A Model of U.S. Monetary Policy and the Global Financial Cycle*

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September 2026

Abstract

We propose a general equilibrium model in which U.S. monetary policy affects global risk premia by revaluing the wealth of currency-mismatched arbitrageurs. Assuming their portfolios are mean-variance efficient, arbitrageurs must be short the dollar. A U.S. tightening thus erodes arbitrageurs' net worth and raises the global price of risk. We discipline this mechanism to rationalize the effects of U.S. monetary policy on international asset prices, and we quantify the associated macroeconomic spillovers and role of policy responses abroad. In a potential future with higher dollar interest rates, arbitrageurs' dollar funding and U.S. monetary policy spillovers are dampened.

JEL codes: E44, E52, F31, G15

Keywords: monetary policy, exchange rates, global financial cycle

*We thank Mark Aguiar, Yacine Ait-Sahalia, Luigi Bocola, Ricardo Caballero, David Chapman, Max Croce, Anil Kashyap, Robert Kollmann, Alexandr Kopytov, Arvind Krishnamurthy, Sylvain Leduc, Wenhao Li, Guido Lorenzoni, Dmitry Mukhin, Brent Neiman, Richard Rogerson, Hyun Song Shin, Alp Simsek, Andreas Stathopoulos, Adrien Verdelhan, and Motohiro Yogo for discussions, and Sviatoslav Tiupin for excellent research assistance. This research has been supported by the National Science Foundation grant SES-2446790.

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1 Introduction

Over the past decade, a growing body of evidence has documented that changes in U.S. monetary policy have particularly large effects on risky asset prices in the rest of the world. The resulting "global financial cycle", with U.S. monetary policy at its center, has attracted an enormous amount of academic and policy attention. Yet important questions remain: what makes U.S. monetary policy special? What are the associated macroeconomic spillovers? How might changes in the economic environment affect these spillovers in the future? In this paper, we propose answers to these questions through the lens of a general equilibrium model that can capture the observed effects of U.S. monetary policy on global risk pricing.

We augment a canonical New Keynesian open economy model with global arbitrageurs who facilitate international capital flows and price assets around the world. The key mechanism is that when arbitrageurs are short the dollar, a U.S. monetary policy tightening that appreciates the dollar thus erodes arbitrageurs' net worth and raises the global price of risk. Consistent with this mechanism, the exchange rate and global bond yield effects of identified U.S. monetary surprises line up with average excess currency and bond returns in the data. Furthermore, observed asset prices imply that arbitrageurs must indeed be short the dollar on average, assuming their portfolios are mean-variance efficient. We estimate their leverage in non-dollar currencies to rationalize the observed effects of U.S. monetary policy shocks on the cross-section of international asset prices, and we quantify the implications for global economic activity and how these depend on policy responses abroad. We finally consider the possibility that dollar interest rates may rise in the future. In such an environment, arbitrageurs take smaller funding positions in dollars and U.S. monetary policy has smaller spillovers on the rest of the world.

We study a multi-country New Keynesian open economy environment with global arbitrageurs facilitating international capital flows. Households in each country supply labor used to produce differentiated final goods sold globally. Firms adjust prices subject to adjustment costs, and firms can differ in their currency of invoicing between any origin and destination market. Monetary policy is characterized by standard policy rules in each country, with an additive shock in the U.S. that is of primary interest. The key distinction relative to workhorse New Keynesian models is in the structure of financial markets. Households can only trade a local currency one-period bond freely,

and hold inelastic positions in all other assets. Global arbitrageurs are a subset of agents that can trade all assets freely and enforce no arbitrage. They maximize mean-variance preferences over wealth in the next period, and their effective risk aversion is declining in current wealth. We solve this model using a perturbation approach in which risk premia do not disappear, are time-varying, and affect allocations to first-order. We focus on trade in one-period bonds and geometrically decaying consols, so the relevant risk premia are currency and bond risk premia.

The propagation of a U.S. monetary policy shock in this model differs from standard New Keynesian models because it affects global risk pricing. We analytically characterize the relevant mechanisms by focusing on the effects on exchange rates in the useful limiting case of fully sticky prices. Absent any change in risk pricing, a U.S. monetary tightening appreciates the dollar versus all foreign currencies. There are two channels through which risk pricing can then change. First, if arbitrageurs are long foreign currencies entering the period, the dollar appreciation erodes their wealth and raises their price of bearing risk. This channel predicts that currencies offering high unconditional average excess returns ("risky" currencies) will depreciate by more. Second, the U.S. tightening induces capital flows from households, which changes the quantity of exchange rate risk absorbed by arbitrageurs — the "portfolio balance" mechanism. If in particular there are capital inflows to the U.S. because households' IES is below one, this channel predicts that currencies whose dollar exchange rate covaries more with the average dollar exchange rate should depreciate by less. The model makes similar predictions for the effect of a U.S. tightening on global consol prices, which reflect both a change in the price and quantity of risk. Anticipating our empirical finding that the portfolio balance mechanism does not well explain the observed asset price responses to conventional U.S. monetary surprises, we focus on the first channel operating through arbitrageur wealth and the price of risk.

Observed international asset prices imply that arbitrageurs must indeed be long non-dollar currencies, assuming that their portfolios are mean-variance efficient. Under mean-variance efficiency, the first and second moments of asset prices are sufficient to characterize the signs and relative size of arbitrageurs' positions in each non-dollar currency. Arbitrageurs' risk aversion only pins down the overall leverage of their portfolio, not the sign of their non-dollar positions. Using interest rates and exchange rates for the nine other "G10" economies besides the U.S. and for 13 emerging markets over 1991-2024, we characterize the implied composition of mean-

variance efficient portfolios. Focusing on trade in three-month and ten-year bonds, we find that a mean-variance efficient investor would be short G10 currencies, long emerging market currencies, and net long all non-dollar currencies in our sample. This conclusion is robust to the base currency of the investor.

The response of international asset prices to identified U.S. monetary surprises both validates and disciplines the model mechanism operating through the price of risk. Using high-frequency data, we first document that a surprise U.S. tightening is indeed associated with a larger depreciation among currencies paying high unconditional average excess returns ("risky" currencies), and a larger increase in yields among long-term bonds that pay high unconditional average excess returns ("risky" bonds). This cross-sectional relationship allows us to identify the risk aversion of arbitrageurs, which governs their leverage and thus wealth revaluation upon a U.S. tightening. We also demonstrate that in the data, a surprise Euro area policy tightening has effects on international asset prices that are more weakly related to average excess returns. The model rationalizes this (completely untargeted) feature of the data, owing to the dominant role of dollar funding on arbitrageur balance sheets.

We use the model to quantify the importance of risk premium spillovers of U.S. monetary policy, and how this depends on policy responses, in the rest of the world. The model first clarifies that the change in risk pricing dominates heterogeneous trade linkages or invoicing patterns in generating cross-country variation in the exchange rate and long-term yield responses to U.S. policy surprises. The change in risk pricing interacts with these other sources of heterogeneity, however, in shaping spillovers to the real economy. In particular, because emerging markets feature higher shares of dollar invoicing than other G10 economies, they face particularly sharp stagflationary consequences of a higher price of risk. These economies can address the risk premium shock induced by a U.S. tightening by selling dollar bonds through foreign exchange intervention. Indeed, comparing the observed exchange rate responses to those predicted by our model given these currencies' unconditional average excess returns, we quantify the extent to which emerging markets may be responding to a U.S. tightening by intervening in this way.

Looking forward, the model implies that a future with relatively higher dollar interest rates would feature dampened U.S. monetary policy spillovers through risk premia. Dollar interest rates may rise because of lower household demand for dollar

assets or increased issuance of dollar debt by governments.[1] Given a higher dollar interest rate and keeping all other first and second moments unchanged, the mean-variance efficient portfolio features a smaller long position in non-dollar currencies. Quantitatively, a $0.5pp$ higher dollar interest rate (modeled as arising from lower U.S. household patience) reduces arbitrageurs' portfolio share in non-dollar bonds by more than a factor of five. This dampens the effect of U.S. monetary policy on emerging markets, given the smaller revaluation of arbitrageur wealth. The dampened effects on risk pricing are reflected in dampened real spillovers: in response to a U.S. tightening, these economies experience a smaller economic contraction.

Related literature Our analysis connects to a large body of empirical work documenting that U.S. monetary policy surprises have significant effects on risky asset prices in global financial markets. Researchers have demonstrated this in the context of a common global factor in these prices (e.g., Rey (2015) and Miranda-Agrippino and Rey (2020)) as well as for individual asset classes, such as global equities (e.g., Boehm and Kroner (2026)), fixed income (e.g., Albagli, Ceballos, Claro, and Romero (2019), Gilchrist, Yue, and Zakrajsek (2019), and Kalemli-Ozcan (2020)), and currencies (e.g., Mueller, Tahbaz-Salehi, and Vedolin (2017)). Using portfolio data, Loualiche, Pecora, Somogyi, and Ward (2026) document that investment funds and global banks rebalance toward riskier currencies in response to a U.S. monetary easing. We contribute to this literature by providing an equilibrium account of the global risk premium effects of U.S. monetary policy, in which currency mismatch generates a link between the dollar exchange rate and arbitrageurs' risk-bearing capacity.

Our focus on wealth revaluation arising from dollar movements builds on a literature studying balance sheet effects in open economies. The partial equilibrium models of Bruno and Shin (2015b) and Jiang, Krishnamurthy, and Lustig (2024) also emphasize balance sheet effects as a key aspect of what makes U.S. monetary policy special. We embed this feature into a general equilibrium New Keynesian setting and trace out the implications for global risk pricing and economic activity. Prior work in the financial accelerator tradition has also studied balance sheet effects in open economy models.[2] Georgiadis, Mueller, and Schumann (2023) and Akinci and

[1] Building on the framework developed in this paper, Kekre and Lenel (2026) argue that a decline in the demand for dollar assets can indeed plausibly account for the joint dynamics of dollar yields and the dollar exchange rate around the announcement of U.S. tariffs on April 2, 2025.

[2] See, for instance, Cespedes, Chang, and Velasco (2004), Aghion, Bacchetta, and Banerjee

Queralto (2024) develop models in this tradition focused on the global spillovers from U.S. monetary policy. A key difference in our model is that imperfect substitutability across assets arises from risk, rather than intermediation frictions. The distinction matters — for instance, together with our focus on the cross-section of countries, it allows us to argue that the spillovers from U.S. monetary shocks are more consistent with a change in the price of risk than with portfolio balance effects.[3]

We conclude that a dollar appreciation erodes arbitrageur net worth based on an analysis of mean-variance efficient portfolios. Our approach complements an emerging asset pricing literature using mean-variance efficient portfolios to shed light on the sources of currency returns (Chernov, Dahlquist, and Lochstoer (2023, 2025)). It also provides a distinct but complementary perspective on dollar specialness relative to the literature on CIP-based measures of dollar convenience (Jiang, Krishnamurthy, and Lustig (2021), Du, Keerati, and Schreger (2025)): the dollar is special both because dollar interest rates are low relative to their replicating portfolio using foreign currency bonds, and because the risk of a dollar-funded portfolio is particularly low. Directly measuring arbitrageur exposures in the data is made difficult by trading in derivative contracts, and by the fact that it requires an initial (and typically untested) assumption about exactly who the arbitrageurs are.[4] That said, we believe our approach can be fruitfully combined with a growing literature estimating international portfolios using granular data.[5] Future work could compare the efficient portfolio to actual data on holdings to shed light on the identity of arbitrageurs that price the cross-section of exchange rates and international yield curves.

Our model builds on recent work in international finance studying risk premia and capital flows in segmented markets. We synthesize the perspectives of Gabaix and Maggiori (2015), Itskhoki and Mukhin (2021), and Kekre and Lenel (2024a) on the

(2007), Gertler, Gilchrist, and Natalucci (2007), Aoki, Benigno, and Kiyotaki (2021), Morelli, Ottonello, and Perez (2022), and Akinci, Kalemli-Ozcan, and Queralto (2026).

[3]We also use the cross-section to distinguish between the propagation of conventional U.S. monetary shocks and quantitative easing/tightening shocks: the latter appears to spill over to G10 currencies and yield curves primarily through the portfolio balance channel.

[4]Arbitrageurs' positions are also not evident from country-level cross-border portfolios because the latter contains the cross-border positions of households as well. This highlights a key distinction between the present model and models focused on the U.S.' external portfolio (i.e., Gourinchas, Rey, and Govillot (2017), Maggiori (2017), Kekre and Lenel (2024b), Dahlquist, Heyerdahl-Larsen, Pavlova, and Penasse (2025), and Sauzet (2025)).

[5]See, for instance, Maggiori, Neiman, and Schreger (2020), Du, Fontana, Jakubik, Koijen, and Shin (2023), Rey and Stavrakeva (2024), Dao, Gourinchas, and Itskhoki (2025), Hacioglu-Hoke, Ostry, Rey, Rousset Planat, Stavrakeva, and Tang (2026), and Koijen and Yogo (2026).

one hand, and Gourinchas, Ray, and Vayanos (2025) and Greenwood, Hanson, Stein, and Sunderam (2023) on the other. Like the first set of papers, we study a general equilibrium model in which arbitrageurs bear currency risk; like the second set of papers, arbitrageurs trade multiple risky assets, enforcing no arbitrage across these markets.[6] A key difference from all of these papers is that the risk-bearing capacity of arbitrageurs is endogenous, owing to the endogenous evolution of their wealth. The important role of arbitrageur wealth echoes themes from He and Krishnamurthy (2013) and Brunnermeier and Sannikov (2014), and is at the core of our explanation for why U.S. monetary policy affects the global price of risk.[7]

Our paper finally relates to a literature on the risk premium effects of monetary policy in the closed economy context. The model is an open economy counterpart to frameworks we have previously developed in which monetary shocks revalue the wealth of heterogeneous agents in equity and fixed income markets (Kekre and Lenel (2022) and Kekre, Lenel, and Mainardi (2026)). Other explanations for the risk premium effects of monetary policy are that monetary shocks affect the price of risk via the cost of taking leverage or habit preferences (Drechsler, Savov, and Schnabl (2018) and Pflueger and Rinaldi (2022)), change the quantity of risk on the balance sheets of agents who price risky assets (Hanson and Stein (2015) and Hanson, Lucca, and Wright (2021)), or communicate news about the policy rule and thereby future asset price comovements (Bianchi, Lettau, and Ludvigson (2021), Bauer, Pflueger, and Sunderam (2024), and Bianchi, Ludvigson, and Ma (2026)). While these mechanisms are all complementary, the wealth revaluation channel naturally explains why U.S. monetary policy generates disproportionate risk premium spillovers in the open economy context, owing to the dollar's dominant role as a funding currency.

