UIP: Insights from Event Studies*

Elias Albagli† Luis Ceballos‡ Sebastian Claro§ Damian Romero¶

January 5, 2023

Abstract

We develop a model of partial arbitrage between foreign and domestic long-term bonds. The covariance between yield differentials and exchange rates is conditional on the nature of shocks, which we identify using event studies. In line with the UIP, tighter US monetary policy appreciates the dollar while increasing US yield differentials. In contrast, global uncertainty shocks appreciate the dollar but reduce differentials, exacerbating the widely documented UIP violation. Both relationships are weaker in emerging countries, consistent with more pervasive currency stabilization policies. Our results suggest that the UIP logic remains helpful in predicting market prices in response to monetary innovations.

JEL Codes: F30, F31, G12, G15.

Keywords: uncovered interest parity, long-term yields, event studies.

*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 would like to thank Borja Larrain, Hyun-Song Shin, Xiaodong Zhu, and seminar participants at PBC School of Finance, Tsinghua University, the International Monetary Fund, the Czech National Bank and Universidad de Los Andes, Chile.

†Central Bank of Chile. Email: ealbagli@bcentral.cl

‡Knauss School of Business, University of San Diego. Email: luisceballos@sandiego.edu

§Corresponding author. Facultad de Ciencias Economicas y Empresariales, Universidad de los Andes, Monseñor Alvaro del Portillo 12455, Santiago 7550000, Chile. Email: sclaro@uandes.cl

¶Central Bank of Chile. Email: dromeroc@bcentral.cl

Declaration of interest

The authors declare that they have no relevant or material financial interests that relate to the research described in this paper.

Elias Albagli
Central Bank of Chile

Luis Ceballos
University of San Diego

Sebastian Claro
Universidad de los Andes, Chile

Damian Romero
Central Bank of Chile

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] Its failure has been mainly attributed to innovations in some form of risk premia and policy actions in countries with fear of floating concerns.[2] For example, Chinn and Meredith (2004) and Chinn (2006) argue that exchange rate shocks, coupled with defensive monetary policy responses, break down the traditional UIP at short horizons. However, this result does not hold when tested in long-term bonds due to the transitory nature of these shocks and central banks' limited ability to influence yields at higher maturities. More recently, Lustig et al. (2014) document a more systematic relationship between interest rate differentials and risk premia. They build a model where the stochastic discount factor of US-based investors is governed by relative US cyclical conditions so that lower interest rates in the US coincide with periods of higher marginal utility and required return for holding foreign securities. In a similar spirit, di Giovanni et al. (2022) document that borrowing in local currency becomes cheaper than in foreign currency during expansionary phases of the global financial cycle, while Hofmann et al. (2019) argue that shifts in the credit risk premia which compress long-term yields in emerging countries tend to appreciate their currencies.

Notwithstanding the risk premium-based dominant explanation for the forward premium puzzle, there is evidence that financial variables respond in the direction suggested by the UIP to some shocks. For instance, foreign currencies tend to depreciate on impact in response to contractionary US monetary policy shocks, contradicting some of the risk-premium-driven explanations, which is indeed the essence of the Engel puzzle (Engel, 2016). In contrast, the literature consistently documents a US dollar appreciation and a fall in US interest rates in response to uncertainty shocks. These results tend to confirm the failure of the UIP based upon a risk-premium explanation.[3] The differential response of financial variables to shocks of different nature suggests a more nuanced interpretation of the UIP: beyond its documented failure, it can perhaps still provide a valuable framework for understanding movements in financial markets in response to specific events.

We develop a framework encompassing monetary policy and uncertainty shocks to shed light on these issues. In particular, we build a model in which domestic US investors partially arbitrage (non-US) long-term bond markets. The model delivers differential predictions for the covariance of exchange rates and interest rates differentials *conditional* on the source of shocks and links its magnitude to the degree of (monetary and FX) policy intervention. We then empirically validate the core predictions of the model using an identification strategy based on event studies that, arguably,

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

[2] Other explanations are related to deviations from the rational expectations hypothesis implicit in the way exchange rate expectations are formed as embedded in the UIP condition or specifics of the market microstructure, like transaction costs or taxes.

[3] See, for example, Caballero and Kamber (2019), and Carrière-Swallow and Cespedes (2013)

can help identify the source of shocks. The results show that monetary policy shocks neither amplify nor dampen any pre-existing deviation of the UIP –that is, movements in exchange rates and interest rate differentials in response to these shocks largely cancel each other out, in line with the no-arbitrage mechanism behind the UIP relationship. In contrast, large uncertainty shocks magnify the forward premium puzzle.

In our model, US investors under rational expectations participate in local long-term bond markets, whose demand exhibits positive but finite elasticity to deviations in the UIP condition. Short-term interest rates are governed by local monetary policy (MP) decisions, while a balance of payments condition sets the exchange rate. There are two dominant sources of innovations: MP (in the US and other countries) and uncertainty/risk-sentiment shocks. When US MP tightens, the rise in the US-yield curve across horizons induces capital outflows of US investors from foreign markets, depreciating local currencies against the US dollar (USD). At the same time, there is a partial narrowing in long-term interest rate differentials as local investors are required to buy the bonds dumped by non-residents, driving up local yields. Conditional on this shock, an econometrician would find that higher interest rate differentials in favor of the US are indeed compensated by a contemporaneous depreciation (and an ensuing appreciation) of domestic currencies. The implication is that ex-ante excess returns of a strategy long on the domestic currency and short on the US dollar do not change around US monetary policy events. However, the strength of this relationship is mitigated if central banks respond through either tighter monetary policy or foreign exchange interventions, as these actions dampen the currency movement while amplifying capital outflows and long-term yield differentials.[4] Naturally, the strength of the relationship could also weaken in the presence of idiosyncratic volatility unrelated to MP shocks, to the extent that such shocks cannot be controlled empirically.

In contrast, uncertainty shocks lead to a negative contemporaneous covariance between the USD and interest rate differentials (US minus local government bond yields), meaning that ex-ante excess returns of such strategies increase in response to a risk premium shock. For example, when a risk-off event leads international investors into the safety of US Treasuries, domestic yields increase relative to US yields even as the USD appreciates, as in Lustig et al. (2014) and Hofmann et al. (2019). This reaction of asset prices enhances the excess returns of strategies that go long in domestic bonds and short in US bonds. At such events, policy interventions also mitigate local exchange rate movements at the cost of increasing capital outflows and thus enlarging long-term interest rate differentials. Just as in the case of MP events, interventions thus weaken the absolute value of the (now negative) contemporaneous correlation between long-term interest rate differentials and US-local currencies exchange rate. Our model thus highlights both the importance of evaluating the covariance between yield differentials and exchange rate *conditioning* on the sources of the shocks, as well as the role played by the domestic policy in mitigating said relationship.

[4] See Albagli et al. (2019).

We test the model using an event study methodology based on changes in long-term interest rates and exchange rates within a two-day bracket around specific events between November 2008 and December 2019. We focus on the post-GFC period as it witnessed a strong trend towards capital market integration and foreign presence in emerging bond markets.[5] Our sample includes 24 countries split equally between developed and emerging economies (simply DEV and EME, henceforth). This approach has two key advantages and one limitation. First, by focusing on the reaction of interest rates and exchange rates around narrowly-defined windows bracketing particular events, we provide more reliable identification of the two key shocks highlighted by the model: MP events, centered around both US and local MP meetings, and uncertainty shocks, defined as days when the VIX exhibits large fluctuations. Second, by focusing on long horizons, we show that changes in exchange rate dynamics are well approximated by the instantaneous movement in the exchange rate, bypassing the need to proxy for exchange rate expectations at long horizons with actual future exchange rates, which is very taxing in terms of data. We provide evidence to validate this methodological assumption using Consensus Forecast data, showing that exchange rate fluctuations lead to a monotonically decaying exchange rate expectations "term structure" that approaches zero well before the 10-year maturity. A similar pattern emerges when analyzing the effect on exchange rate expectations of monetary policy innovations. After controlling for orthogonal variations in the exchange rate to these shocks, we find that monetary policy shocks affect FX expectations over the short term. However, this effect vanishes over long horizons.[6]

The main limitation of this event study approach is that it is not well suited for evaluating the covariance between exchange rates and interest rate differentials using short-term bonds. This is because even if the rational expectations hypothesis were valid,[7] one could not assume that the one-month expectation of the exchange rate at day $t-1$ – say, the day before the event – is well approximated by the $t+29$-day value of the exchange rate, since such value will undoubtedly be a closer reflection of the updated information set of investors after day $t+1$, after the shock. In other words, we would need to observe the change in exchange rate expectations at short maturities around specific events, which is not available. Moreover, the equivalent assumption of no-change in exchange rate expectations around the event bracket is inconsistent with our additional evidence about the persistence of exchange rate expectations to innovations in spot rates. Focusing on long-run interest rates and exchange rate differences across events is convenient for sidestepping this problem.

