $\Delta y_t^{*(h)} = \hat{a} + \hat{b} \cdot \Delta\text{VIX}_t + \nu_t$				
US	-5.92***	-6.67***	-7.41***	-7.59***
Panel E: Number of events and overlap (as fraction of VIX events)				
# events	139	139	106	106
Overlap (%)			52.17	

[37]The total number of events is quite similar using the VIX definition and the Bekaert et al. (2022) definition.

F Unconditional currency excess returns

Table F.1: Unconditional currency excess returns for long-term bonds

	DEV		EME	
	Mean	Std. Dev.	Mean	Std. Dev.
Panel A: 1-day holding period ($[t-1,t]$)				
Excess return	0.41	86.05	0.38	151.71
Interest rate	0.54	57.71	0.77	125.75
Domestic	0.98	44.99	1.20	115.35
US	0.43	52.60	0.43	52.60
Exchange rate	-0.14	67.37	-0.39	67.69
Panel B: 2-day holding period ($[t-1,t+1]$)				
Excess return	0.66	113.49	0.63	213.19
Interest rate	1.04	73.58	1.46	172.92
Domestic	1.96	64.23	2.39	161.13
US	0.92	72.28	0.92	72.28
Exchange rate	-0.37	94.30	-0.83	95.68

Notes: This table presents unconditional descriptive statistics for excess returns and their components. Excess returns and components are constructed as in Eq.(15). Panel A computes returns in the $[t-1,t]$ window. Panel B computes returns in the $[t-1,t+1]$ window. All variables in basis points.

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