Outline In section 2 we outline our proposed model, and in section 3 we analytically characterize the effects of U.S. monetary shocks. In section 4 we study mean-variance efficient portfolios and the effects of U.S. monetary shocks on international asset prices in the data. In section 5 we quantify the model using this evidence and study

[6] In contemporaneous and complementary work, Caballero and Simsek (2026) and Oskolkov and Perez (2026) also study open economy general equilibrium models in which arbitrageurs trade multiple risky assets. The focus of these papers — how the monetary rule affects the propagation of exogenous risk premium shocks, and the factor structure in exchange rates, respectively — differs from our focus on the conditional effects of monetary shocks on asset prices and allocations.

[7] Our focus on monetary shocks also harkens back to the earlier work of Alvarez, Atkeson, and Kehoe (2002, 2009) studying monetary policy, exchange rates, and risk premia in segmented markets.

a potential future with higher dollar interest rates. Finally, in section 6 we conclude.

2 Model

We first outline a multi-country New Keynesian model with global arbitrageurs who facilitate international capital flows and price the cross-section of assets.

2.1 Environment

We study an environment with $n > 1$ countries. We index the U.S. as country 1, and we refer to its unit of account (currency) as the dollar.

Households There are measure ζ_j households in each country j, with the normalization $\zeta_1 = 1$. The representative household in j has time separable preferences over consumption $c_{j,t}$ and labor supply $\ell_{j,t}$ with flow utility

$$u(c_{j,t}) = \frac{c_{j,t}^{1-1/\psi} - 1}{1 - 1/\psi} - \bar{\varphi}\frac{\ell_{j,t}^{1+1/\varphi}}{1 + 1/\varphi},$$

given an elasticity of intertemporal substitution ψ, Frisch elasticity of labor supply φ, disutility of labor supply controlled by $\bar{\varphi}$, and discount factor β_j. Consumption is a CES aggregator over bundles of differentiated goods from each country

$$c_{j,t} = \left(\sum_{k=1}^{n} \eta_{jk}^{\frac{1}{\varsigma}} \left(c_{jk,t}\right)^{\frac{\varsigma-1}{\varsigma}}\right)^{\frac{\varsigma}{\varsigma-1}},$$

where ς denotes the elasticity of substitution across domestic goods and imports, and η_{jk} controls the household's preference for goods originating in k. The consumption of goods produced in k is in turn a symmetric CES aggregator over a unit mass of varieties

$$c_{jk,t} = \left(\int_0^1 c_{jk,t}(s)^{\frac{\varepsilon-1}{\varepsilon}} ds\right)^{\frac{\varepsilon}{\varepsilon-1}},$$

with elasticity of substitution ε.

Relative to a standard open economy New Keynesian model à la Gali and Monacelli (2005), the key distinction in the present environment is the structure of asset markets. Households only trade one-period bonds and geometrically decaying consols

in each currency.[8] A one-period bond in currency k pays $1+i_{k,t}$ at $t+1$. The consol in currency k trades at price $P^{\delta}_{k,t}$ at date t, pays a coupon of one at $t+1$, and decays at rate δ (which thus controls its duration). We assume that household j can trade its local currency one-period bond frictionlessly, but holds exogenous positions in its local currency consol and all foreign bonds

$$\begin{aligned} B^{\delta}_{jj,t} &= P_{j,t}\bar{b}^{\delta}_{jj}, \\ B_{jk,t} &= P_{k,t}\bar{b}_{jk}, \ \ \forall k \neq j, \\ B^{\delta}_{jk,t} &= P_{k,t}\bar{b}^{\delta}_{jk}, \ \ \forall k \neq j, \end{aligned}$$

where $B_{jk,t}$ is the value of its position in the one-period bond in currency k, $B^{\delta}_{jk,t}$ is the value of its position in the consol in currency k, and $P_{k,t}$ is k's consumer price index. We can relax the assumption of exogenous positions in these bonds to assume that the household trades them subject to finite adjustment costs, without changing any of our main results.

The representative household enters the period with $A_{j,t}$ in financial wealth and receives a wage $W_{j,t}$ in units of local currency. It also receives lump-sum transfers $T_{j,t}$ and profits of local firms $\Pi_{j,t}$, and U.S. households receive net transfers from arbitrageurs $\xi\Pi_{a,t} - \xi P_{1,t}\omega_a$, all of which are described further below. Taken together, the resource constraint of household j is

$$\begin{aligned} \sum_{k=1}^{n}\int_0^1 P_{jk,t}(s)c_{jk,t}(s)ds + E_{j,t}\sum_{k=1}^{n} E^{-1}_{k,t}\left(B_{jk,t} + B^{\delta}_{jk,t}\right) = \\ W_{j,t}\ell_{j,t} + T_{j,t} + \Pi_{j,t} + 1\{j=1\}\left(\xi\Pi_{a,t} - \xi P_{1,t}\omega_a\right) + A_{j,t}, \end{aligned}$$

where $P_{jk,t}(s)$ is the price of variety s produced in k and sold in j, and $E_{j,t}$ is the nominal exchange rate (j's currency per dollar).[9] Financial wealth evolves according to

$$A_{j,t+1} = E_{j,t+1}\sum_{k=1}^{n} E^{-1}_{k,t+1}\left((1+i_{k,t})B_{jk,t} + \left(\frac{1+\delta P^{\delta}_{k,t+1}}{P^{\delta}_{k,t}}\right)B^{\delta}_{jk,t}\right).$$

[8] The effects of monetary policy on risk pricing would be evident with trade only in one-period bonds. We allow for trade in consols as well because currency and interest rate trading are typically integrated in practice. We leave extensions to additional assets, such as equities, to future work.

[9] Hence, we have $E_{1,t} = 1$.

Firms A unit mass of monopolistically competitive firms in j hire labor and produce differentiated varieties which are then sold globally. Following Gopinath, Boz, Casas, Diez, Gourinchas, and Plagborg-Moller (2020), we assume that a fraction θ_{kj}^m of firms set prices in destination k in invoicing currency m, where $\sum_{m=1}^n \theta_{kj}^m = 1$. This nests as special cases producer currency pricing ($m = j$), local currency pricing ($m = k$), and dollar pricing ($m = 1$). Consider a firm $s \in [0, 1]$ exporting from j to k with invoicing currency $\mathcal{M}(j, k, s) = m$:[10] we denote its chosen price $P_{kj,t}^m(s)$ and thus $P_{kj,t}(s) = E_{k,t}E_{m,t}^{-1}P_{kj,t}^m(s)$ the price facing households in currency k. We assume the firm faces Rotemberg (1982) adjustment costs when setting its price

$$\frac{P_{kj,t}^m y_{kj,t}^m \chi_p}{2}\left(\frac{P_{kj,t}^m(s)}{P_{kj,t-1}^m(s)} - 1\right)^2,$$

where χ_p controls the magnitude of price adjustment costs and $P_{kj,t}^m$ and $y_{kj,t}^m$ are aggregates that the firm takes as given.[11] The firm's profits from sales in k (expressed in its own currency) are thus

$$\Pi_{kj,t}^m(s) = E_{j,t}E_{m,t}^{-1}\left[P_{kj,t}^m(s)y_{kj,t}^m(s) - \frac{P_{kj,t}^m y_{kj,t}^m \chi_p}{2}\left(\frac{P_{kj,t}^m(s)}{P_{kj,t-1}^m(s)} - 1\right)^2\right] - W_{j,t}\ell_{kj,t}^m(s),$$

where $y_{kj,t}^m(s)$ is the quantity the firm sells in destination k and $\ell_{kj,t}^m(s)$ is the corresponding labor demanded for those sales. The firm internalizes its residual demand implied by household optimization across varieties

$$y_{kj,t}^m(s) = \left(\frac{P_{kj,t}(s)}{P_{kj,t}}\right)^{-\varepsilon} y_{kj,t},$$

where

$$P_{kj,t} = \left(\int_0^1 P_{kj,t}(s)^{1-\varepsilon} ds\right)^{\frac{1}{1-\varepsilon}}$$

is taken as given. Finally, the firm uses a linear technology, so that

$$y_{kj,t}^m(s) = \ell_{kj,t}^m(s).$$

[10] Thus, $\theta_{kj}^m = \int_0^1 1\{\mathcal{M}(j, k, s) = m\}ds$.

[11] We assume for simplicity that these adjustment costs are paid to the government and rebated lump-sum, so they do not affect aggregate resources.

In the usual way, the firm's vector of prior period prices are state variables in the firm's problem, and it makes optimal price setting decisions in each market using household j's stochastic discount factor.

Global arbitrageurs A unit mass of global arbitrageurs freely trade in all bonds and facilitate international capital flows. The representative arbitrageur operating at date t has wealth $W_{a,t}$ (denominated in dollars) and trades in bonds to satisfy the budget constraint

$$\sum_{j=1}^{n} E_{j,t}^{-1} \left(B_{aj,t} + B_{aj,t}^{\delta} \right) = W_{a,t}, \tag{1}$$

where $B_{aj,t}$ and $B_{aj,t}^{\delta}$ are the value of its position in currency j's bonds and consols, respectively. The profits earned by this representative arbitrageur in the next period are

$$\Pi_{a,t+1} = \sum_{j=1}^{n} E_{j,t+1}^{-1} \left((1 + i_{j,t}) B_{aj,t} + \left(\frac{1 + \delta P_{j,t+1}^{\delta}}{P_{j,t}^{\delta}} \right) B_{aj,t}^{\delta} \right). \tag{2}$$

A fraction ξ of arbitrageurs exit at the start of the next period and rebate their profits to U.S. households. These are replaced by new arbitrageurs endowed with $P_{1,t+1}\omega_a$ dollars from U.S. households, so that the wealth of the representative arbitrageur evolves according to

$$W_{a,t+1} = \xi P_{1,t+1}\omega_a + (1 - \xi)\Pi_{a,t+1}. \tag{3}$$

When making their portfolio choice decisions, we assume that the representative arbitrageur maximizes the myopic mean-variance objective

$$E_t \Pi_{a,t+1} - \frac{\gamma}{2W_{a,t}} Var_t \Pi_{a,t+1},$$

where γ controls the risk aversion of arbitrageurs. Crucially, their effective risk aversion is declining in their aggregate resources on hand $W_{a,t}$. This objective delivers approximately the same portfolio choice conditions as (myopic) CRRA utility with coefficient of relative risk aversion γ.[12]

[12]The resulting portfolio choice is also similar to that of risk neutral arbitrageurs facing a value-at-risk constraint. As we discuss in an extension with such arbitrageurs in section 5, the only difference in the value-at-risk microfoundation is the dynamic response of arbitrageur leverage to shocks.

Policy Let us define consumer price inflation in each country as

$$\pi_{j,t} = \log P_{j,t}/P_{j,t-1},$$

where the consumer price indices are

$$P_{j,t} = \left(\sum_{k=1}^{n} \eta_{jk} P_{jk,t}^{1-\varsigma} \right)^{\frac{1}{1-\varsigma}}.$$

We assume that monetary policy in each economy follows a Taylor (1993) rule with inertia and subject to a transitory shock

$$\log(1+i_{j,t}) = \log(1+i_j) + \rho_i \left(\log(1+i_{j,t-1}) - \log(1+i_j) \right) + (1-\rho_i)\phi_\pi \pi_{j,t} + \nu_{j,t},$$

where $\nu_{j,t}$ follows an AR(1) process with persistence ρ_j. The U.S. monetary shock, $\nu_{1,t}$, is our main stochastic disturbance of interest. We later consider richer policy responses by foreign central banks, including through exchange rate interventions. Finally, we assume that the fiscal authority rebates the price adjustment costs back to households using

$$T_{j,t} = E_{j,t} \sum_{k=1}^{n} \sum_{m=1}^{n} \theta_{kj}^{m} \frac{E_{m,t}^{-1} P_{kj,t}^{m} y_{kj,t}^{m} \chi_p}{2} \left(\frac{P_{kj,t}^{m}}{P_{kj,t-1}^{m}} - 1 \right)^2.$$

This uses that all firms in j selling to k in invoicing currency m will, in equilibrium, choose the same prices.

Market clearing The environment is closed with standard market clearing conditions in labor, goods, and bonds. The model can admit many more driving forces than the monetary policy shocks which will be our focus, such as to household tastes, technologies, and demands for foreign bonds or consols. As we further describe in the next section, our solution of the equilibrium will sidestep having to take a stand on all these driving forces or their stochastic properties.

2.2 Equilibrium and solution

The equilibrium admits a recursive representation in the real value of net foreign assets $a_{j,t} \equiv A_{j,t}/P_{j,t}$ for $j > 1$, the real wealth of arbitrageurs $w_{a,t} \equiv W_{a,t}/P_{1,t}$, the lagged real relative prices of goods relative to the consumption bundles $p^m_{jk,t-1} \equiv E_{j,t-1}E^{-1}_{m,t-1}P^m_{jk,t-1}/P_{j,t-1}$ and $p_{jk,t-1} \equiv P_{jk,t-1}/P_{j,t-1}$,[13] the lagged real exchange rates $q_{j,t-1} \equiv E_{j,t-1}P_{1,t-1}/P_{j,t-1}$, and the lagged nominal exchange rates $E_{j,t-1}$, as well as all exogenous state variables. Appendix A lists the model's equilibrium conditions.