We highlight three main empirical results, all consistent with the model's key predictions. First, following US MP around Federal Open Market Committee (FOMC) meetings, exchange rates, and interest rate differentials for long-term bonds react in the direction suggested by a UIP condition

[5] See Fratzscher (2012), Albagli et al. (2019), Miranda-Agrippino and Rey (2020), Doidge et al. (2020) and Lilley et al. (2019) among others.

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

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

for most DEV, both individually and when tested as a panel. Specifically, a regression between the appreciation in country-j's currency and the change in the 10-year yield differential of that country vis–a–vis the US around the bracketed event yields a positive, statistically, and economically significant coefficient. An analogous result holds for domestic MP shocks in our DEV sample, albeit at a lower statistical significance. Second, as suggested by the model, the response of exchange rates and yield differences amplify any violation of 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 – with the notable exceptions of Switzerland and Japan, two well-known safe-haven countries.[8]

Third, these results are weaker for EME. In particular, we find a coefficient of essentially zero around MP events (both US and domestic) and a negative but only marginally significant relationship between interest rate differentials and exchange rates for uncertainty events in EME. These results are consistent with the tendency of emerging markets to endogenously adjust policy to stabilize their currency, as documented by Sarno and Taylor (2001), Ghosh et al. (2016) and Fratzscher et al. (2019). Recent evidence suggests that these reactions might end up feeding changes in the risk premium, eventually undoing the stabilization objectives of central banks and contributing to the empirical failure of the UIP.[9] Of course, the weaker absolute value of the correlations linked to our specific events could reflect the presence of other idiosyncratic shocks that we fail to control for empirically. Indeed, we document more significant volatilities of the key financial variables of interest in EME, both unconditionally and during event days.

Following Mueller et al. (2017), we explore an alternative empirical approach to evaluate the response of asset prices to monetary policy and uncertainty shocks. We compute the ex-post excess return of a one- and two-day investment strategy which goes long/short on domestic/US 10-year bonds and estimate how this excess return varies around monetary policy and uncertainty events. The behavior of ex-post excess returns reflects the reaction of yield differentials and exchange rates around these narrow windows. Consistent with our previous evidence, ex-post excess returns do not differ significantly around US monetary events relative to other days, revealing that yield differences 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, revealing that a domestic currency depreciation also accompanies an increase in yields in domestic bonds (vis-a-vis US bonds). Together, these results amplify any pre-existing deviation of the UIP condition. In synthesis, our results suggest that the UIP logic remains useful in predicting movements in financial markets following key monetary policy events, which we believe is a valuable policy insight. Moreover, our results are consistent with those in Lustig et al. (2019), who show that for a group of developed economies, unconditional average excess returns decline as the maturity of bonds increase. In our paper, around FOMC meetings the behaviour of longer-term

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

[9]See Pasquariello (2010) and Kalemli-Özcan (2019).

yields compensate the movement in exchange rates in such a way that there is no significant change in excess returns in DEVs. Different results follow from uncertainty shocks, which highlight the relevance of distinguishing across different types of shocks. This is, indeed, the main contribution of our paper.

Our paper is also related to a recent literature aimed at analyzing the endogenous relationship between exchange rates and long-term yields (see, for example, Gourinchas et al. (2022) and Greenwood et al. (2022)). In these papers – as in ours – limited bond market intergration leads to movements in term premia and exchange rates through a mechanism derived from the UIP logic investors demands 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. (2022) 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 behaviour 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 compementary to theirs.

Our work also connects to a broader literature that studies the relationship between capital flows and asset prices around specific events. A growing body of evidence documents a significant impact of US MP on a variety of asset prices, including long-term bonds in the US and abroad and stock markets.[10] Many have linked such relationship to the existence of a risk-taking channel of US MP –both conventional and unconventional–, whereby a more expansionary stance in the US leads to capital inflows into other countries, especially EMEs.[11] Although these papers do not analyze the implications on asset prices in line with a UIP relationship itself, we borrow from them the notion that US MP shocks, while essentially different from episodes of large fluctuations in uncertainty, are not completely orthogonal to movements in risk premia. In our model, we incorporate the possibility that monetary policy shocks can trigger changes in risk appetite. This assumption implies that US MP shocks tend to strengthen the positive relationship between interest rate differentials in favor of a given country and a contemporaneous appreciation (hence expected depreciation) of its currency. The fact that such relation is empirically non-significant for EME, precisely the group of countries where the US MP risk-taking channel is likely to be stronger,[12] is highly suggestive of the role played by defensive exchange rate policies.

Finally, our findings also relate to the literature evaluating the impact of uncertainty events on asset prices. Several recent papers document significant action of capital flows and asset prices towards EMEs around episodes of VIX movements and the effects of policy interventions, including

[10] See among others Bernanke and Kuttner (2005), Savor and Wilson (2014), Hanson and Stein (2015), Albagli et al. (2019) and Gilchrist et al. (2019).

[11] See Rey (2013), Obstfeld (2015), Bruno and Shin (2015), Kalemli-Özcan (2019) and Bhattarai et al. (forthcoming).

[12] See Kalemli-Özcan (2019), Akinci and Queralto (2019), Caballero and Kamber (2019).

conventional MP, foreign exchange interventions (FXI), and capital controls.[13] We borrow key insights from these papers to inform modeling choices governing variations in risk sentiment and capital flows around uncertainty events. We apply those insights within a simple framework that stresses the conditionality of the covariance of exchange rates and interest rate differentials on the origins of shocks, which we can then test under the laboratory of event studies.

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 reports our main empirical results around monetary policy and uncertainty events, respectively. Section 6 presents the evidence on the response of (ex-post) excess returns of different investment strategies to monetary policy and uncertainty shocks. Finally, 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 sources of 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

[13] di Giovanni et al. (2022) show evidence of the UIP failure in Turkey, finding that increases in the VIX raise firm-level credit risk and constrain bank access to international financing, driving up the cost of local currency loans relative to foreign currency borrowing. Caballero and Kamber (2019), Carrière-Swallow and Cespedes (2013) and Akinci et al. (2022) show that risk-off shocks raise long-term yields and depreciate local currencies in EMEs. Bhattarai et al. (2020) further discuss how policy responses vary among EMEs and the consequences they have on asset prices and capital flow volatility. Farhi and Gabaix (2016) show that risky countries command higher risk-reversal premia linked to world disasters, while Gourio et al. (2013) links the failure of the UIP to variations in disaster probability. More recently, Akinci et al. (2022) develop a model with endogenous risk premia leading to a high correlation between risk sentiment and UIP premia on foreign currencies.

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) into country j consist of foreign portfolio allocation into short-term (1-year) 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). Then, the level of K 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

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

is the price-inelastic capital flow into country-j's long-term bond (while all parameters can, in principle, be idiosyncratic to each country j, we will omit superscripts below for notational simplicity). Notice that we consider four components to this element. The first is a response to the risk-taking channel associated to US MP, with a country-specific loading of $\delta > 0$ (a tightening of US MP induces a net outflow). Second, we also allow for domestic MP to potentially trigger a risk-taking channel, with country-specific loading λ. Third, price-inelastic flows also load on the global uncertainty shocks v_t. While the shock is systematic to the international financial system, we assume a country-specific loading of κ (so a risk-off event, with $v_t > 0$, will imply a retreat in the price-inelastic component of flows if $\kappa > 0$). The fourth 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. 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:

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

Domestic MP is given by the shock m_t (a proxy for overall macroeconomic conditions, with persistence

$0 < \psi < 1$) and an endogenous reaction to exchange rate movements whose strength is given by the country-specific parameter d. Such reaction may reflect concerns about inflationary pressures, as well as financial stability considerations in heavily dollarized economies. Besides traditional MP, central banks may also stabilize their currency with FX interventions (I). Following Blanchard et al. (2015), we assume an offset parameter ϕ, such that $I_t = -\phi \cdot K_t$, and that 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 critical assumption of the model is that the demand of foreign investors responds positively to deviations from a version of the UIP condition for long-term interest rates. In other words, the difference between domestic yield differentials against US long-term bonds, 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, and defining $b \equiv \beta/(1 + \beta) < 1$ we obtain 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

$$a_1 = \frac{(1-\phi)\,(1+b/\beta^* + b\delta))}{\gamma + (1-\phi)\,(1+d-\rho+(1-\rho^h)b/h)},\; a_2 = \frac{(1-\phi)\,(1-\lambda b)}{\gamma+(1-\phi)\,(1+d-\psi+(1-\psi^h)b/h)}, \quad (9)$$
$$a_3 = \frac{(1-\phi)b\,(\kappa - 1/\beta^*)}{\gamma+(1-\phi)\,(1+d+b/h)}, \quad \text{and} \quad a_4 = \frac{(1-\phi)b}{\gamma+(1-\phi)\,(1+d+b/h)}.$$

While the sign of most coefficients is ambiguous ex-ante, the evidence below provides valuable guidance. The effect on the exchange rate of US MP (coefficient a_1) is always positive under the assumption that the loading of the "risk-taking channel" of US MP shocks on country-j's capital flows is negative ($\delta > 0$). Intuitively, an increase in US MP leads to capital outflows, both directly through a widening of the short term interest rate differentials (the endogenous response to deviations in the UIP condition) and indirectly through a lower risk appetite (the exogenous loading of capital flows given by δ). Capital outflows require a depreciation of local currencies to clear the balance of payment condition and reestablish equilibrium in bond markets by increasing expected returns in the local currency. The FXI parameter ϕ and local monetary policy response to exchange rate movements d mitigate 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.

The impact on the exchange rate of domestic monetary policy shocks is given by a_2. Consider first the case when $\lambda = 0$, implying that domestic monetary policy shocks do not trigger investors' risk-taking behavior. In this case, a_2 is unambiguously positive: a tightening of domestic policy has the (intuitive) effect of strengthening the domestic currency. This relationship, however, might become negative if a contractionary MP leads investors to anticipate a deterioration in domestic fundamentals and command a more 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. This scenario is the case for many EMEs,[14] and will be confirmed in our data for this group of countries as well.

Regarding an uncertainty shock, a "risk-off" episode (a positive v_t) will depreciate local currencies ($a_3 > 0$) as long as κ –the loading of capital flows into other countries in response to the uncertainty shock– is positive and sufficiently large. Indeed, a large value for κ implies that capital outflows depreciate the local currency, despite the compression in US long-term yield. This observation holds for most countries in the sample, with the notable exceptions of Japan and Switzerland, which respond as "safe havens" and appreciate relative to the USD following risk-off events ($a_3 < 0$). Finally, $a_4 > 0$, reflecting that a purely idiosyncratic capital flow shock into the local bond market will appreciate the currency as foreign capital enters the capital account.

To solve for local bond yields and the yield differential $y_t^{(h)} - y_t^{*(h)}$, we iterate forward Eq. (8),

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

use US-yields from Eq. (2), and replace both in the domestic bond-market clearing Eq. (7):

$$
\begin{aligned}
y_t^{(h)} - y_t^{*(h)} = (1-b)\Bigg[&- m_t^* \left(\beta/\beta^* - \delta + \frac{a_1(1-\rho^h)}{h}\right) + m_t \left(\lambda + \frac{a_2(1-\psi^h)}{h}\right) \\
&+ v_t \left(\beta/\beta^* + \kappa - \frac{a_3}{h}\right) - u_t \left(1 - \frac{a_4}{h}\right). \Bigg]
\end{aligned}
\tag{10}
$$

According to Eq. (10), tighter US monetary policy (a rise in m_t^*) reduces the domestic yield differential through the direct increase in $y_t^{*(h)}$ (Eq. 2) but also through the intensity of the risk-taking channel of US MP through country j's loading on m_t^*, δ. If δ is not too large, the risk-off in a specific country associated with tighter US MP will increase domestic yields but less so than in the US, thus lowering yield differentials. The effect of US MP on long-term yield differentials also depends on policy reaction, as reflected in a_1. If central banks aggressively intervene in the FX market or respond through defensive MP, a_1 will be small, but this enhances the reaction of domestic yields, closing the gap relative to the US. We document below that contractionary US MP generally lowers domestic long-term yield differentials (US Treasury yields go up by more than local yields).

The effect of domestic MP policy on long-term yield differentials, on the other hand, will be positive as long as λ is not too negative. The data below will generally confirm this positive relationship. Regarding uncertainty shock, a risk-off event ($v_t > 0$) will increase yield differentials with respect to the US whenever the country-loading κ is not too negative, also generally true in our sample. Finally, an increase in the idiosyncratic capital flow ($u_t > 0$) will, by construction, lower long-term yield differentials given the definition of the shock.[15]

2.4 From the model to the data

A good starting point for our analysis is the basic UIP condition, which can be written as:

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

where $y_t^{(h)}$ ($y_t^{*(h)}$) is the h-period yield on the domestic (foreign) instrument, s_t is the log of the spot exchange rate (US dollar value of unit of domestic currency), $E\left[s_{t+h}|\Omega_t\right]$ is the exchange rate expectation h-periods ahead, and $\eta_{t,t+h}$ is a risk-premium term. Because both $E\left[s_{t+h}|\Omega_t\right]$ and $\eta_{t,t+h}$ are unobservable, the empirical strategy usually imposes rational expectations, under which exchange rate h-periods ahead equals its contemporaneous expectation plus an error term; $s_{t+h} = E\left[s_{t+h}|\Omega_t\right] + \xi_{t,t+h}$. Together with Eq. (11), this yields $(s_t - s_{t+h})/h = y_t^{(h)} - y_t^{*(h)} + \varepsilon_{t,t+h}$,

[15] The precise conditions for the sign of the derivative of yield differentials with respect to shocks are: $\partial(y_t^{(h)} - y_t^{*(h)})/\partial m_t^* < 0$ if $\delta < \beta/\beta^* + \frac{(1-\phi)(1-\rho^h)}{h(\gamma+(1-\phi)(1+d-\rho))}(1+\beta/\beta^*)$; $\partial(y_t^{(h)} - y_t^{*(h)})/\partial m_t > 0$ if $\lambda > \frac{-(1-\psi^h)(1-\phi)}{h(\gamma+(1-\phi)(1+d-\psi))}$; and $\partial(y_t^{(h)} - y_t^{*(h)})/\partial v_t > 0$ if $\kappa > -\beta/\beta^*\left(1 + \frac{(1-\phi)}{h(\gamma+(1-\phi)(1+d))}\right)$.

where $\varepsilon_{t,t+h} = \eta_{t,t+h} - \xi_{t,t+h}/h$. Empirically, the most common specification of this condition is:

$$(s_t - s_{t+h})/h = a_0 + b_0(y_t^{(h)} - y_t^{*(h)}) + \varepsilon_{t,t+h}. \tag{12}$$

This equation, which should be valid at any horizon h, has been mainly tested for short horizons –usually between 1 and 12 months– as it becomes increasingly taxing in terms of data over longer horizons.[16] For example, for $h = 120$ (ten years), one would have to forfeit the last ten years of data –precisely our sample of interest.

Our ultimate object of interest is the covariance between the exchange rate and interest rate differentials, conditional on the presence of different shocks. We propose an event-study methodology that helps identify the relationship between interest rate differentials and exchange rates around specific events. In terms of Eq. (11), as long as it holds at every t, it must be the case that it holds for two close dates $t_1 < t < t_2$. Then, differentiating Eq. (11) we get

$$(\Delta s_t - \Delta E\left[s_{t+h}|\Omega_t\right])/h = \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_{t,t+h}, \tag{13}$$

where $\Delta x_t = x_{t_2} - x_{t_1}$ for variable x, and $\nu_{\tau,\tau+h} \equiv \Delta\eta_{t,t+h}$. 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. 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. This is simply to say that using future spot exchange rates to proxy for contemporaneous expectations may be valid to estimate Eq. (13) when considering multiple events that transpire over a month or a year, but invalid to study its behavior around specific ones. Thus, the estimation of Eq. (13) for shorter-term maturities requires the observation of exchange rate expectations at such maturities, which are not available at daily frequencies.

While this problem might be prevalent over any horizon, Appendix C presents evidence suggesting that these concerns become second-order for longer horizons. 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 usually lead to significant revisions in exchange rate expectations over relatively short horizons. However, this exchange rate expectations

[16] In the vast empirical literature estimating Eq. (12), the null-hypothesis values of $a_0 = 0$ and $b_0 = 1$ are often rejected in favor of a specification where risk premia leads to excess return predictability. See, for example, Lustig et al. (2014) and di Giovanni et al. (2022).

"term-structure" converges to zero for forecasts more than a few years ahead.[17] As our empirical strategy evaluates Eq. (13) around specific events, we provide additional evidence on the impact of MP shocks. In particular, we estimate regressions between exchange rate expectations at different horizons and our measures of monetary policy shocks, controlling for the orthogonal component (to MP shocks) of exchange rate fluctuations. The results show that exchange rate expectations respond in line with the shocks over the short run, meaning, for instance, that monetary shocks that lower the spot value of a currency would also weaken its expected value over short horizons, but these effects disappear over time.