We build on Itskhoki and Mukhin (2021), Borovicka, Hansen, and Sargent (2023), and Kekre and Lenel (2024a) in studying a first-order approximation of the equilibrium around a steady-state in which risk premia do not disappear. In particular, letting σ denote the perturbation parameter scaling the volatility of all innovations to exogenous state variables, we assume that the risk aversion of arbitrageurs is itself a function of this perturbation parameter, and that

$$\gamma(\sigma)\sigma^2 \to \Gamma$$

as $\sigma \to 0$. Under this assumption, and defining the excess returns to the one-period bond and consol in currency j

$$\begin{aligned} ex_{j,t+1} &\equiv \frac{E_{j,t}}{E_{j,t+1}}(1+i_{j,t}) - (1+i_{1,t}), \\ ex^{\delta}_{j,t+1} &\equiv \frac{E_{j,t}}{E_{j,t+1}}\frac{1+\delta P^{\delta}_{j,t+1}}{P^{\delta}_{j,t}} - (1+i_{1,t}), \end{aligned}$$

appendix A demonstrates that as $\sigma \to 0$ arbitrageurs' optimal choice of one-period bonds in currency j becomes

$$ex_j = \frac{\Gamma}{w_a}\left[\sum_{k=2}^{n} Cov(ex_{j,+1}, ex_{k,+1})q_k^{-1}b_{ak} + \sum_{k=1}^{n} Cov(ex_{j,+1}, ex^{\delta}_{k,+1})q_k^{-1}b^{\delta}_{ak}\right], \qquad (4)$$

where $w_{a,t} \equiv W_{a,t}/P_{1,t}$, $b_{ak,t} \equiv B_{ak,t}/P_{k,t}$, and $b^{\delta}_{ak,t} \equiv B^{\delta}_{ak,t}/P_{k,t}$. That is, in steady-state, arbitrageurs' required excess return on one-period bonds in currency j must compensate them for the associated risk, where this risk depends on the covariance of

[13] As demonstrated in appendix A, in the first-order solution of the model, we can exploit a bloc recursivity property to characterize invoicing-specific variables separately from all others. This allows us to drop $p^m_{jk,t-1}$ as state variables, which helps in reducing the computational burden.

the trade with all other trades and the arbitrageurs' positions. To first order around steady-state, appendix A demonstrates that this same optimality condition becomes

$$E_t \hat{ex}_{j,t+1} = -ex_j \hat{w}_{a,t} + \frac{\Gamma}{w_a} \left[\sum_{k=2}^{n} Cov(ex_{j,+1}, ex_{k,+1}) q_k^{-1} \left(-b_{ak} \hat{q}_{k,t} + \hat{b}_{ak,t} \right) + \sum_{k=1}^{n} Cov(ex_{j,+1}, ex_{k,+1}^{\delta}) q_k^{-1} \left(-b_{ak}^{\delta} \hat{q}_{k,t} + \hat{b}_{ak,t}^{\delta} \right) \right], \quad (5)$$

where $\hat{x}_t \equiv \log x_t - \log x$ except for variables which can be negative (such as excess returns and bond positions), in which case $\hat{x}_t \equiv x_t - x$. Analogous results hold for ex_j^{δ} and $E_t \hat{ex}_{j,t+1}^{\delta}$ for all j. In this way, currency and bond risk premia are time-varying even using a first-order approximation. The way in which currency and bond risk premia adjust in response to flows on arbitrageurs' balance sheet echoes Gourinchas et al. (2025) and Greenwood et al. (2023). The key differences from these papers are that these flows reflect the optimizing behavior of households in a multi-country New Keynesian environment; risk premia respond to changes in arbitrageur wealth, not just flows; and changes in risk premia affect short rates and consumption. More generally, risk premia and macroeconomic dynamics are jointly determined.

In equilibrium, the covariances of excess returns in (4) and (5) need to be consistent with the first-order dynamics of the model around the steady-state. These covariances depend on all of the potential driving forces in the model. But crucially, these covariances are "sufficient statistics" for the presence and effects of other shocks on the propagation of U.S. monetary shocks up to first order. In our analytical results in the next section, we will thus characterize the effects of U.S. monetary shocks taking as given these covariances. In our quantitative calibration later in the paper, we will treat these covariances as parameters disciplined directly by the data, avoiding the complexity of taking a stand on all the driving forces and their stochastic properties.[14] The implicit assumption is that there is some set of driving forces in this model which can rationalize the observed covariances of currency and bond excess returns. In parallel (ongoing) work, we are using this model to study the structural shocks which can rationalize the cross-section of currency and bond excess returns.

[14] Of course, where this can bite is when we study counterfactual environments, in which case these covariances might endogenously change.

3 Analytics of U.S. monetary policy transmission

We now analytically characterize the effects of a U.S. monetary policy shock in this environment. We focus on the impact effects on exchange rates and long-term bond prices given our focus on equilibrium risk pricing. If arbitrageurs hold foreign currency assets financed in dollars, a dollar appreciation induced by tighter U.S. monetary policy erodes their net worth and raises the global price of risk. Household capital flows induced by a U.S. monetary policy shock also change the risk exposures on arbitrageur balance sheets, affecting asset prices through portfolio balance.

3.1 Effects of U.S. monetary shock on exchange rates

We first characterize the effect of the U.S. monetary shock $d\nu_{1,t}$ on the full set of bilateral dollar exchange rates.

We provide this result in the simplifying case of fully sticky prices ($\chi_p \to \infty$) and local currency pricing ($\theta_{kj}^k = 1$ for all j, k), which imply that U.S. real interest rates fully inherit the dynamics of the monetary shock and foreign interest rates do not respond. While exchange rates have no expenditure switching role in this limiting case, they still jointly equilibrate goods and financial markets: they affect country budget constraints via the local currency value of foreign sales, and they affect the relative attractiveness of bonds denominated in different countries. We also assume zero consol positions $\{b_{jk}^\delta\}$, which simplifies the portfolio of arbitrageurs to be in one-period bonds alone. This does not prevent us from describing how a U.S. monetary shock affects consol prices, which we do in the next section. Finally, we assume no interest rate inertia ($\rho_i = 0$) and that the steady-state excess returns $\{ex_j, ex_j^\delta\}$, one-period bond positions $\{b_{jk}\}$, and effective risk aversion Γ/w_a are small. These assumptions simplify the mathematics without affecting the key economics.

We obtain:

Proposition 1. *Assume $\chi_p \to \infty$; $\theta_{kj}^k = 1$ for all j, k; zero $\{b_{jk}^\delta\}$ and ρ_i; and small $\{ex_j, ex_j^\delta, b_{jk}\}$ and Γ/w_a. Then given a U.S. monetary shock $d\nu_{1,t}$,*

$$\hat{E}_{j,t} = \hat{q}_{j,t} = \Bigg(\left[\beta^{-1} - \rho_1\right]^{-1} \left[1 + (\beta^{-1} - 1)(1 - \rho_1)^{-1}\psi\right] + \left[\beta^{-1} - \rho_1\right]^{-1} \left[1 + (\beta^{-1} - 1)(1 - \rho_1)^{-1}\psi\right] \beta \frac{1 - \xi}{\xi} \frac{\sum_{k=2}^n q_k^{-1} b_{ak}}{w_a} ex_j$$

$$- \left[\beta^{-1} - \rho_1\right]^{-2} \left[\beta^{-1} - 1\right]^{-1} [1 - \psi] \frac{\Gamma}{w_a} Cov \left(ex_{j,+1}, \sum_{k=2}^{n} \eta_{1k} ex_{k,+1} \right) \Bigg) d\nu_{1,t}. \quad (6)$$

The first line of (6) describes the response of exchange rates in the absence of any change in equilibrium risk pricing. It implies that a U.S. monetary tightening causes a depreciation of all foreign currencies. The foreign depreciation is rising in the persistence of the U.S. tightening ρ_1, and becomes arbitrarily large as the discount factor β and persistence ρ_1 approach one, reflecting the logic of Engel and West (2005). The foreign depreciation is also rising in households' intertemporal elasticity of substitution ψ, reflecting that U.S. consumption falls by more upon a U.S. monetary tightening, and thus the required depreciation to restore the local currency value of exports to the U.S. and equilibrate goods markets must rise. Relatedly, the foreign depreciation is symmetric across all economies regardless of their trade linkages. This reflects that, while economies that sell more to the U.S. experience a larger decline in revenues following the reduction in U.S. demand, they also experience a larger increase in revenues following a local currency depreciation versus the dollar.

The second line of (6) describes the effect on exchange rates arising from a revaluation in arbitrageur wealth and thus a change in risk pricing — our key mechanism of interest. When arbitrageurs hold foreign currency exposure $\sum_{k=2}^{n} q_k^{-1} b_{ak}$, the dollar appreciation induced by a U.S. monetary tightening reduces their wealth and thus raises their effective risk aversion. As a result, "risky" currencies ($ex_j > 0$) further depreciate relative to the benchmark described in the prior paragraph: these exchange rates further depreciate on impact so that they can deliver a higher expected appreciation, and thus higher excess return, going forward. Symmetrically, "safe" currencies ($ex_j < 0$) appreciate relative to the benchmark without a change in risk pricing. The key statistic governing the strength of this channel is

$$\frac{1 - \xi}{\xi} \frac{\sum_{k=2}^{n} q_k^{-1} b_{ak}}{w_a},$$

which is rising in the leverage of arbitrageurs in foreign currency assets, $\frac{\sum_{k=2}^{n} q_k^{-1} b_{ak}}{w_a}$, and the persistence of their wealth, $1 - \xi$. The former governs the initial wealth revaluation from a dollar appreciation, while the latter governs how persistently currency risk premia are affected, and thus how much exchange rates must move on impact.[15]

[15]These results suggest that the effect of a change in risk pricing on exchange rates can be

The third line of (6) clarifies that risk premia in this model are also affected by capital flows that change the quantity of risk arbitrageurs are exposed to — the "portfolio balance" mechanism. The U.S. monetary tightening induces cross-border capital flows in equilibrium. For $\psi < 1$, the U.S. experiences capital inflows, while for $\psi > 1$, foreign economies experience capital inflows. In the first case, this means that arbitrageurs are less exposed to foreign currencies. If excess returns on currency j covary positively with the trade-weighted average of excess returns (higher $Cov\left(ex_{j,+1}, \sum_{k=2}^{n} \eta_{1k} ex_{k,+1}\right)$), j's currency risk premium is diminished, and its exchange rate appreciates relative to the benchmark without a change in risk pricing. Note that $Cov\left(ex_{j,+1}, \sum_{k=2}^{n} \eta_{1k} ex_{k,+1}\right)$ can equivalently be interpreted as the covariance of j's dollar exchange rate with the trade-weighted average of dollar exchange rates — what the asset pricing literature describes as the "dollar beta" of currency j (Verdelhan (2018)). Interestingly, dollar betas are the relevant measure of risk exposures to capital flows in/out of the U.S. in intermediated financial markets.

3.2 Effects of U.S. monetary shock on consol prices

We now characterize the effect of the U.S. monetary shock on consol prices.

Under our simplifying assumption that household positions are fixed in consols, the presence of consols only affects the allocation by affecting arbitrageurs' risk pricing, and thus willingness to intermediate capital flows in one-period bonds. But when household consol positions are in particular zero — as assumed in this section alone — even this effect is absent, and the pricing of consols does not affect the rest of the allocation. Nonetheless, it is useful to still characterize the response of consol prices to a U.S. monetary shock, both to underscore that these responses also reflect a change in the price and quantity of risk, and to clarify what aspects of cross-country heterogeneity govern these channels in the case of long-term bonds.

We obtain:

Proposition 2. *Assume the same parametric conditions as in Proposition 1. Then given a U.S. monetary shock* $d\nu_{1,t}$*,*

arbitrarily large if ξ is sufficiently low, holding fixed arbitrageurs' leverage. This is an artifact of the fact that (6) is an approximation for small $\{ex_j\}$. As we later emphasize in our quantitative results, because high expected excess returns endogenously recapitalize arbitrageurs following a decline in their wealth, this constrains how much a decline in their wealth can move exchange rates on impact.

$$\hat{P}^{\delta}_{j,t} = \Bigg(- \left[1 - \delta\beta\rho_1\right]^{-1} 1\{j=1\}$$
$$- \left[\beta^{-1} - \rho_1\right]^{-1} \left[1 + (\beta^{-1} - 1)(1-\rho_1)^{-1}\psi\right] \frac{1-\xi}{1-\delta(1-\xi)} \frac{\sum_{k=2}^{n} q_k^{-1} b_{ak}}{w_a} \left(ex^{\delta}_j - ex_j\right)$$
$$+ \left[\beta^{-1} - \rho_1\right]^{-1} \left[\beta^{-1} - \delta\rho_1\right]^{-1} \left[1 - \psi\right] \frac{\Gamma}{w_a} Cov\left(ex^{\delta}_{j,+1} - ex_{j,+1}, \sum_{k=2}^{n} \eta_{1k} ex_{k,+1}\right)\Bigg) d\nu_{1,t}. \tag{7}$$

The first line of (7) describes the response of consol prices in the absence of any change in risk pricing. It implies that the U.S. consol price falls upon the U.S. tightening via the expectations hypothesis, with a bigger fall the higher is δ (which controls duration), discount factor β, and persistence ρ_1. With fully sticky prices in local currency, foreign consol prices are unaffected in the absence of any change in risk pricing, because the expected path of foreign interest rates is unchanged.

The second line of (7) describes the response of consol prices to a revaluation in arbitrageur wealth and thus a change in their price of bearing risk. The key insight here is that it is the average "local currency bond carry" return,

$$ex^{\delta}_j - ex_j = \frac{1 + \delta P^{\delta}_j}{P^{\delta}_j} - (1 + i_j),$$

which governs the response of consol j's price to a change in the price of risk. This is the excess return to a strategy of borrowing a unit of currency j using j's one-period bond and investing in j's consol. Intuitively, holding consols in currency j exposes arbitrageurs to two sources of risk: the risk that currency j depreciates, and the risk that j's consol price falls. A change in the arbitrageur's willingness to bear the former risk will be reflected in the exchange rate, while it is only a change in the arbitrageur's willingness to bear the latter risk that is reflected in the consol price. The latter depends in turn on how risky it is to hold j's consol abstracting from exchange rate risk, summarized by the average local currency bond carry return. Hence, in currencies with a steeper yield curve, a reduction in arbitrageur wealth would reduce long-term bond prices by more.

Away from the case of zero consol positions, the change in consol prices would feed back to affect arbitrageur wealth and thus their price of bearing risk. For small consol positions, small average excess returns, and small Γ/w_a, only the revaluation in wealth

driven by the expectations hypothesis would matter. In particular, if arbitrageurs are long dollar consols, their wealth would fall upon a U.S. tightening through a distinct channel than arbitrageurs' currency mismatch.[16] We focus on the currency mismatch channel in this paper because the duration channel is not specific to U.S. monetary policy: as we will later describe, the data implies that arbitrageurs have positive positions in long-term bonds in many currencies.

Finally, the third line of (7) describes the response of consol prices to capital flows in one-period bonds. As described in the last subsection, for $\psi < 1$, the U.S. experiences capital inflows and arbitrageurs are less exposed to foreign currencies. If j's local currency bond carry covaries positively with the average excess return on foreign currencies, then j's bond risk premium is diminished, and its consol price rises. It is again the covariance with the local currency bond carry return that is relevant because this isolates the risk in the consol price alone.

Taking stock A U.S. monetary tightening affects global risk pricing via a change in arbitrageurs' wealth and via the cross-border capital flows induced by the shock. The strength of the first mechanism relies on the composition of arbitrageur portfolios, and the relative importance of these mechanisms is in principle evident from the cross-section of asset price responses to identified U.S. monetary surprises. In the next section, we examine these features of the data.

4 Empirical analysis

In the data, the observed first and second moments of currency returns imply that mean-variance efficient investors are indeed net long foreign currencies, particularly emerging markets, and short the dollar. The cross-section of exchange rate and global yield curve responses to identified U.S. monetary shocks is consistent with an increase in the price of risk following a U.S. tightening, as due to lower arbitrageur wealth.

4.1 Mean-variance efficient portfolios

The first and second moments of currency returns pin down arbitrageurs' portfolios, up to leverage. This follows from the fact that arbitrageurs' steady-state portfolios

[16] In a closed economy, the revaluation of arbitrageur wealth through duration is the focus of Kekre et al. (2026).

solve a standard Markowitz (1952) problem. Letting $\boldsymbol{ex}$ denote the column vector of $\{ex_j, ex_j^\delta\}$, $\boldsymbol{C}$ denote the covariance matrix of $\{Cov\left(ex_{j,+1}, ex_{k,+1}\right), Cov\left(ex_{j,+1}, ex_{k,+1}^\delta\right), Cov\left(ex_{j,+1}^\delta, ex_{k,+1}^\delta\right)\}$, and $(\boldsymbol{q}^{-1}\boldsymbol{b}_a)$ denote the column vector of $\{q_j^{-1}b_{aj}, q_j^{-1}b_{aj}^\delta\}$, we can express the solution, (4) and the analog for ex_j^δ, as

$$\frac{1}{w_a}(\boldsymbol{q}^{-1}\boldsymbol{b}_a) = \frac{1}{\Gamma}\boldsymbol{C}^{-1}\boldsymbol{ex}, \tag{8}$$

assuming $\boldsymbol{C}$ is invertible. It follows that the expected excess returns $\boldsymbol{ex}$ and (conditional) covariance matrix of excess returns $\boldsymbol{C}$ together govern the sign and relative size of arbitrageurs' positions in non-dollar currencies, while risk aversion Γ governs the overall leverage in these non-dollar currencies.[17] While mean-variance efficiency is typically used to explain expected excess returns given covariances (since the representative agent must "hold the market"), here we invert this logic in a segmented markets environment: we use estimates of expected excess returns and covariances to pin down the portfolio of the marginal agent, up to leverage.

We use daily data on yield curves and exchange rates for nine advanced economies (which together with the U.S. are referred to as the "G10") and 13 emerging markets (EMs). The G10 economies besides the U.S. are Australia (AUD), Canada (CAD), Euro area (EUR), Japan (JPY), Norway (NOK), New Zealand (NZD), Sweden (SEK), Switzerland (CHF), and the U.K. (GBP). The emerging markets are Brazil (BRL), Colombia (COP), Hungary (HUF), India (INR), Indonesia (IDR), Mexico (MXN), Malaysia (MYR), Peru (PEN), Philippines (PHP), Poland (PLN), South Africa (ZAR), South Korea (KRW), and Thailand (THB). For each currency, we collect three-month and ten-year yields as well as dollar exchange rates from Bloomberg from April 1991 through February 2024.[18,19] We use government bond yields as our baseline measure of yields, but explore the consequences of using IBOR-based yields

[17] In equilibrium, of course, these positions and excess returns must be consistent with the optimizing behavior of households and market clearing.

[18] We choose this sample period because a consistent set of Bloomberg tickers are available over it, and it broadly overlaps with the period over which the U.S. monetary surprises studied in the next section are available. For many emerging markets, we start the sample period later to avoid including periods of pegs or "falling freely" exchange rates, following the classification of Ilzetzki, Reinhart, and Rogoff (2019, 2022). Appendix C describes the start date of all series used.

[19] Three months is the shortest tenor for government bond yields consistently available from Bloomberg across countries. When we study our quantitative model in the next section, we will calibrate it to a weekly frequency which we view as a more appropriate frequency at which active market participants make trading decisions. We will assume the annualized yield on a bond with maturity one week equals that on a bond with three months in the data.

	CHF, EUR, JPY	Other non-U.S. G10	EM	Sum
$\Gamma = 1$	-4.8	-2.2	13.2	6.2
$\Gamma = 1$, G10 avg as base	-4.5	-1.6	13.2	7.1
$\Gamma = 1$, G10 IBOR yields	-5.6	-0.6	9.8	3.7
$\Gamma = 1$, no 10-year bonds	-1.5	-3.5	14.0	9.0
$\Gamma = 1$, 2008-2024	-6.8	-2.4	15.4	6.2
$\Gamma = 1$, $0.5pp$ higher dollar interest rate	-5.6	-2.4	9.1	1.1

Table 1: mean-variance efficient portfolios

Notes: each entry reports total share of wealth invested in those currencies (both short- and long-term bonds). Last column is the total share of wealth invested in non-dollar currencies.

for the G10 economies as well. Since Bloomberg reports government par yields, we bootstrap zero coupon yields using this data.[20]

We use this historical data to construct estimates of $\boldsymbol{ex}$ and $\boldsymbol{C}$. We estimate $\{ex_j\}$ by computing average quarterly excess returns for each currency, running a cross-sectional regression of these on the average three-month interest rate differential versus the dollar, and finally using the fitted values as our estimates. We estimate $\{ex_j^\delta\}$ by computing average quarterly excess returns on ten-year bonds in each currency financed by the three-month dollar bond,[21] running a cross-sectional regression of these on the average yield differential between ten-year bonds in that currency and the three-month dollar interest rate, and using the fitted values as our estimates. We use these fitted values from cross-sectional regressions to reduce the noise in our estimates of excess returns; appendix C demonstrates that the average interest rate differentials and yield curve slopes do a very good job explaining the cross-section of average excess returns. We finally estimate $\boldsymbol{C}$ by computing covariances of log daily changes in exchange rates and log daily changes in ten-year bond prices.[22]

[20] We assume that all countries pay coupons semiannually except for Germany, Norway, Sweden, Switzerland, and Poland, which we assume pay annually, based on our understanding of typical practice for sovereign bonds in each country.

[21] We estimate the quarterly return on a ten-year bond purchased at quarter t and sold at quarter $t+1$ assuming that the yield on a bond with maturity nine years and nine months at time $t+1$ is the same as the yield on a bond with maturity ten years at that same time.

[22] We use the subsample beginning in July 1998 to estimate the covariance matrix, so that we at

Using the estimated $\boldsymbol{ex}$ and $\boldsymbol{C}$, the first row of Table 1 reports that a mean-variance efficient investor with $\Gamma = 1$ would have been on average short G10 currencies, long emerging market currencies, and net long all non-dollar currencies over the sample period. Quantitatively, the investor would borrow 5.2 times their wealth in dollars; borrow 4.8 times their wealth in Swiss franc, euros, and yen (so 1.6 times their wealth in each); borrow an additional 2.2 times their wealth in other G10 currencies; and invest 13.2 times their wealth in emerging markets. The portfolio shares in each currency include three-month and ten-year bonds. In total, the investor would hold a leveraged position of 6.2 times their wealth in non-dollar currencies.[23] As we make more precise below, the fact that the efficient portfolio features a large funding position in dollars even though dollar interest rates are not the lowest among the currencies in our sample reflects that excess returns on a dollar-funded position are high relative to their risk. We will later describe how the relatively small funding position in non-dollar "safe" currencies allows us to explain why U.S. monetary policy has larger effects on risk pricing than monetary policy abroad.

The case with unitary risk aversion is useful because, following Solnik (1974) and Sercu (1980), with conditionally lognormal excess returns the mean-variance efficient portfolio is invariant to the base currency of the investor.[24] In practice, of course, excess returns are not conditionally lognormally distributed, so the base currency could matter. However, the second row reports the average of portfolio shares using each of the possible G10 currencies as the base currency, and demonstrates that even in this case the mean-variance efficient portfolio is largely short the dollar. Away from unitary risk aversion, the base currency would matter even with conditionally lognormal excess returns, because it governs the currency in which the investor further leverages up/down the portfolio reported for $\Gamma = 1$. Our assumption that the representative

least have a balanced panel for the G10 economies. To account for measurement error, we regularize this covariance matrix following Ledoit and Wolf (2004). In particular, the target preserves the sample variances exactly and replaces each off-diagonal entry with the average sample correlation pooled within one of four blocks: within currencies, within 10-year bonds, between a currency and the same-currency 10-year bond, and between a currency and a different-currency 10-year bond. The closed-form Frobenius-MSE-minimizing shrinkage intensity of Ledoit and Wolf (2004) is small, and the shrunken estimator lies essentially on top of the sample covariance.

[23] While it is not reported in the table, we also note that the investor would hold a position of 3.9 times their wealth in the ten-year dollar bond. As described in the last section, this means that in response to a U.S. tightening which erodes the value of long-term dollar bonds, such an investor would take additional balance sheet losses.

[24] This is closely related to the result in Coeurdacier and Gourinchas (2016) that a log investor does not have a real exchange rate hedging motive.

arbitrageur cares about payoffs in dollars is effectively an assumption that the risk aversion of dollar-based arbitrageurs is lower than that of other arbitrageurs.

The next three rows of Table 1 demonstrate that the investor's large funding position in dollars is robust to several features of the underlying data. We first use IBOR-based yields rather than government bond yields for the G10 economies, which may be more appropriate if private investors cannot take large short positions in G10 government bonds at the prevailing yields. The investor still takes a meaningful short position in dollars. This illustrates that it is excess returns relative to the dollar interest rate, not just the U.S. Treasury bill rate, that are high relative to their risk. We next drop ten-year bonds from the menu of assets available to the investor or focus on the second half of our sample period alone to investigate stability over time. In both cases, we find that the investor still holds a leveraged position in emerging market currencies funded largely in dollars. Appendix C similarly demonstrates that this portfolio is robust to using realized average excess returns for $\boldsymbol{ex}$ rather than cross-sectional projections of these average excess returns on interest rate differentials and yield curve slopes, and using a factor model to construct $\boldsymbol{C}$.

The last row of Table 1 finally contemplates a counterfactual world in which the U.S. dollar interest rate is higher relative to all other interest rates. In particular, we reduce each entry of $\boldsymbol{ex}$ by $0.5pp$ (annualized), we keep the covariances in $\boldsymbol{C}$ unchanged, and we continue to assume $\Gamma = 1$. Relative to the first or second rows, the portfolio share in non-dollar bonds now meaningfully falls. This counterfactual scenario is of interest when contemplating the consequences of a potential rise in dollar interest rates in the future, as driven by fiscal pressures or a shift of households away from dollar-denominated assets. We will return to it in section 5.4.

What makes the dollar special? To more formally understand why the dollar is the dominant funding currency in the efficient portfolio, it is useful to focus on the case with another currency as the base currency of the investor. For simplicity, let us focus on short-term bonds alone. Let $ex^k_{j,t}$ denote excess currency returns on short-term bonds in currency j relative to short-term bonds in base currency k. Project $ex^k_{j,t}$ on all other excess returns in a linear regression

$$ex^k_{j,t} = \alpha^k_j + \boldsymbol{\beta}^k_j \boldsymbol{ex}^k_{-j,t} + \varepsilon^k_{j,t},$$

where $\boldsymbol{ex}^k_{-j,t}$ is a column vector of these other excess returns and $\boldsymbol{\beta}^k_j$ is a row vector describing the weights in the portfolio that best replicates $ex^k_{j,t}$. Then standard results for the mean-variance efficient portfolio (e.g., Stevens (1998)) imply that a base-k investor with risk aversion Γ chooses a portfolio share in currency j of

$$\frac{1}{\Gamma}\frac{\alpha^k_j}{\left(\sigma^k_{\varepsilon,j}\right)^2},$$

where $\left(\sigma^k_{\varepsilon,j}\right)^2$ is the variance of $\epsilon^k_{j,t}$. That is, the efficient portfolio share in currency j is increasing in the unspanned excess return on currency j, and falling in magnitude with the unspanned variance of currency j.

The dollar is the dominant funding currency in the efficient portfolio both because its return is low relative to its replicating portfolio, and because its residual risk is low (that is, its payoff is highly replicable with those other currencies). Using each of the G10 currencies as a possible base currency, we compute α^k_j and $\sigma^k_{\varepsilon,j}$ for $j \neq k$. We average over these for each of the G10 base currencies. Figure 1 depicts these average values for each j, annualized. Short-term bonds in the dollar have the second lowest return relative to their replicating portfolio, and the lowest unspanned variance.[25,26] As a consequence, even though dollar rates are not the lowest over the sample, the dollar is the preferred funding currency for mean-variance investors.

Who are the arbitrageurs? We characterize mean-variance efficient portfolios to circumvent the challenges involved in directly measuring arbitrageur exposures in the data. That said, having estimated exposures using our approach, it is a fair question who in the data is actually long emerging market currencies, financed in the dollar.