Notice that these results are also consistent with our model, in which (persistent) monetary policy shocks affect the spot value of exchange rates and their rational expectation over short horizons. Over sufficiently long horizons, the model predicts no effect on the expected exchange rate of monetary policy shocks. Similar results follow from swings in risk-on/risk-off sentiments, whose effect on exchange rate expectations vanish over time. Building on this insight, we impose the condition $\Delta E\left[s_{t+h}|\Omega_t\right] \rightarrow 0$ for h sufficiently high, i.e. 10 years, on Eq. (13). We then get the following equation for estimating the relationship between exchange rates and long-term yield differences in response to specific shocks:

$$\Delta s_t/h = \hat{a} + \hat{b} \cdot \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_{t,t+h}. \tag{14}$$

where $\Delta x_t = x_{t_2} - x_{t_1}$ for variable x around two close dates $t_1 < t_2$. This empirical strategy is valid as long as we can safely isolate the nature of the shocks. In particular, we set an event study methodology around the sources of shocks, which include domestic and foreign MP shocks, as well as uncertainty events. As will become clear in Section 3, this strategy helps isolating the first three components of z_t from each other given the minimal overlap of events. However, the event window will nevertheless be contaminated by some idiosyncratic volatility, introduced in the model by the country-specific shock u_t. This may affect the strength, and as we will see, even the sign of the correlation between interest rate differentials and exchange rates.

Our model yields clear predictions about parameter $\hat{b}$ in Eq. (14), i.e., about the parameters that govern the response of exchange rates and long-term yields differences to different shocks.[18] Using Eqs. (8) and (10), we obtain $\hat{b}|_x$, where x is a specific shock in the regression $s_t/h = \hat{a} + \hat{b}\left(y_t^{(h)} - y_t^{*(h)}|x\right) + \mu_t$:

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

[18] 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.

$$\begin{aligned}
\hat{b}|_{m^*} &= \frac{\sigma_{m^*}^2\left(\beta/\beta^* - \delta + \frac{a_1}{h}(1-\rho^h)\right)\frac{a_1}{h} - \sigma_u^2\left(1-\frac{a_4}{h}\right)\frac{a_4}{h}}{(1-b)\left[\sigma_{m^*}^2\left(\beta/\beta^* - \delta + \frac{a_1}{h}(1-\rho^h)\right)^2 + \sigma_u^2\left(1-\frac{a_4}{h}\right)^2\right]}, \\
\hat{b}|_{m} &= \frac{\sigma_m^2\left(\frac{a_2(1-\psi^h)}{h} + \lambda\right)\frac{a_2}{h} - \sigma_u^2\left(1-\frac{a_4}{h}\right)\frac{a_4}{h}}{(1-b)\left[\sigma_m^2\left(\frac{a_2(1-\psi^h)}{h} + \lambda\right)^2 + \sigma_u^2\left(1-\frac{a_4}{h}\right)^2\right]}, \\
\hat{b}|_{v} &= \frac{-\sigma_v^2\left(\beta/\beta^* + \kappa - \frac{a_3}{h}\right)\frac{a_3}{h} - \sigma_u^2\left(1-\frac{a_4}{h}\right)\frac{a_4}{h}}{(1-b)\left[\sigma_v^2\left(\beta/\beta^* + \kappa - \frac{a_3}{h}\right)^2 + \sigma_u^2\left(1-\frac{a_4}{h}\right)^2\right]}.
\end{aligned} \tag{15}$$

Eq. (15) sets the stage for the key comparative statics of the model, which in turn guide its testable predictions. We summarize them in the next Proposition:

Proposition 1 (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* $\sigma_{m^*}^2(1/b + 1/\beta^* + \delta)\tilde{R}_1(\beta/\beta^* - \delta) > \sigma_u^2$, *where* $\tilde{R}_1 \geq 1$ *depends on model parameters. ii) Effect of policy:* $\partial\hat{b}|_{m^*}/\partial d < 0$ *and* $\partial\hat{b}|_{m^*}/\partial\phi < 0$ *iff* $\sigma_{m^*}^2(1/b + 1/\beta^* + \delta)\tilde{R}_1^2(\beta/\beta^* - \delta) > \sigma_u^2$. *iii) Effect of noise:* $\partial\hat{b}|_{m^*}/\partial\sigma_u^2 < 0$ *iff* $\beta/\beta^* > \delta$.

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

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

Proof. See Appendix A. □

Proposition 1 gives the key prediction of the model: the sign of the correlation between the exchange rate and interest rate differentials is conditional on the nature of the shock. For US MP shocks, the condition $\delta \geq 0$ implies that, following a contractionary US MP shock, the "risk-taking channel" of US MP will always ensure a depreciation of domestic currencies against the USD –which holds in our data for all countries. This mechanism strengthens the currency whose relative yields increase under two additional conditions. Firstly, the capital outflow term $\delta \cdot m_t^*$ must not be too high. Otherwise, long-term domestic yields will rise even more than US long-term yields, changing the sign of the yield differential.[19] The empirical results will show that, for all countries, long-term yield differentials shrink in response to contractionary US monetary policy shocks. However, this is not all. As Part a.i) of the proposition states, another way to break a UIP-type condition is the presence of idiosyncratic noise. Indeed, when σ_u^2 is large enough, the noise coming from unobservable idiosyncratic capital flows is strong enough to tilt the correlation in the opposite direction.[20]

[19] Formally, this condition is $\beta/\beta^* \geq \delta$.

[20] Notice that 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 a.iii) sheds light on this point by showing that an increase in the volatility of the idiosyncratic shock reduces the covariance. Another way to reduce the correlation $\hat{b}|_{m^*}$ is through policy interventions –either outright FXI (higher ϕ) or a defensive domestic MP (higher d)–, as both mitigate the response of the exchange rate and hence lower the absolute value of $\hat{b}|_{m^*}$. This latter mechanism will be crucial below when exploring the potential causes of a negative and significant UIP relationship for a few countries in our EME sample, conditional on US MP shocks.

Turning to local MP shocks, Proposition 1 indicates analogous conditions for a positive value $\hat{b}|_m$. A first restriction is $\lambda(1/b - \lambda) > 0$. The first root in this expression ($\lambda = 0$) captures the fact that, if $\lambda < 0$ (and discarding lower-order terms as $(1-b)/h$ and $1/h^2 \to 0$), a contractionary local MP shock will induce positive capital inflows, which compress long-term yields while appreciating the domestic currency, thus leading to a negative sign. The second root ($\lambda = 1/b$) reflects that when $\lambda > 1/b$, 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. Nevertheless, this is not all, as 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, the presence of noise reduces the correlation (and can even overturn it), as is the case for US MP shocks (i.e., even for $\lambda(1/b - \lambda) > 0$, large enough σ_u^2 will eventually shift the sign of $\hat{b}|_m$ into negative territory).[21]

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. The first root ($\kappa = 1/\beta^*$) follows from the exchange rate effect, as $\kappa > 1/\beta^*$ is necessary for the local currency to depreciate against the USD after an increase in global uncertainty (a risk-off event). On the other hand, the second root ($\kappa = -\beta/\beta^*$) establishes the condition for the local yield to go up relative to US long-term rates, which happens whenever $\kappa \geq -\beta/\beta^*$, a less restrictive condition than the former due to the assumption that uncertainty events compress US long-term yields. Part c.ii) shows that policies (FXI and MP response) 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 shows a simulation in which the upper, center and bottom panels plot the combination of local exchange rates (vertical axis) and long-term yield differentials (horizontal axis) 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

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

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 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.[22] 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 the number of monetary policy meetings for each country in the sample period.[23] 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 difference in the 2-year bond yield in each country.[24] 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.[25]

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.

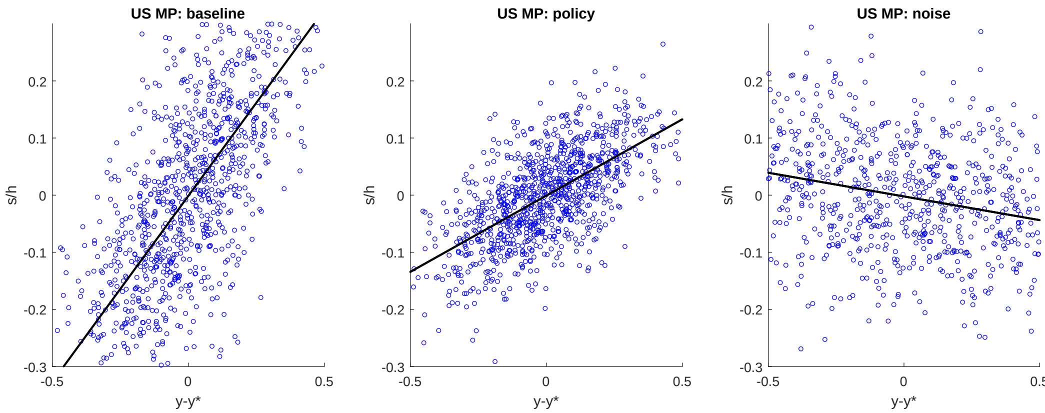
Table 1 documents the overlap between events, reporting the share of MPMs in DEV and EME

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

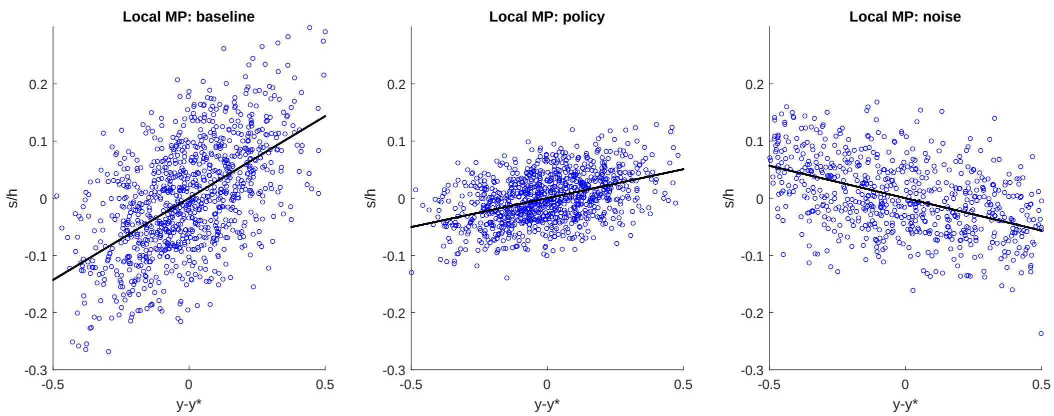
[23] The list of selected countries responds to data availability. Because our focus is on the UIP relationship around events, it is critical to collect daily data for the analysis. This differentiates our work relative to previous literature, which covers a more extended period but uses lower frequency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

[24] For example, for the FOMC meeting held on Wednesday, December 11, 2019, we compute the difference between the 2-year treasury yield at the close of Tuesday, December 10 and the close of Thursday, December 12.

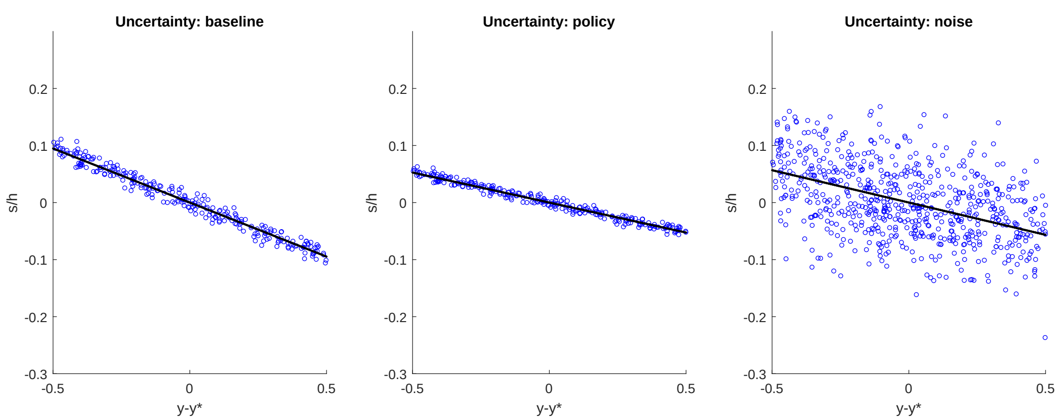
[25] All results hold when we compute the standard deviation of the daily change in the VIX for the November 2008-December 2019 sample. In Appendix E we propose a series of robustness checks for our definition and identification of VIX events.



(a) US MP shocks



(b) Local MP shocks



(c) Uncertainty shocks

Figure 1: Exchange rates and yield differences around events: model simulation

Notes:This figure simulates the relationship between exchange rates and yield differences derived from the model. 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 the analogous comparative statics for the case of local MP shocks, and uncertainty shocks, respectively. In the baseline simulation model parameters are as follows: i) shock 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: $\delta = 0.1$, $\lambda = 0.5$, $\kappa = 6$; iv) Local macroeconomic parameters: $\gamma = 0.1$, $\psi = 0.5$, $d = \phi = 0$, $\beta = 3$.

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, 2020; Doidge et al., 2020; Lilley et al., 2019). 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)".[26] 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.[27] 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 compare 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

[26] Why Americas financial plumbing has seized up, *The Economist*, March 21, 2020 edition.

[27] During this period, we have 89 FOMC meetings, 1161 and 1228 domestic MPM for DEV and EME, respectively, and 138 VIX events. As shown in table 1, there is limited overlap across these events, both for DEV and EME.

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.

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	-0.39	7.95***	-1.23	18.97***	0.18	0.00
VIX	-1.76	9.43***	-0.27	24.75***	0.15	0.00
No event	-0.01	6.10***	-0.10	15.83***	0.15	0.00
Unconditional	-0.20	6.61	-0.25	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	-2.69	124.95***	-1.76	100.01**	1.56	0.00
VIX	-18.81	140.87***	-25.27	142.87***	0.97	0.57
No event	1.45	82.42***	1.28	86.71***	0.90	0.00
Unconditional	-0.50	94.06	-0.91	95.49	0.97	0.01

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 9.43 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 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. (14), $\Delta s_t/h = \hat{a} + \hat{b} \cdot \Delta(y_t^{(h)} - y_t^{*(h)}) + \nu_{t,t+h}$ around US MP events. 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), (10) 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.

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.[28]

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 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

[28] 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				
DEV	0.47***	(0.08)	0.01	34584
EME	-0.10*	(0.06)	0.00	34584
Panel B: US MP and exchange rates				
DEV	-1.00***	(0.07)	0.01	34584
EME	-0.59***	(0.08)	0.00	34584
Panel C: US MP and 10-year differentials				
DEV	-0.97***	(0.06)	0.02	34584
EME	-0.94***	(0.09)	0.00	34584
Panel D: US MP and US 10-year yield				
US	1.41***	(0.17)	0.03	2882

Notes: This table presents regressions around FOMC events. Panel A 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 local and the US 10-year yields, respectively. Panel B 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. Panel C runs the regression $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b}\Delta i_t^* + \nu_{t,t+h}$. Finally, Panel D runs the regression $\Delta y_t^{*(h)} = \hat{a} + \hat{b}\Delta i_t^* + \nu_{t,t+h}$. In all specifications $\Delta x_t = x_t - x_{t-2}$, when in $t-1$ there is a FOMC meeting. 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.

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 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 increase 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 policy reaction by central banks, either through FXI or defensive MP reaction. 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.[29] 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. (14) 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 highlighted by the model.

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

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.[30] 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. Our model can interpret these results as if unconventional policies trigger a larger risk-taking behavior, which is accounted for by parameter δ. A higher δ unambiguously amplifies the exchange rate response but has an ambiguous effect on yield differences because a larger response of US long-term yields is also accompanied by a larger reaction of foreign long yields.

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.

4.2 Domestic MP shocks

Panel A in Table 5 reports results of Eq. (14) 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

[30] 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.

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												
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: US MP shock and NER												
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: US MP shock and 10-year differential												
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 MP shock and US 10-year yield												
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. Panel A 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 local and the US 10-year yields, respectively. Panel B 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. Panel C runs the regression $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b}\Delta i_t^* + \nu_{t,t+h}$. Finally, Panel D runs the regression $\Delta y_t^{*(h)} = \hat{a} + \hat{b}\Delta i_t^* + \nu_{t,t+h}$. In all specifications $\Delta x_t = x_t - x_{t-2}$, when in $t-1$ there is a FOMC meeting. 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). 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.

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 5 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.

TABLE 5: Around domestic MPM events

	$\hat{b}$	SE	R^2	N
Panel A: UIP				
DEV	0.27*	(0.12)	0.00	28820
EME	-0.10	(0.05)	0.00	34584
Panel B: Domestic MP shock and NER				
DEV	0.40*	(0.20)	0.00	28820
EME	-0.15***	(0.05)	0.00	34584
Panel C: Domestic MP shock and 10-year differential				
DEV	0.39***	(0.07)	0.01	28820
EME	0.90***	(0.15)	0.01	34584
Panel D: Domestic MP shock and U.S. 10-year yield				
DEV	0.20*	(0.11)	0.00	28820
EME	0.02	(0.01)	0.00	34584

NOTES: This table presents regressions around domestic MPM events. Panel A 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. Panel B 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. Panel C runs the regression $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b}\Delta i_t + \nu_{t,t+h}$. Finally, Panel D runs the regression $\Delta y_t^{*(h)} = \hat{a} + \hat{b}\Delta i_t + \nu_{t,t+h}$. In all specifications $\Delta x_t = x_t - x_{t-2}$, when in $t-1$ there is a domestic MPM. 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.