Our view is that this portfolio describes well the balance sheet in many emerging market economies in which firms and governments borrow in dollars and earn income or revenues in local currency. Our interpretation is that these markets are not isolated, but rather connected through the risk-bearing capacity of the global financial system, which borrows funds from global savers (largely in dollars) to lend to these emerging

[25]This is consistent with foreign central banks disproportionately stabilizing their exchange rates relative to the U.S. dollar (Ilzetzki et al. (2019), Hassan, Mertens, and Zhang (2023)), but foreign interest rates on average still being higher than dollar rates.

[26]A very similar picture is obtained using IBOR yields for the G10, rather than government bond yields. This underscores that Figure 1 speaks to dollar specialness, not Treasury bill specialness.

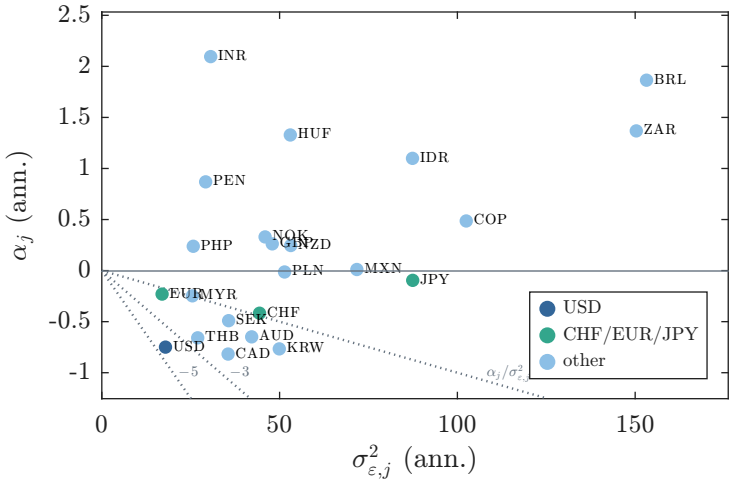


Figure 1: unspanned excess return and variance of short-term bonds in each currency

Notes: using each of G10 currencies as possible base currency k, we compute α_j^k and $\sigma_{\varepsilon,j}^k$ for $j \neq k$ in spanning regression $ex_{j,t}^k = \alpha_j^k + \boldsymbol{\beta}_j^k \boldsymbol{ex}_{-j,t}^k + \varepsilon_{j,t}^k$, where $\left(\sigma_{\varepsilon,j}^k\right)^2$ is variance of $\varepsilon_{j,t}^k$. Marker for currency j depicts average α_j^k and $\sigma_{\varepsilon,j}^k$ over each of G10 base currencies k. Dotted lines depict values with constant $\alpha_j/\sigma_{\varepsilon,j}^2$.

markets.[27] While specific investors such as dealer banks may be largely hedged,[28,29] these results suggest that the system as a whole is not.

4.2 U.S. monetary surprises and global asset prices

We now estimate the cross-section of exchange rate and global yield curve responses to identified U.S. monetary shocks. This evidence is useful both in discriminating between channels of transmission to asset prices, and in disciplining the quantitative model in the next section.

In the data, we follow Bauer and Swanson (2023) in measuring surprises to U.S. nominal interest rates. This is the first principal component of 30-minute changes

[27] We also do not take a stance on the motives of the savers which seek to hold dollar assets and fund arbitrageurs (in the language of the model, we do not microfound foreign households' inelastic positions in dollar bonds). See, for instance, Bocola and Lorenzoni (2020) and Jiang et al. (2024) for perspectives emphasizing insurance and dollar convenience yields, respectively.

[28] For instance, Barbiero, Brauning, Joaquim, and Stein (2026) and Du and Huber (2026) demonstrate that risk limits constrain the exposure of U.S. dealers to currency movements.

[29] As we clarify in an extension with arbitrageurs facing a value-at-risk constraint in section 5, such a constraint might only amplify the asset pricing responses to shocks provided such investors hold a similar portfolio (even if less levered) as mean-variance investors.

in the nearest four quarterly Eurodollar contracts around U.S. monetary policy announcements.[30] We project two-day changes in global yields and exchange rates on the two-day change in the two-year dollar yield around FOMC announcements, instrumented by the intraday monetary surprise. Two-day changes are computed as the closing asset price on the day after the U.S. monetary announcement, less the closing asset price on the day before the announcement. We use an instrumental variables design only to ease interpretation: our estimates correspond to the effect of a $1pp$ increase in the two-year dollar yield induced by the monetary surprise.

Figure 2 suggests that an increase in the price of risk, as due to lower arbitrageur wealth, drives the response of exchange rates and global yield curves to a U.S. monetary tightening. In the first row, we compare the exchange rate responses to our estimated average excess currency returns, ex_j. In the second row, we compare the 10-year yield responses to our estimated average local currency bond carry returns, $ex_j^\delta - ex_j$.[31] The left panels report responses for the G10 economies, while the right panels report them for emerging markets. Consistent with the wealth revaluation channel described in Propositions 1 and 2, currencies earning higher average excess returns depreciate by more, and long-dated yields for currencies with higher average excess bond carry returns rise by more, upon a U.S. tightening. This pattern in exchange rate responses for G10 economies echoes Mueller et al. (2017); to our knowledge, this pattern in exchange rate responses for emerging markets, and these patterns in global yield curve responses for both sets of economies, are novel.

Appendix C demonstrates that it is difficult to explain the response of these asset prices through the portfolio balance mechanism. As demonstrated in Propositions 1 and 2, the portfolio balance channel would imply that currencies (long-term bond prices) that covary more with the trade-weighted average currency carry trade return

[30] We use the version of this measure orthogonalized with respect to macro news released prior to the monetary announcement. Bauer and Swanson (2023) argue that monetary surprises are predictable by this news because markets are learning about the central bank's policy response to this news from monetary announcements. They argue that orthogonalizing the monetary surprise is not needed to study the high-frequency response of asset prices, but we still do so to sharpen interpretation and compare to the model: it suggests that the asset price effects we estimate are not driven by learning about the policy rule, but are driven by surprise innovations around the rule.

[31] Appendix C demonstrates that similar results are again obtained using realized average excess returns, rather than our projections on interest rate differentials and yield curve slopes, for ex_j and $ex_j^\delta - ex_j$. Similar results are also obtained replacing the 10-year yield with five-year ahead, five-year forward rates (validating our interpretation that the response of 10-year yields is driven by a change in global term premia), and excluding all announcements from July 2008 through June 2009 (clarifying that these results are not specific to the global financial crisis).

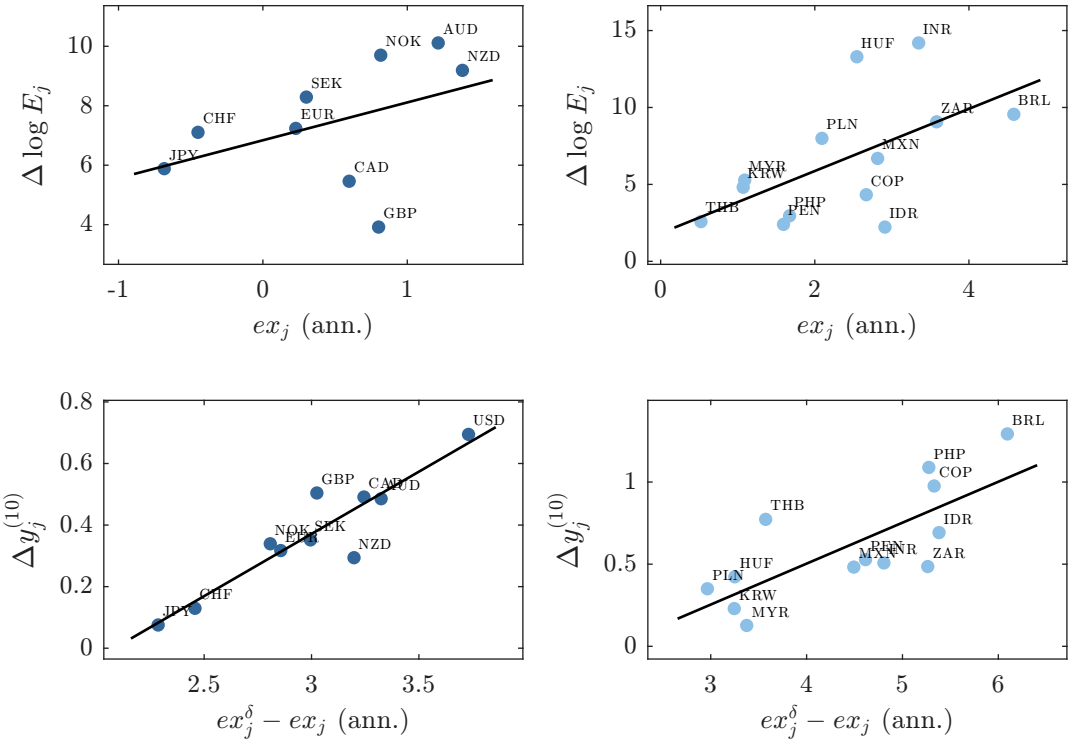


Figure 2: asset price responses to U.S. monetary tightening in data

Notes: response to U.S. monetary tightening estimated in data using Bauer and Swanson (2023) measure of monetary surprises and two-day changes in asset prices around surprises. Responses reported for surprise that raises two-year dollar yield by 1*pp*.

should depreciate by less (fall by less) following a U.S. tightening, assuming capital inflows to the U.S. This prediction does not hold in the data. For instance, currencies with a higher dollar beta depreciate by *more*.[32,33] Portfolio balance effects from other potential capital flows on impact of a U.S. tightening, such as an increase in foreigners' desired position in dollar bonds $\bar{b}_{j1}$ following a rise in the U.S. nominal interest rate, face similar challenges in accounting for the data.

These results on the transmission of interest rate surprises stand in contrast to the transmission of large-scale asset purchases (LSAPs), which *are* well explained

[32]If a U.S. tightening instead implies capital outflows from the U.S., then the predictions of portfolio balance would hold for currencies and emerging market bonds, but still not for G10 bonds.

[33]This approach of comparing asset price responses against (conditional) covariances of asset returns builds on the approach in Bauer and Neely (2014). Our analysis contrasts the portfolio balance channel with one operating through the price of risk, and is focused on interest rate surprises (though, as described in the text, we also study quantitative tightening surprises).

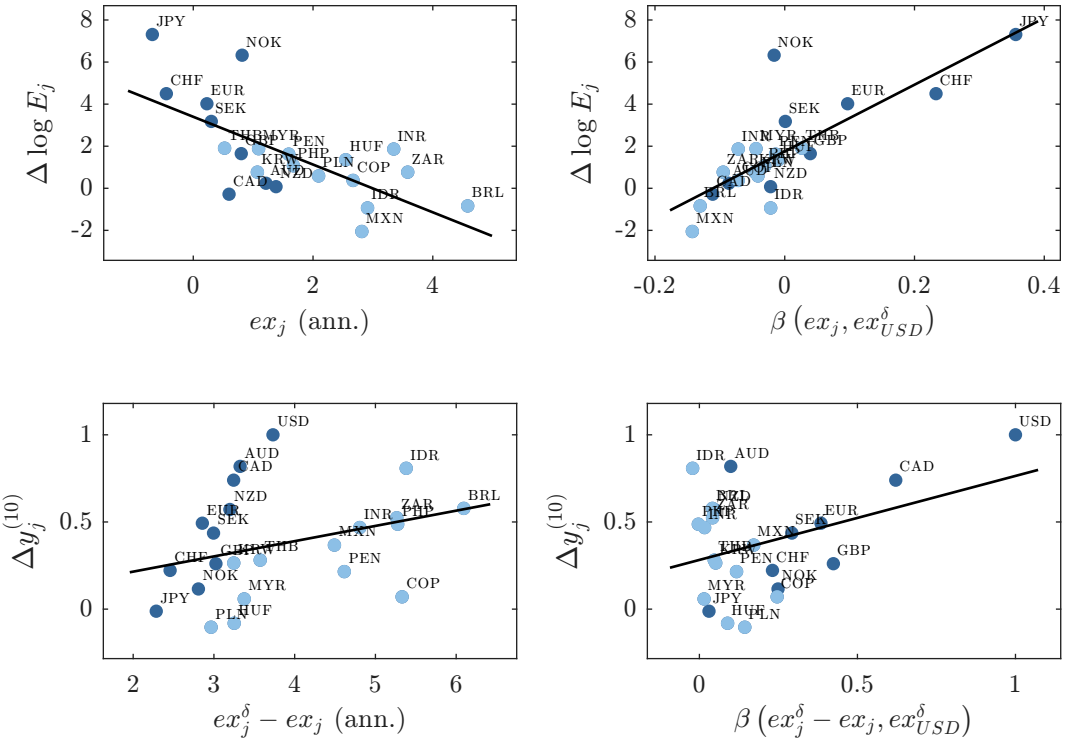


Figure 3: asset price responses to U.S. quantitative tightening in data

Notes: response to U.S. quantitative tightening estimated in data using Swanson (2021) intraday LSAP factor and two-day changes in asset prices around surprises. Responses reported for surprise that raises 10-year yield in U.S. by 1*pp*. $\beta(y,x) \equiv Cov(y,x)/Var(x)$.

by the portfolio balance mechanism. We measure LSAP surprises using Swanson (2021)'s LSAP factor constructed from the high-frequency response of long yields in tight intraday windows around Federal Reserve announcements. We project two-day changes in exchange rates and global yields on two-day changes in the U.S. 10-year yield, instrumented with the intraday Swanson (2021) LSAP surprise. The left panels of Figure 3 depict the effects against average excess returns, while the right panels depict the effects against covariances with the U.S. bond carry return.

The responses to LSAP surprises are consistent with the portfolio balance mechanism. In response to more long-term dollar bonds on arbitrageur balance sheets, currencies with a higher covariance with the U.S. bond carry trade should depreciate by more, and long-term bond prices with a higher covariance with the U.S. bond carry trade should fall by more (and thus long yields should rise by more). The right

panels of Figure 3 demonstrate that this is broadly what we see. The exchange rate responses are particularly striking because currencies with a higher covariance with the U.S. bond carry trade tend to pay *lower* excess returns on average.[34] Thus, a higher price of risk, as in the case of an interest rate surprise, would generate the opposite patterns in exchange rate responses than in Figure 3.