These findings are informative of the nature of the failure of the UIP relationship in EME, where the reaction of the exchange rate to domestic MP policy shocks is inconsistent with a UIP-type relationship. In particular, we observe that, on average, domestic currencies depreciate relative to the US dollar when domestic MP is perceived as more contractionary. While perhaps counterintuitive, similar findings have been documented by Kohlscheen (2013) and Hnatkovska et al. (2016). In our model, such effect obtains for values $\lambda > 1/b$, as in such case, domestic MP shocks trigger large enough (price-inelastic) outflows to more than compensate for the more intuitive effect of higher short-term interest rates. Our evidence supports this interpretation for EME, as domestic MP shocks both increase domestic long-term yield differentials and weaken EME currencies.

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.

5 Uncertainty Shocks

Table 6 reports the results of regression (14) 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. This result suggests that in line with Propositions 1.c.ii) and 1.c.iii), the degree of policy reaction or idiosyncratic noise may be playing some role.

Table 6: Around VIX events

	$\hat{b}$	SE	R^2	N
Panel A: UIP				
DEV	-0.19***	(0.05)	0.00	28820
EME	-0.17*	(0.10)	0.01	34584
Panel B: VIX and exchange rates				
DEV	-5.73***	(0.00)	0.02	28820
EME	-6.10***	(0.00)	0.02	34584
Panel C: VIX and 10-year differentials				
DEV	3.38***	(0.00)	0.01	28820
EME	8.12***	(0.00)	0.01	34584
Panel D: VIX and US 10-year yield				
US	-5.73***	(0.00)	0.03	2882

Notes: This table presents regressions around risk-premium events. Panel A 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. Panel B runs the regression $\frac{\Delta s_t}{10} = \hat{a} + \hat{b}\Delta\text{VIX}_t + \nu_t$. Panel C runs the regression $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b}\Delta\text{VIX}_t + \nu_{t,t+h}$. Finally, Panel D runs the regression $\Delta y_t^{*(h)} = \hat{a} + \hat{b}\Delta\text{VIX}_t + \nu_{t,t+h}$. In all specifications $\Delta x_t = x_t - x_{t-2}$, when in $t-1$ there is 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 627 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 6 report the results of regressions that analyze 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 627 bps– we normalize the change in VIX by said standard deviation to facilitate the interpretation of the regression coefficients.[31]

Regarding exchange rates, a one-standard deviation in the VIX increase leads to domestic currency depreciation of 0.57% and 0.61% 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 3.4 bps in DEV and 8.1 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. This result is consistent with the findings in Kalemli-Özcan and Varela (2021) 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.

Country-level results are reported in Appendix D. Similar to our baseline results, we find that $\hat{b}$ 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 6, 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 E.2. Aside from extending 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

[31] 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.

between financial variables during March 2020 lead us to discard this specification as the preferred one.

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 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 in the context of event studies.

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 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 and are, by themselves, novel results in the literature.

Following Mueller et al. (2017), 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 to day $t+k$. We compute expost currency excess returns using h-period bonds as:

$$rx_{t+k}^{(h)} = r_{t+k}^{(h)} - {r_{t+k}^{*}}^{(h)} + \Delta s_{t+k}, \tag{16}$$

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 (16) 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:

$$rx_{t+k}^{(h)} = \alpha_{t+k}^{(h)} + \beta_{t+k}^{(h)}\text{Event}_t + \epsilon_{t+k}^{(h)}, \tag{17}$$

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 (16), as well as for each of its components (domestic and US interest rate returns as well as exchange rate returns). Panel A of Table 7 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 7: 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***	-136.92***
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***	-62.01***

NOTES: This table presents regressions for excess returns and their components around events as in Eq. (17). Numbers correspond to basis points. Excess returns and components are constructed as in Eq. (16) 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 of Table 7 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 7. There is some (weak) evidence that excess returns increase between 11 and 17 bps around FOMC events relative to non-FOMC days. The increase in excess returns for EME comes from the more muted response of exchange rate returns to monetary policy shocks. Indeed, this is consistent with the evidence presented in Table 3, which shows that exchange rate responses in EMEs to US monetary policy

[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 = 0 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).

shocks are weaker than in DEV. This translates into a higher ex-post excess return of a strategy that is long in long-term domestic bonds.

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. (17) 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 7. For DEV and EME, we find significant decreases in excess returns in response to uncertainty shocks. These results reveal both the fall in interest rate differentials in favor of the US and the domestic currency depreciation. Both adjustments contribute to generating a fall in ex-post excess returns, consistent with the estimates in Section 5. Also, the large effect in excess returns 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, revealing much larger capital outflows from EMEs.

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 57 bps, while in EME is 107 bps. This means that, as reported in Panel B in Table 7, excess returns in EME around VIX events are 1.3 (=137/107) standard deviations larger than otherwise, and 1.40 time (=80/57) 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 countries, their currencies depreciate. 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. In line with our model, this could result from extensive use of defensive policies to stabilize exchange rates, which has been widely documented for this group of countries.

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. 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 long-term interest rate volatility –perhaps a less appreciated trade-off.

References

Akinci, O., Ş. Kalemli-Özcan, and A. Queralto (2022): "Uncertainty Shocks, Capital Flows, and International Risk Spillovers," Staff Reports 1016, Federal Reserve Bank of New York.

Akinci, O. and A. Queralto (2019): "Exchange Rate Dynamics and Monetary Spillovers with Imperfect Financial Markets," *mimeo*.

Albagli, E., L. Ceballos, S. Claro, and D. Romero (2019): "Channels of US Monetary Policy Spillovers to International Bond Markets," *Journal of Financial Economics*, 134, 447–473.

Bernanke, B. and K. Kuttner (2005): "What Explains the Stock Market's Reaction to Federal Reserve Policy?" *Journal of Finance*, 60, 1221–1257.

Bhattarai, S., A. Chatterjee, and W. Park (2020): "Global Spillover Effects of U.S. Uncertainty," *Journal of Monetary Economics*, 114, 71–89.

——— (forthcoming): "Effects of US Quantitative Easing on Emerging Market Economies," *Journal of Economic Dynamics and Control*.

Blanchard, O., G. Adler, and I. de Carvalho Filho (2015): "Can Foreign Exchange Intervention Stem Exchange Rate Pressures from Global Capital Flow Shocks?" Working Paper 21427, National Bureau of Economic Research, Cambridge, MA.

Bruno, V. and H. Shin (2015): "Cross-border Banking and Global Liquidity," *Review of Economic Studies*, 82, 535–564.

Caballero, R. and G. Kamber (2019): "On the Global Impact of Risk-off Shocks and Policy-put Frameworks," Working Paper 26031, National Bureau of Economic Research, Cambridge, MA.

Carrière-Swallow, Y. and L. Cespedes (2013): "The Impact of Uncertainty Shocks in Emerging Economies," *Journal of International Economics*, 90, 316–325.

Chinn, M. (2006): "The (Partial) Rehabilitation of Interest Rate Parity in the Floating Rate Era: Longer Horizons, Alternative Expectations, and Emerging Markets," *Journal of International Money and Finance*, 25, 7–21.

Chinn, M. and G. Meredith (2004): "Monetary Policy and Long-Horizon Uncovered Interest Parity," *IMF Staff Papers*, 51, 409–430.

di Giovanni, J., Ş. Kalemli-Özcan, M. Fatih Ulu, and Y. Soner Baskaya (2022): "International Spillovers and Local Credit Cycles," *Review of Economic Studies*, 89, 733–773.

Doidge, C., A. Karolyi, and R. Stulz (2020): "Is Financial Globalization In Reverse After The 2008 Global Financial Crisis? Evidence From Corporate Valuations," Working Paper 27022, National Bureau of Economic Research.

Domanski, D., E. Kohlscheen, and R. Moreno (2016): "Foreign exchange market intervention in EMEs: what has changed?" *BIS Quarterly Review*, September, 65–79.

Engel, C. (2014): "Exchange Rates and Interest Parity," in *Handbook of International Economics*, ed. by G. Gopinath, E. Helpman, and K. Rogff, Elsevier, vol. 4, 453–522.

——— (2016): "Exchange Rates, Interest Rates, and the Risk Premium," *American Economic Review*, 106, 436–474.

Fama, E. (1984): "Forward and Spot Exchange Rates," *Journal of Monetary Economics*, 14, 319–338.

Farhi, E. and X. Gabaix (2016): "Rare Disasters and Exchange Rates," *The Quarterly Journal of Economics*, 131, 1–52.

Fratzscher, M. (2012): "Capital Flows, Push Versus Pull Factors and the Global Financial Crisis," *Journal of International Economics*, 88, 341–356.

Fratzscher, M., O. Gloede, L. Menkhoff, L. Sarno, and T. Stöhr (2019): "When is Foreign Exchange Intervention Effective? Evidence from 33 Countries," *American Economic Journal: Macroeconomics*, 11, 132–156.

Froot, K. and T. Ito (1989): "On the Consistency of Short-run and Long-run Exchange Rate Expectations," *Journal of International Money and Finance*, 8, 487–510.

Ghosh, A., J. Ostry, and M. Chamon (2016): "Two Targets, Two Instruments: Monetary and Exchange Rate Policies in Emerging Market Economies," *Journal of International Money and Finance*, 60, 172–196.

Gilchrist, S., V. Yue, and Zakrajsek (2019): "U.S. Monetary Policy and International Bond Markets," *Journal of Money, Credit and Banking*, 51, 127–161.

Gourinchas, P., W. Ray, and D. Vayanos (2022): "A Preferred-Habitat Model of Term Premia, Exchange Rates, and Monetary Policy Spillovers," Working Paper 29875, National Bureau of Economic Research.

Gourio, F., M. Siemer, and A. Verdelhan (2013): "International Risk Cycles," *Journal of International Economics*, 892, 471–484.

Greenwood, R., S. Hanson, J. Stein, and A. Sunderam (2022): "A Quantity-Driven Theory of Term Premia and Exchange Rates," *mimeo.*

Habib, M. and L. Stracca (2012): "Getting Beyond Carry Trade: What Makes a Safe Haven Currency?" *Journal of International Economics*, 87, 50–64.

Hansen, L. and R. Hodrick (1980): "Forward Exchange Rates as Optimal Predictors of Future Spot Rates: An Econometric Analysis," *Journal of Political Economy*, 88, 829–853.

Hanson, S. and J. Stein (2015): "Monetary Policy and Long-term Real Rates," *Journal of Financial Economics*, 115, 429–448.

Hassan, T. and R. Mano (2019): "Forward and Spot Exchange Rates in a Multi-currency World," *Quarterly Journal of Economics*, 134, 397–450.

Hnatkovska, V., A. Lahiri, and C. Vegh (2016): "The Exchange Rate Response to Monetary Policy Innovations," *American Economic Journal: Macroeconomics*, 8, 137–181.

Hofmann, B., I. Shim, and H. Shin (2019): "Bond Risk Premia and the Exchange Rate," *BIS Working Paper*, 775.

Kalemli-Özcan, Ş. (2019): "U.S. Monetary Policy and International Risk Spillovers," Working Paper 26297, National Bureau of Economic Research, Cambridge, MA.

Kalemli-Özcan, Ş. and L. Varela (2021): "Five Facts about the UIP Premium," Working Paper 28923, National Bureau of Economic Research.

Kohlscheen, E. (2013): "The Impact of Monetary Policy on the Exchange Rate: A High Frequency Exchange Rate Puzzle in Emerging Economies," *Journal of International Money and Finance*, 44, 69–96.

Lilley, A., M. Maggiori, B. Neiman, and J. Schreger (2019): "Exchange Rate Reconnect," *mimeo.*

Lustig, H., R. Nikolai, and A. Verdelhan (2014): "Countercyclical Currency Risk Premia," *Journal of Financial Economics*, 111, 527–553.

Lustig, H., A. Stathopoulos, and A. Verdelhan (2019): "The Term Structure of Currency Carry Trade Risk Premia," *American Economic Review*, 109, 4142–4177.

Meese, R. and K. Rogoff (1983): "Empirical Exchange Rate Models of the Seventies. Do They Fit Out of Sample?" *Journal of International Economics*, 14, 3–24.

Miranda-Agrippino, S. and H. Rey (2020): "The Global Financial Cycle After Lehman," *AEA Papers and Proceedings*, 110, 523–528.

Mueller, P., A. Tahbaz-Salehi, and A. Vedolin (2017): "Exchange Rates and Monetary Policy Uncertainty," *Journal of Finance*, 72, 1213–1252.

Obstfeld, M. (2015): "Trilemmas and Trade-offs: Living with Financial Globalisation," *BIS Working Paper*, 480.

Pasquariello, P. (2010): "Central Bank Intervention and the Intraday Process of Price Formation in the Currency Markets," *Journal of International Money and Finance*, 29, 1045–1061.

Ranaldo, A. and P. Soderlind (2010): "Safe Haven Currencies," *Review of Finance*, 14, 385–407.

Rey, H. (2013): "Dilemma Not Trilemma: The Global Financial Cycle and Monetary Policy Independence," Working Paper 21162, National Bureau of Economic Research, Cambridge, MA.

Sarno, L. and M. Taylor (2001): "Official Intervention in the Foreign Exchange Market: Is it Effective and, if so, How Does it Work?" *Journal of Economic Literature*, 39, 839–868.

Savor, P. and M. Wilson (2014): "Asset Pricing: A Tale of Two Days," *Journal of Financial Economics*, 113, 171–201.

Zhang, T. (2022): "Monetary Policy Spillovers through Invoicing Currencies," *Journal of Finance*, 77, 129–161.

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

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

We then apply the method of undetermined coefficients to solve for the equilibrium values of a_1 through a_4, which yields expression (9) in the main text.

A.2 Proof of Proposition 1

a.i) US MP shocks: sign

From expression (15),

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

Taking the limit as $1/h^2 \to 0$, we get $\hat{b}|_{m^*} > 0$ iff $\frac{a_1}{a_4}\left(\beta/\beta^*-\delta\right)\sigma_{m^*}^2 > \sigma_u^2$. Computing the ratio a_1/a_4 from (9) gives the condition in the proposition, where $\tilde{R}_1 \equiv \frac{\gamma+(1-\phi)(1+d+b/h)}{\gamma+(1-\phi)(1+d-\rho+b(1-\rho^h)/h)} \geq 1$.

a.ii) US MP shocks: effect of policy

We begin by noting that the derivative of the denominator of $\hat{b}|_{m^*}$ in expression (15) with respect to both policies is proportional to $(1-b)/h$, and can thus be ignored in the proposed limit. It follows that

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

Taking the ratio of the coefficient derivatives, we get that $\partial\hat{b}|_{m^*}/\partial(1-\phi) > 0$ (and thus $\partial\hat{b}|_{m^*}/\partial\phi < 0$) iff the condition of the proposition holds. It is straightforward to show that the same results apply for the derivative with respect to the other policy parameter, d.

a.iii) US MP shocks: effect of noise

Taking the derivative of $\hat{b}|_{m^*}$ in expression (15) with respect to the variance of the idiosyncratic shock, and taking the proposed limit, yields:

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

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

b.i) Domestic MP shocks: sign

From expression (15),

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

Taking the limit as $1/h^2 \to 0$, and computing the ratio a_2/a_4 from (9) gives the condition in the proposition, where $\tilde{R}_2 \equiv \frac{\gamma+(1-\phi)(1+d+b/h)}{\gamma+(1-\phi)(1+d-\psi+b(1-\psi^h)/h)} \geq 1$.

b.ii) Domestic MP shocks: effect of policy

We begin by noting that the derivative of the denominator of $\hat{b}|_m$ in expression (15) with respect to both policies is proportional to $(1-b)/h$, and can thus be ignored in the proposed limit. It follows that

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

Taking the ratio of the coefficient derivatives, we get that $\partial \hat{b}|_m / \partial(1-\phi) > 0$ (and thus $\partial \hat{b}|_m/\partial\phi < 0$) iff the condition of the proposition holds. It is straightforward to show that the same results apply for the derivative with respect to the other policy parameter, d.

b.iii) Domestic MP shocks: effect of noise

Taking the derivative of $\hat{b}|_m$ in expression (15) 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 (15),

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

Taking the limit as $1/h^2 \to 0$, and computing the ratio a_3/a_4 from (9) 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 (15) 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 $\left[\sigma_v^2(\kappa - 1/\beta^*)(\beta/\beta^* + \kappa) + \sigma_u^2\right] > 0$. It is straightforward to show that the same results apply for the derivative with respect to the other policy parameter, d. That is, the sign of the derivative of $\hat{b}|_v$ with respect to both policies is the exact opposite than the sign of $\hat{b}|_v$.

c.iii) Uncertainty shocks: effect of noise

Taking the derivative of $\hat{b}|_v$ in expression (15) 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 138.

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.