An interpretation is that, while U.S. quantitative tightening revalues downward the wealth of arbitrageurs like a U.S. interest rate hike (both due to the stronger dollar and lower long-term dollar bond prices), its main effects on global risk premia operate through the portfolio balance mechanism.[35] This result both validates the asset substitution elasticities implied by mean-variance optimization, and sharpens our finding that interest rate policy spills over through a distinct mechanism.

5 Quantification

We now discipline the model to rationalize the high-frequency effects of U.S. interest rate policy on international asset prices. We show that arbitrageurs' dollar funding is central to explaining these effects. The risk premium spillovers of U.S. monetary policy imply a stagflationary trade-off for foreign central banks, particularly in emerging markets, which can in part be managed through FX intervention. In a potential future with relatively higher dollar interest rates, these spillovers of U.S. monetary policy are dampened.

5.1 Model parameterization

We solve the model at a weekly frequency. We set parameters in three steps.

We first set a number of parameters without solving the model, summarized in Table 2. We set households' elasticity of intertemporal substitution ψ to 0.5, the trade elasticity ς to 1.5, and the Frisch elasticity of labor supply φ to 1, all standard values in the literature.[36] We set the elasticity of substitution across product varieties ϵ to

[34] Intuitively, when interest rates fall globally, they tend to fall least in low interest rate countries like Japan and Switzerland, so these currencies appreciate as long-term (U.S.) bonds pay well.

[35] This result is also consistent with the recent evidence of Krohn, Uthemann, Vala, and Yang (2026), who study global asset price responses around U.S. Treasury auctions.

[36] See, for instance, Barsky, Juster, Kimball, and Shapiro (1997), Backus, Kehoe, and Kydland (1994), and Chetty, Guren, Manoli, and Weber (2011), respectively.

	Description	Value
ψ	elast. of intertemporal subs.	0.5
ς	trade elasticity	1.5
φ	Frisch elasticity	1
ϵ	elast. of subs. across varieties	11
χ_p	price adj. cost	26,520
ϕ_π	interest rate coeff. on CPI inflation	1.5
ρ_i	interest rate inertia	$0.4^{1/52}$
δ	consol decay	duration 10 years
$Cov(ex_{j,+1}, ex_{k,+1})$	conditional cov. of carry returns	see text

Table 2: baseline parameters set without solving model

Notes: model is solved at a weekly frequency.

11 and the price adjustment cost χ_p to $26,520$; together, these imply a Calvo (1983)-equivalent frequency of price adjustment around one year. We set the monetary policy coefficient on inflation to the standard value $\phi_\pi = 1.5$, and we set interest rate inertia ρ_i to 0.4 annualized, which we assume for simplicity are identical everywhere. We set δ so that the duration of consols is approximately 10 years. Finally, we set the conditional covariance of excess returns to exactly match our baseline estimates of the covariance matrix $\boldsymbol{C}$ described in the prior subsection. As previously noted, it is only through this covariance matrix that the presence and stochastic properties of other shocks affect the propagation of a U.S. monetary policy shock. Our implicit assumption is thus that there exists a set of shocks in this environment that can rationalize the estimated $\boldsymbol{C}$.

The second set of parameters are calibrated after solving the model, but without reference to U.S. monetary surprises. These are reported in the first four rows of Table 3, and all calibrated to targets estimated over the 1991 Q2 – 2024 Q1 period.[37,38] We set the populations ζ_j to target average GDP relative to the U.S. as reported in the IMF national accounts.[39] We set the trade shares η_{jk} to target expenditure shares

[37]Appendix D reports the value of each of these parameters in our baseline calibration, as well as the model's success in matching the targeted moments closely.

[38]We solve the model for all G10 and emerging market economies used in our empirical analysis, and a "rest of world". Appendix D describes how we set parameters for the "rest of world".

[39]We define GDP in j, in units of local consumption bundle, as $y_j = q_{j,t} \sum_{k=1}^{n} q_{k,t}^{-1}(P_{kj,t}/P_{k,t})y_{kj,t}$.

	Description	Moments	Targets
ζ_j	populations	$q_j^{-1} y_j / y_1$	relative GDP
β_j	discount factors	i_1, ex_j	dollar rate, excess returns
$\bar{b}_{1j}^{\delta}$	consol demand	ex_j^{δ}	excess returns
η_{jk}	trade shares	$P_{jk} c_{jk} / (P_j c_j)$	expend. shares by origin
θ_{jk}^m	curr. invoicing shares	$\sum_k, \sum_j \theta_{jk}^m \zeta_j E_j^{-1} P_{jk} c_{jk}$	trade by invoicing curr.
Γ	arb. risk aversion	resp. to U.S. monetary surprise (see text)	
ξ	arb. exit rate	resp. to U.S. monetary surprise (see text)	
$\rho_{\nu i 1}$	persist. U.S. MP shock	resp. to U.S. monetary surprise (see text)	
ω_a	arb. endowment	resp. to U.S. QE surprise (see text)	

Table 3: baseline parameters set by solving model

Notes: data moments estimated over 1991 Q2 – 2024 Q1. Disutility of labor $\bar{\varphi}$ and Taylor rule intercepts i_j are set to normalize $\ell_1 = 1$ and $\pi_j = 0$. For simplicity $\{\bar{b}_{1j} = 0\}$ and $\{\bar{b}_{jk} = 0, \bar{b}_{jk}^{\delta} = 0\}$ for $j > 1$.

using the OECD's Trade in Value Added statistics. We set currency invoicing shares θ_{jk}^m using the evidence compiled by Boz, Bruggen, Casas, Georgiadis, Gopinath, and Mehl (2025). This involves projecting invoicing shares for each destination, origin, and currency onto fixed effects for each destination-currency and origin-currency pair, since the available data only provides invoicing shares for imports in and exports from each country (not each trading pair). Finally, we set the discount factors β_j and household positions in foreign bonds and consols $\{\bar{b}_{jk}, \bar{b}_{jk}^{\delta}\}$ to target the expected excess returns $\boldsymbol{ex}$ estimated in the prior subsection, as well as the average three-month dollar interest rate from Bloomberg.[40]

There are (infinitely) many ways to set households' exogenous positions in foreign bonds and consols $\{\bar{b}_{jk}, \bar{b}_{jk}^{\delta}\}$ so that asset markets clear, given the efficient portfolio chosen by arbitrageurs. As a clean benchmark, our baseline approach implies zero net foreign asset positions: we assume U.S. households take the other side of arbitrageurs' non-dollar and consol exposures. Appendix D reports the results of a richer calibration matching observed net foreign asset positions. In this calibration, we assume that households in each non-U.S. country take the other side of arbitrageurs' positions in that currency, and we use these households' position in one-period dollar bonds to

[40] As described in footnote 19, we assume the annualized yield on a bond with maturity one week in the model is identical to that on a bond with three months in the data.

match the observed NFAs. To rationalize the fact that arbitrageurs are short the dollar even though the rest of the world has a positive net foreign asset position vis-à-vis the U.S., this calibration requires that the rest of the world holds dollar-denominated assets in excess of their borrowing in local currencies. This is consistent with estimates of the currency composition of NFAs (e.g., Allen and Juvenal (2025)).

The remaining parameters are calibrated after solving the model, and targeting the high-frequency asset price responses to U.S. policy surprises. These are arbitrageur risk aversion Γ, the exit rate of arbitrageurs ξ, the persistence of the monetary shock ρ_1, and the initial endowment of arbitrageurs ω_a.

We discipline the first three parameters to match the high-frequency evidence from U.S. interest rate surprises studied in the last section. Since the model matches average excess returns and conditional covariances of excess returns in the data, arbitrageurs' risk aversion Γ is the only remaining degree of freedom we have to pin down their portfolio, as it governs their leverage (recall the discussion in section 4.1). Consistent with Propositions 1 and 2, arbitrageurs' leverage, ξ, and ρ_1 shape the asset price responses to U.S. monetary surprises, and can thus be identified by them.[41]

Formally, we calibrate these parameters to minimize the deviations in exchange rate, two-year yield, and 10-year yield responses between data and model,[42] simulating the U.S. monetary surprise as a positive innovation to $\nu_{1,t}$ which raises the two-year dollar yield by $1pp$ as in the data. We target moments for the G10 economies alone, given that emerging markets may have richer policy reactions functions than the inflation targeting rule we have assumed (we return to this below).

Conditional on these parameters, we then discipline arbitrageurs' initial endowment ω_a to match the response of U.S. yields to quantitative easing (QE) announcements. Intuitively, conditional on Γ, arbitrageurs' steady-state wealth governs the price impact of a dollar change in their bond positions, and is thus well-suited to match evidence from QE. Roughly matching the Federal Reserve's purchase of 10-year equivalent bonds in QE1, we simulate a positive U.S. household consol demand shock ($\Delta\bar{b}^{\delta}_{11,t}$) that is 6.5% of annual U.S. GDP, and with half-life of roughly four years.[43] We discipline ω_a to match a $1.05pp$ decline in the 10-year dollar yield on

[41] Appendix D depicts how Γ and ξ are identified in particular.

[42] We obtain two-year yields in the model by pricing consols with duration of two years, even though arbitrageurs only have non-zero positions in the consols with duration 10 years.

[43] Gagnon, Raskin, Remache, and Sack (2011) report purchases of 10-year equivalents of \$850bn between December 2008 and March 2010. To adjust for the fact that the resulting price impact may

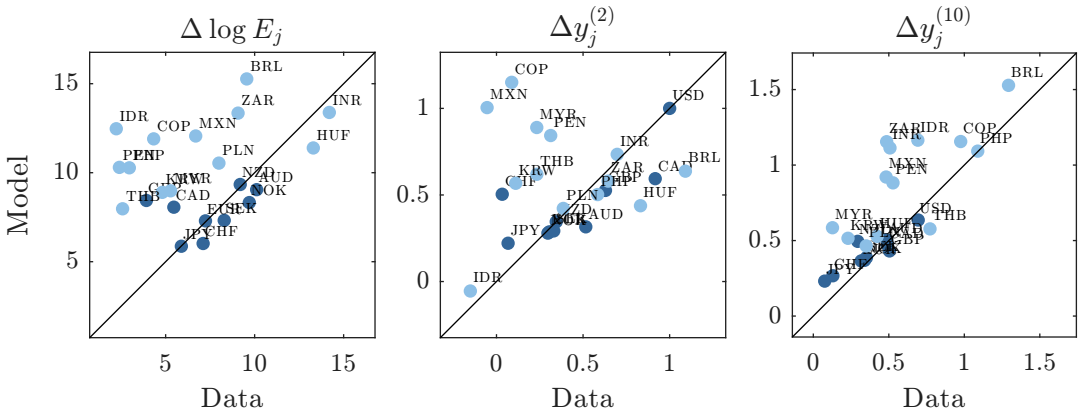


Figure 4: asset price responses to U.S. monetary tightening in model vs. data

Notes: both model and data responses reported for surprise that raises two-year dollar yield by 1*pp*.

impact of the shock.[44]

Figure 4 demonstrates that the model is quite successful in matching the asset price responses to U.S. interest rate surprises in the data. While these moments are targeted for the G10 economies, the success for these economies is still notable given that the estimation is substantially overidentified (only three parameters are being chosen to match 28 moments). The model is even fairly successful for emerging markets, which were not at all used in the GMM. The discrepancies that do appear for emerging markets may reflect that we are especially mismeasuring risk premia for these economies, or that these economies engage in exchange rate intervention following a U.S. tightening. We further discuss the latter possibility below.

The estimated Γ implies that arbitrageurs' leverage in non-dollar bonds

$$\frac{\sum_{k=2}^{n} q_k^{-1}(b_{ak} + b_{ak}^{\delta})}{w_a} \approx 18.$$

The model requires a high degree of arbitrageur leverage to rationalize the data because arbitrageurs endogenously earn high returns in the periods following a U.S. tightening, dampening the initial response of asset prices to the shock. Indeed, ap-

have been larger than in non-crisis times, we round this upwards to \$1tn. Relative to annual GDP of roughly \$15tn, this is approximately 6.5%.

[44]This is the cumulative change in 10-year dollar yield on 11/25/2008, 12/1/2008, 12/16/2008, 1/28/2009, and 3/18/2009, the first five U.S. QE "event days" widely studied in the literature. We cumulate the change in yields from the day after the event, relative to the day before the event.

pendix D demonstrates that arbitrageur wealth is largely back at its steady-state value one year after a U.S. tightening. This reflects that expected excess returns are high and that arbitrageurs take very large leverage after their wealth is depleted.

This motivates an extension of our baseline model with a second set of arbitrageurs who choose the same steady-state portfolio, but who must *shrink* their assets when wealth declines. In particular, suppose there is a second set of arbitrageurs — "value-at-risk arbitrageurs" — who are risk-neutral and choose their portfolio to maximize

$$E_t \Pi_{a_2,t+1} \tag{9}$$

subject to the value-at-risk constraint

$$Var \Pi_{a_2,t+1} \leq \kappa W_{a_2,t}^2. \tag{10}$$

Their period t budget constraint, $t+1$ profits, and evolution of wealth are analogous to those of the mean-variance arbitrageurs in (1)-(3). Appendix D demonstrates that, under a condition relating κ to Γ, the steady-state portfolio chosen by the value-at-risk arbitrageurs is identical to that of the mean-variance arbitrageurs. The key difference between these arbitrageurs is the dynamic response of portfolios to shocks: in response to a U.S. tightening that erodes their wealth, value-at-risk arbitrageurs must shed assets to satisfy the constraint (10).[45]

With the addition of these value-at-risk arbitrageurs, we can match the same responses to a U.S. monetary tightening with a much smaller degree of arbitrageur leverage. In particular, since value-at-risk arbitrageurs will shrink their balance sheets, mean-variance arbitrageurs must expand theirs more than they otherwise would, amplifying their increase in required returns and thus the initial response in asset prices. The strength of this mechanism is increasing in the relative size of value-at-risk arbitrageurs' balance sheets. For alternative values of the steady-state wealth share of value-at-risk arbitrageurs w_{a_2}/w_a (achieved with the relative endowments of the arbitrageurs), Figure 5 depicts the estimated leverage in non-dollar bonds $\frac{\sum_{k=2}^{n} q_k^{-1}(b_{ak}+b_{ak}^{\delta})}{w_a}$ required to best match the data in Figure 4. For instance, when 95% of arbitrageur wealth is held by value-at-risk arbitrageurs, arbitrageurs only require leverage of 2 in

[45] In this sense, the value-at-risk constraint is more consistent with the procyclical leverage of global banks documented by Adrian and Shin (2010), Bruno and Shin (2015a), and Miranda-Agrippino and Rey (2020).

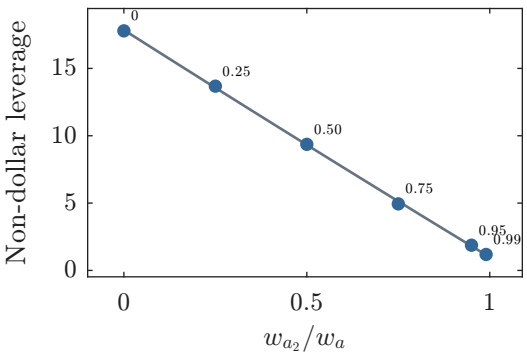


Figure 5: estimated arbitrageur leverage given two types of arbitrageurs

Notes: second type of arbitrageur is risk neutral but faces value-at-risk constraint (10), with κ chosen to achieve same portfolio of mean-variance arbitrageur in steady-state (see appendix D for details). x-axis varies relative endowments of arbitrageurs and thus steady-state wealth share of value-at-risk arbitrageur, and y-axis reports estimated leverage of arbitrageurs in non-dollar currencies, $\frac{\sum_{k=2}^{n} q_k^{-1}(b_{ak}+b_{ak}^{\delta})}{w_a}$, best matching data in Figure 4.

non-dollar bonds to match the same asset price data on impact of a U.S. tightening.

Having matched the asset price responses to U.S. monetary policy shocks, we now use our model to study several questions of interest. We do so using our baseline calibration, but note that similar results are obtained using the alternative calibrations matching NFAs and with value-at-risk arbitrageurs also described in this section.

5.2 Effects of U.S. monetary policy on global asset prices

We first use the model to decompose the asset pricing effects of U.S. monetary policy, and to clarify what makes U.S. monetary policy special.

The model implies that the change in the global price of risk is the primary driver of heterogeneity in asset price responses to a U.S. tightening. Figure 6 compares exchange rate and yield curve responses in the calibrated model versus a counterfactual environment in which average excess returns across currencies and bonds are zero (because discount factors are equated and household consol positions are all zero) and all other primitives are unchanged. Spillovers through the price of risk are thus shut down, so that the only remaining source of heterogeneous spillovers across countries are heterogeneous trade linkages, country sizes, invoicing patterns, and return covariances. While the latter (particularly invoicing patterns) do account for some of the heterogeneous responses in short-term interest rates, they account for essentially

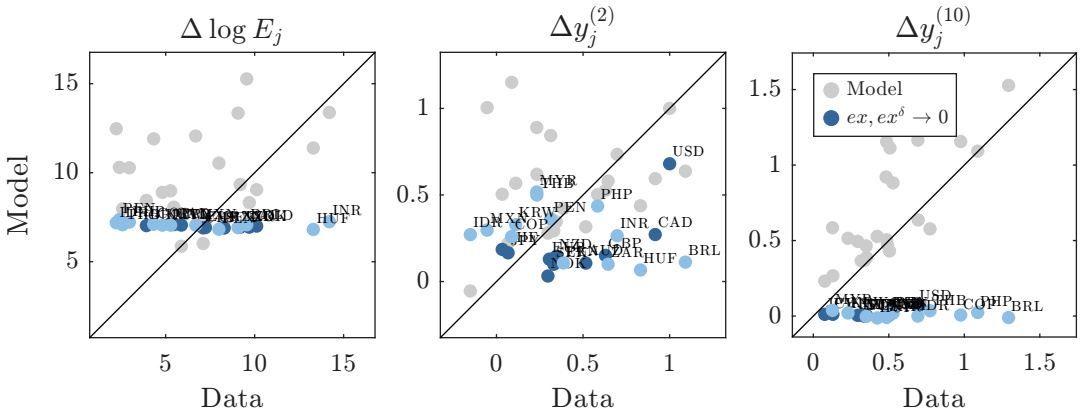


Figure 6: comparing transmission mechanisms of U.S. tightening

Notes: both baseline model and data responses reported for surprise that raises two-year dollar yield by 1*pp*. Size of U.S. monetary shock in counterfactual environment with $\Gamma, \boldsymbol{ex}, \boldsymbol{ex}^{\delta} \to 0$ kept same as baseline model.

none of the exchange rate and long-term yield responses that we see in the data.

The global asset price responses to U.S. monetary surprises are different than their responses to surprises by other central banks. Figure 7 compares the high-frequency responses of exchange rates to Federal Reserve surprises versus European Central Bank (ECB) surprises. We focus on ECB surprises because the Euro area is an economy of comparable size to the U.S. We measure ECB surprises using the intraday "forward guidance" surprise measure from Altavilla, Brugnolini, Gurkaynak, Motto, and Ragusa (2019).[46] In both panels, we consider a surprise that raises the two-year domestic yield by 1*pp*. By comparison to the responses to Federal Reserve surprises in the first panel, we see that the exchange rate effects of ECB surprises are much more weakly related to average excess returns. Through the lens of Proposition 1, this is consistent with weaker effects of the ECB surprise on global risk pricing.[47]

The model clarifies that the dollar's dominant role as a funding currency is es-

[46]This explains most variation in OIS rates around the monetary announcement with maturity between two and five years, though similar results are obtained for the "timing factor" that explains most variation in OIS rates with maturity less than or equal to one year. We use the surprise measures for the event window that encapsulates both the press release and press conference. The surprises begin in January 1999 and we study all such surprises through December 2019, for comparability with the sample of Federal Reserve surprises studied in the last section.

[47]While other researchers have compared spillovers induced by the Federal Reserve and other central banks (e.g., Gerko and Rey (2017) and Ca'Zorzi, Dedola, Georgiadis, Jarocinski, Stracca, and Strasser (2023)), to our knowledge this evidence leveraging the cross-section of international asset price responses is novel.

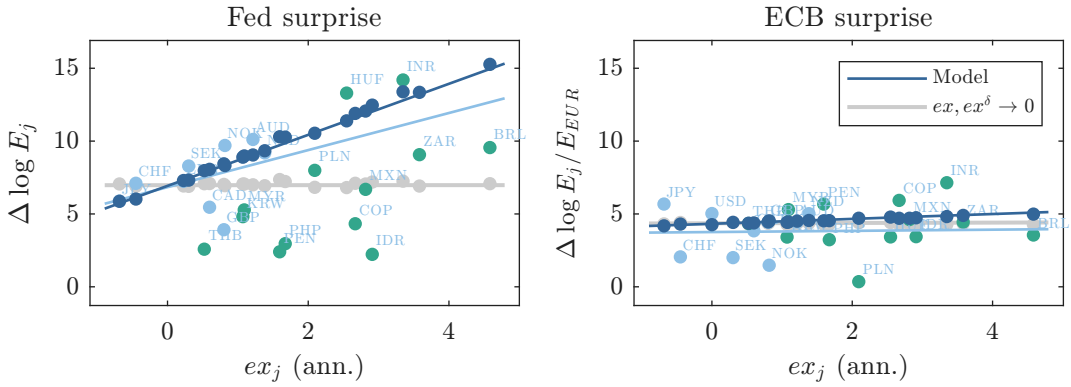


Figure 7: asset price responses to U.S. vs. Euro area monetary tightening

Notes: both model and data responses reported for surprise that raises two-year dollar yield by $1pp$ (left panel), and two-year euro yield by $1pp$ (right panel). The response to Euro area monetary tightening estimated in data using Altavilla et al. (2019) "forward guidance" measure of monetary surprises and two-day changes in asset prices around surprises. Light blue and green markers depict data, dark blue markers depict model (and associated line of best fit), and grey markers depict counterfactual environment with $\boldsymbol{ex}, \boldsymbol{ex}^{\delta} \to 0$ (and associated line of best fit).

sential to explain why U.S. monetary policy is special in this way. Figure 7 first demonstrates that our model generates the observed asymmetry in asset price responses to U.S. and Euro area monetary surprises. We keep all parameters fixed from our simulation of a U.S. monetary surprise in the last section, except that we estimate the persistence of a Euro area monetary surprise $\rho_{i,EUR}$ to match the average exchange rate response relative to the euro. As in the data (and completely untargeted), the model generates an essentially non-existent relationship between the exchange rate responses to a Euro area surprise and average excess currency returns. This follows from arbitrageurs funding only a small fraction of their portfolio in euros, as described in Table 1: while a euro appreciation thus erodes their wealth, the wealth revaluation is an order of magnitude smaller than in response to a dollar appreciation. Indeed, if all average excess returns were zero and thus spillovers through arbitrageurs' wealth revaluation were shut down, Figure 7 demonstrates that the cross-sectional effects of a U.S. and Euro area tightening would essentially be the same.

5.3 Real consequences and policy responses

We now quantify the real effects of U.S. monetary spillovers through risk pricing, and how these can be shaped by alternative policy responses in the rest of the world.

The model implies that the effects of U.S. monetary policy on global risk pricing have first-order consequences for spillovers to inflation and output abroad. Figure 8 compares impulse responses to the same U.S. monetary policy shock simulated in the prior figures. We contrast simple averages of impulse responses for the other G10 economies against those for emerging markets. Even absent a change in the price of risk, the dollar appreciation following a U.S. tightening causes emerging markets to experience a bigger jump in nominal prices and decline in consumption than G10 economies.[48] This reflects the preponderance of dollar pricing for emerging markets, which in turn causes them to tighten policy rates more aggressively than G10 economies. The relatively stagflationary effects of a U.S. tightening on emerging markets are amplified by the change in global risk pricing, since the amplified depreciation in emerging markets passes through into even higher domestic prices and thus policy rates. Quantitatively, recalling that the U.S. tightening is scaled to raise the two-year dollar yield by $1pp$, the endogenous change in risk pricing amplifies the peak decline in consumption in emerging markets by $0.4pp$. These effects of a change in risk pricing on the real economy are meaningful but also fairly transitory, reflecting that the change in risk premia are (endogenously) transitory as well.[49]

Modifying the monetary policy rule in emerging markets to depart from pure inflation targeting, and to instead put weight on stabilizing the exchange rate, can mitigate the depreciation and thus imported inflation in these economies. However, because this requires that these emerging markets further tighten policy rates in response to a U.S. tightening, it comes at the expense of a further decline in consumption. Appendix D quantifies this trade-off for Brazil, a representative large emerging market.

Here we instead focus on an alternative policy response — foreign exchange intervention — which more directly addresses the change in the external risk premium at its source. We model such intervention as the Brazilian government selling one-

[48] After the initial jump in prices, both emerging markets and G10 economies experience persistent deflation owing to the decline in global demand induced by the increase in policy rates. Nonetheless, nominal prices remain permanently higher than in the absence of the U.S. tightening.

[49] Empirical local projections corresponding to the impulse responses in Figure 8 are typically very noisy (indeed, this is why we use the model to infer the real effects of a U.S. tightening, having disciplined it to match the better identified impact effects on asset prices). That said, available estimates validate our model's implied magnitudes — Mendoza-Fernandez and Pelin (2026), for instance, estimate that a $1pp$ surprise increase in the Fed funds rate raises policy rates in emerging markets by $[0.3pp, 0.8pp]$ and advanced economies by $[-0.1pp, 0.3pp]$ in their first meeting after the Fed surprise (90% confidence intervals based on their Table 2). These responses are only a bit below those in Figure 8, recognizing that it depicts a $1.5pp$ increase in the annualized U.S. policy rate.

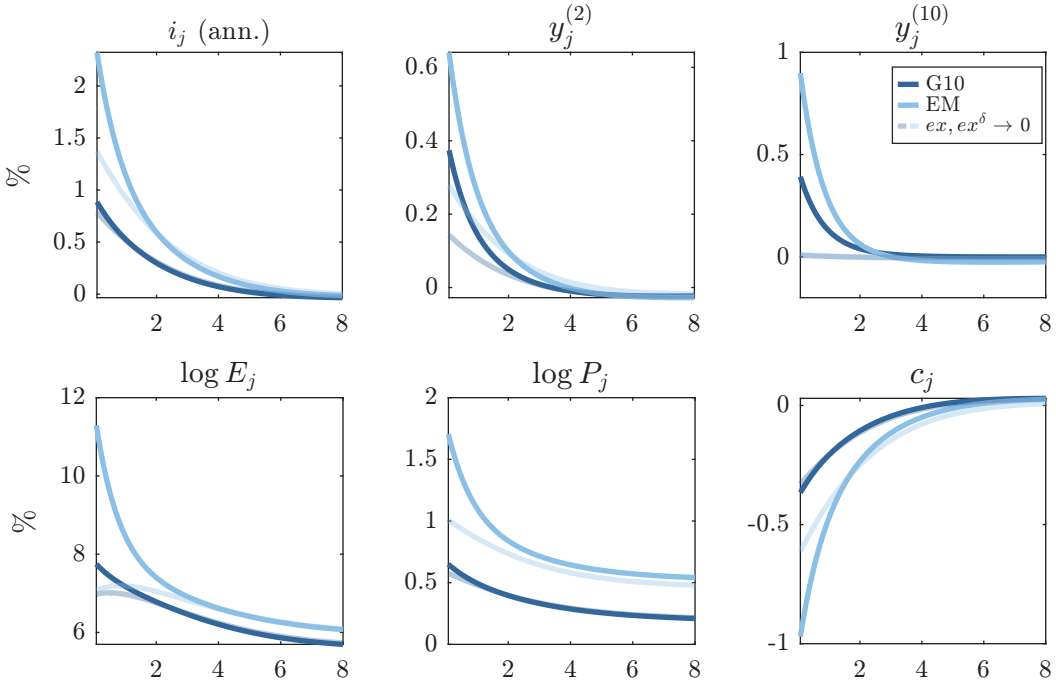


Figure 8: impulse responses to U.S. monetary tightening

Notes: responses reported for same $\nu_{1,t}$ innovation depicted in Figure 4. x-axis denotes quarters since the shock.

period dollar bonds and buying one-period Brazilian bonds. Since Brazil is Ricardian with respect to its local currency bonds (because households can freely trade them), this amounts to the Brazilian government selling one-period dollar bonds (lowering $\bar{b}_{BRL1}$) on behalf of local households. These households in turn seek to smooth consumption by buying local currency bonds. With fewer such bonds on the balance sheet of arbitrageurs, Figure 9 demonstrates that the local currency risk premium is dampened, in turn dampening the Brazilian depreciation, increase in policy rates, and decline in consumption.[50] Moreover, because the excess return on Brazilian real covaries positively with the excess return on many other currencies, and positively with the local currency bond carry trade in Brazil,[51] intervention in Brazil tends to

[50]Interestingly, these effects are more persistent than those caused by the increase in the price of risk upon a U.S. tightening, depicted in Figure 8. The reason is that FX intervention implies that the Brazilian government earns expected returns that would otherwise accrue to foreign arbitrageurs. The associated cross-border wealth transfers imply a persistent increase in consumption and Brazilian appreciation relative to the counterfactual without intervention.