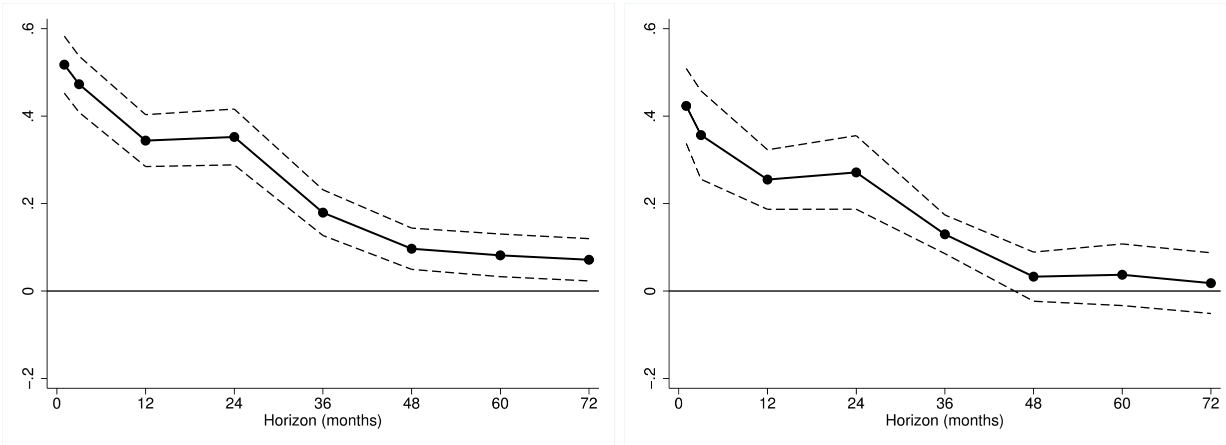
Consensus Forecast fixes the horizon of the expectation only in the case of 1, 3, and 12 months ahead. For $t+1$ through $t+6$, such horizon is mobile, depending on the month of the survey. For example, consider the expectations reported in January 2015. Since the expectation for the year $t+1$ corresponds 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 5. 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 5. 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 (14) 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)	-8.10***	(0.00)	1.52***	(0.00)
Japan	0.53***	(0.17)	5.32***	(0.00)	5.06***	(0.00)
United Kingdom	0.11	(0.22)	-4.98**	(0.00)	2.47***	(0.00)
Germany	0.10	(0.18)	-2.30	(0.00)	2.34***	(0.00)
Italy	-0.16**	(0.08)	-2.30	(0.00)	8.09***	(0.00)
France	-0.09	(0.16)	-2.30	(0.00)	4.03***	(0.00)
Australia	-0.24	(0.15)	-10.43***	(0.00)	1.56	(0.00)
New Zealand	-0.24	(0.16)	-9.10***	(0.00)	3.64***	(0.00)
Norway	-0.28*	(0.15)	-7.56***	(0.00)	2.23***	(0.00)
Sweden	-0.05	(0.20)	-6.20**	(0.00)	2.37***	(0.00)
Switzerland	0.37	(0.27)	0.71	(0.00)	3.85***	(0.00)
Czech Republic	-0.40**	(0.16)	-4.08	(0.01)	5.56***	(0.00)
Pooled	-0.09	(0.08)	-4.28***	(0.00)	3.56***	(0.00)
Pooled (ex JPN/SZ)	-0.19***	(0.05)	-5.73***	(0.00)	3.38***	(0.00)
Panel B: EME						
Chile	-0.06	(0.11)	-6.19***	(0.00)	5.42***	(0.00)
Colombia	-0.53***	(0.12)	-8.69***	(0.00)	11.06***	(0.00)
Hungary	-0.32***	(0.08)	-6.84*	(0.01)	16.56***	(0.00)
India	-0.11*	(0.06)	-3.63***	(0.00)	6.18***	(0.00)
Indonesia	-0.17***	(0.02)	-2.99***	(0.00)	15.69***	(0.00)
Israel	-0.02***	(0.01)	-2.87**	(0.00)	-2.63	(0.01)
Korea	-0.39**	(0.16)	-6.56***	(0.00)	5.27***	(0.00)
Mexico	-0.70***	(0.15)	-10.27***	(0.00)	10.04***	(0.00)
Poland	-0.66***	(0.14)	-7.15**	(0.00)	9.00***	(0.00)
Singapore	-0.15	(0.10)	-4.25***	(0.00)	3.59***	(0.00)
South Africa	-0.79***	(0.08)	-12.62***	(0.00)	11.22***	(0.00)
Thailand	-0.03	(0.02)	-1.08**	(0.00)	6.00***	(0.00)
Pooled	-0.17*	(0.10)	-6.10***	(0.00)	8.12***	(0.00)

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 627 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, 5 and 6, 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				
DEV	0.39***	(0.09)	0.43***	(0.08)
EME	-0.12*	(0.06)	-0.13*	(0.06)
Panel B: US MP and exchange rates				
DEV	-0.85***	(0.06)	-0.99***	(0.07)
EME	-0.52***	(0.07)	-0.58***	(0.08)
Panel C: US MP and 10-year differentials				
DEV	-0.89***	(0.06)	-0.96***	(0.06)
EME	-0.76***	(0.11)	-0.93***	(0.10)
Panel D: US MP and US 10-year yields				
US	1.27***	(0.19)	1.41***	(0.17)

Notes: This table presents regressions around FOMC events. Panel A 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 local and US 10-year yields, respectively. Panel B 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. Panel C runs the regression $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b}\Delta i_t^* + \nu_{t,t+h}$. Finally, Panel D runs the regression $\Delta y_t^{*(h)} = \hat{a} + \hat{b}\Delta i_t^* + \nu_{t,t+h}$. In all specifications $\Delta x_t = x_t - x_{t-2}$, when in $t-1$ there is a FOMC meeting. 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, 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				
DEV	0.20*	(0.10)	0.25*	(0.12)
EME	-0.10*	(0.06)	-0.09*	(0.05)
Panel B: Domestic MP shock and NER				
DEV	0.36	(0.20)	0.39*	(0.20)
EME	-0.11	(0.06)	-0.12**	(0.05)
Panel C: Domestic MP shock and 10-year differential				
DEV	0.39***	(0.08)	0.40***	(0.07)
EME	0.86***	(0.14)	0.88***	(0.15)
Panel D: Domestic MP shock and U.S. 10-year yield				
DEV	0.21*	(0.11)	0.20*	(0.11)
EME	0.03	(0.02)	0.03	(0.02)

Notes: This table presents regressions around domestic MPM events. Panel A 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. Panel B 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. Panel C runs the regression $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b}\Delta i_t + \nu_{t,t+h}$. Finally, Panel D runs the regression $\Delta y_t^{*(h)} = \hat{a} + \hat{b}\Delta i_t + \nu_{t,t+h}$. In all specifications $\Delta x_t = x_t - x_{t-2}$, when in $t-1$ there is a domestic MPM. 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. In all panels, the regression for developed countries excludes Japan and Switzerland. 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.

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 40, from 138 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 which 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								
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: VIX and exchange rates								
DEV	-4.88***	(0.00)	-4.90***	(0.00)	-5.28***	(0.00)	-5.52***	(0.00)
EME	-5.22***	(0.00)	-5.15***	(0.00)	-5.70***	(0.00)	-6.35***	(0.00)
Panel C: VIX and 10-year differentials								
DEV	3.06***	(0.00)	3.04***	(0.00)	3.22***	(0.00)	2.84**	(0.00)
EME	6.82***	(0.00)	6.72***	(0.00)	7.49***	(0.00)	7.45***	(0.00)
Panel D: VIX and US 10-year yields								
US	-4.00***	(0.00)	-3.89***	(0.00)	-5.46***	(0.00)	-4.68***	(0.00)

Notes: This table presents regressions around uncertainty events. Panel A 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. Panel B runs the regression $\frac{\Delta s_t}{10} = \hat{a} + \hat{b}\Delta\text{VIX}_t + \nu_t$. Panel C runs the regression $\Delta(y_t^{(h)} - y_t^{*(h)}) = \hat{a} + \hat{b}\Delta\text{VIX}_t + \nu_{t,t+h}$. Finally, Panel D runs the regression $\Delta y_t^{*(h)} = \hat{a} + \hat{b}\Delta\text{VIX}_t + \nu_{t,t+h}$. In all specifications $\Delta x_t = x_t - x_{t-2}$, when in $t-1$ there is 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 627 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.

F Unconditionl 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	0.33	56.74	0.31	106.59
Interest rate	0.52	36.79	0.73	86.46
Domestic	0.98	32.11	1.19	80.57
US	0.46	36.14	0.46	36.14
Exchange rate	-0.19	47.15	-0.42	47.84

Notes: This table presents unconditional descriptive statistics for excess returns and their components. Excess returns and components are constructed as in Eq.(16). 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.