[51]The latter reflects that the real tends to depreciate when Brazilian long-term yields rise. This finding echoes points made in Kekre and Lenel (2024a) and De Leo, Keller, Simoncelli, Villamizar-

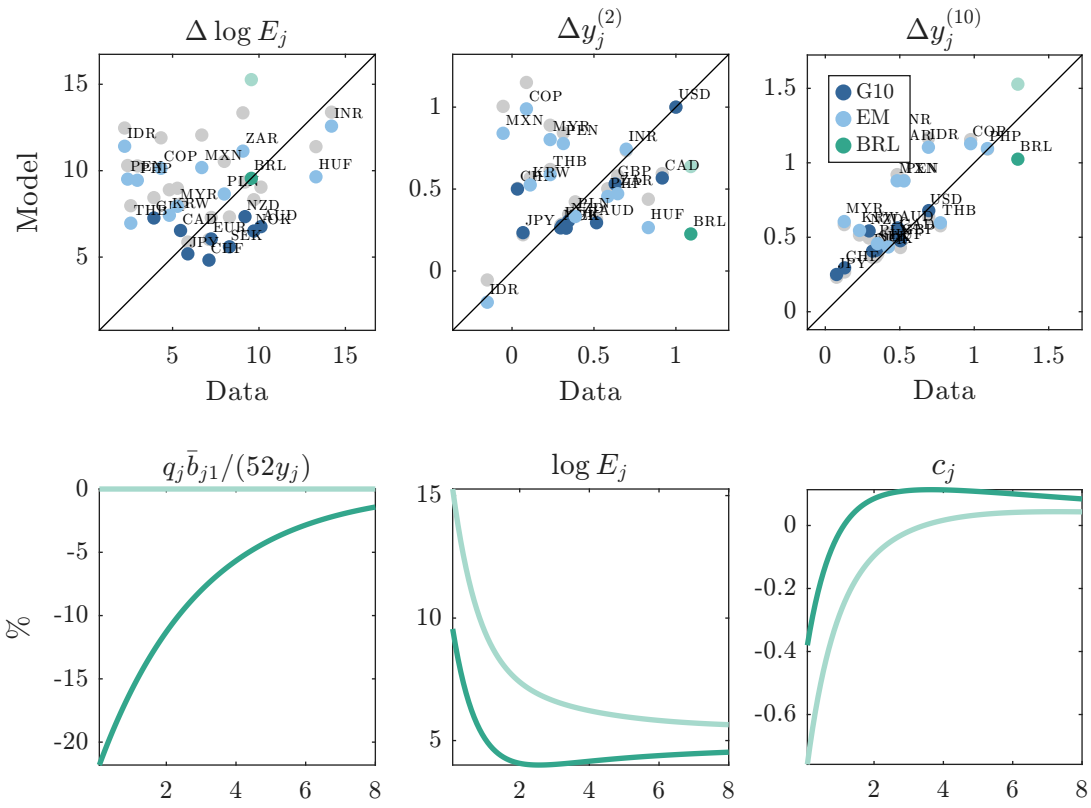


Figure 9: effects of FX intervention in response to U.S. monetary tightening

Notes: responses reported for same $\nu_{1,t}$ innovation depicted in Figure 4. x-axis denotes quarters since the shock.

reduce other currency risk premia and Brazil's term premium as well.

The magnitude of the required intervention to rationalize the data is plausible. Given an annualized persistence of the decline in $\bar{b}_{BRL1}$ of 0.25, Figure 9 demonstrates that a decline in $\bar{b}_{BRL1}$ by 21% of Brazil's annual GDP can rationalize the observed Brazilian depreciation to a U.S. monetary tightening.[52,53] Credible data on exchange rate interventions is thin; our model offers a way to infer (expected) interventions

Villegas, and Williams (2026) for emerging markets more broadly.

[52] As a point of comparison, Milesi-Ferretti (2026) reports that Brazil's reserve balances excluding gold have averaged 15-25% of its annual GDP in recent years. Moreover, recall that the U.S. monetary surprise underlying Figure 9 was scaled to raise the two-year dollar yield by 1*pp*, but typical monetary surprises are an order of magnitude smaller.

[53] Appendix D explores the sensitivity to two key parameters governing the price impact of interventions: the value of arbitrageurs' initial endowment ω_a and thus steady-state wealth (disciplined to match the price impact of QE1), and the persistence of the intervention itself.

based on the responses of asset prices instead. In response to U.S. monetary surprises, interventions of a plausible size may explain why asset prices in emerging markets do not respond by as much as their unconditional excess returns would suggest.

5.4 A future with higher dollar interest rates

We finally study a potential future with relatively higher dollar interest rates.

We model a higher steady-state dollar interest rate in the model via more impatience of U.S. households (lower β_1). It would also arise from lower foreign demand for dollar-denominated assets (lower $\bar{b}_{j1}$ for any j), provided we modify the model to make U.S. households unwilling to elastically supply such assets at the prevailing interest rate.[54] Under a similar condition, increased issuance of dollar-denominated debt by governments would also cause an increase in the dollar interest rate.

Here we consider a decrease in β_1 that raises the steady-state dollar interest rate by $0.5pp$ (annualized). Consistent with the discussion of the last row of Table 1, arbitrageurs take smaller equilibrium positions in emerging market currencies. Quantitatively, keeping all other economic primitives fixed, arbitrageurs' leverage in non-dollar bonds falls from its value of roughly 18 in the baseline model, to 3 here.

The implications for U.S. monetary policy transmission reflect a smaller revaluation of arbitrageur wealth due to dampened currency mismatch. The first row of Figure 10 compares the asset price responses to a U.S. monetary tightening in this counterfactual environment versus the baseline model (from Figure 4). Following a U.S. monetary tightening in this counterfactual environment, the exchange rate depreciation of currencies paying high excess returns is dampened, as is the increase in bond yields of currencies with high excess bond carry returns. The second row demonstrates that these dampened effects on asset prices are in turn reflected in meaningfully smaller spillovers to policy rates and consumption in emerging markets.

[54] This could be achieved by an adjustment cost in the dollar bond position of U.S. households. Note that this is a statement about steady-state — in response to shocks to foreign dollar demand, U.S. households will not perfectly elastically accommodate them because they wish to smooth consumption. This is why FX interventions have price impact in the prior subsection.

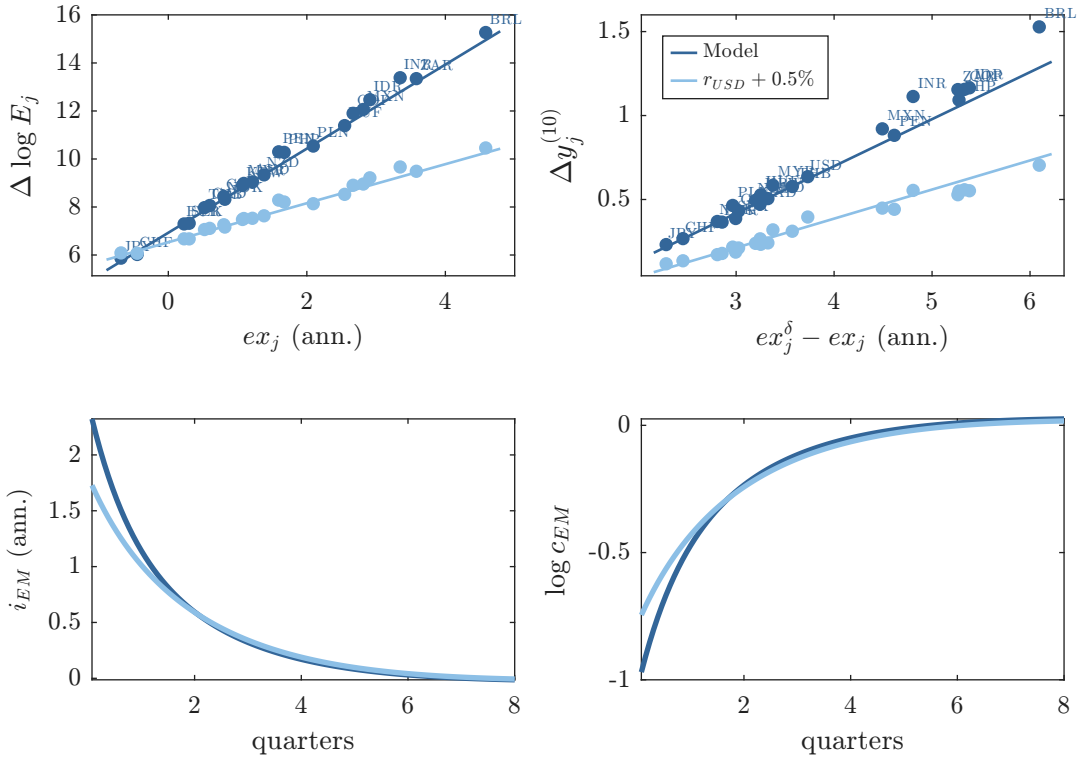


Figure 10: responses to U.S. monetary tightening with 0.5*pp* higher U.S. interest rate

Notes: responses reported for same $\nu_{1,t}$ innovation in Figure 4. x-axis in bottom panels denote quarters since the shock.

6 Conclusion

This paper has proposed a general equilibrium model of U.S. monetary policy spillovers through changes in risk premia in global financial markets. In an otherwise standard multi-country New Keynesian environment, capital flows are facilitated by global arbitrageurs that price the cross section of assets. If arbitrageurs are short the dollar and long foreign currencies, tighter U.S. monetary policy erodes arbitrageurs' net worth and raises the global price of risk. We conclude from observed asset prices that arbitrageur portfolios must indeed have this structure, assuming they are mean-variance efficient. Consistent with an increase in the price of risk, the high-frequency responses to a U.S. tightening line up with average excess returns. We estimate arbitrageurs' leverage to rationalize this evidence, and we quantify the implications for real activity and how these depend on policy responses abroad. In a potential future with higher

dollar interest rates, we find that arbitrageurs are less currency-mismatched and the spillovers of U.S. monetary policy through risk pricing are dampened.

We see two natural directions to extend the analysis in this paper, both of which could take advantage of the tractable linearized framework we propose. First, we have focused on the transmission of U.S. monetary policy, given its central role in discussions of a "global financial cycle." But the mechanisms we emphasize would also apply to any driving force that appreciates the dollar (such as a decline in U.S. output), which would endogenously raise risk premia in our model.[55] Future work could study many such shocks in this model and estimate their relative importance in explaining the comovements between the dollar and global risk pricing.[56] Second, our model can be used to study normative questions. In particular, because the price of risk depends on the broad dollar exchange rate, it may imply novel gains from policy coordination in responding to spillovers induced by U.S. monetary policy (or other shocks). We leave the analysis of these questions to future work.

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[55] While the transmission from the dollar exchange rate to risk pricing is the focus of the present paper, it is also the case that an exogenous shock to arbitrageur risk aversion would appreciate the dollar owing to arbitrageurs' dollar funding, as emphasized in Kekre and Lenel (2024a).

[56] For a description of these comovements see, for instance, Obstfeld and Zhou (2023).

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Appendix for Online Publication

A Equilibrium

In this appendix we outline the equilibrium conditions pertaining to the model described in section 2 and the first-order approximation studied throughout the paper.

A.1 Equilibrium conditions

In addition to the variables defined in the main text, let us define

$$\begin{aligned}
w_{j,t} &\equiv W_{j,t}/P_{j,t}, \\
\pi_{kj,t} &\equiv \log P_{kj,t}/P_{kj,t-1}, \\
p^m_{kj,t} &\equiv E_{k,t}E^{-1}_{m,t}P^m_{kj,t}/P_{k,t}, \\
\pi^m_{kj,t} &\equiv \log P^m_{kj,t}/P^m_{kj,t-1}, \\
b_{jk,t} &\equiv B_{jk,t}/P_{k,t}, \\
b^\delta_{jk,t} &\equiv B^\delta_{jk,t}/P_{k,t}.
\end{aligned}$$

The optimality conditions of the representative households in each country are:

$$1 = E_t\beta_j \left(\frac{c_{j,t+1}}{c_{j,t}}\right)^{-1/\psi} \exp(-\pi_{j,t+1})(1+i_{j,t}), \ \ \forall j, \tag{11}$$

$$b_{jk,t} = \bar{b}_{jk}, \ \ \forall j, k \neq j, \tag{12}$$

$$b^\delta_{jk,t} = \bar{b}^\delta_{jk}, \ \ \forall j, k, \tag{13}$$

$$\bar{\varphi}\ell^{1/\varphi}_{j,t}c^{1/\psi}_{j,t} = w_{j,t}, \ \ \forall j, \tag{14}$$

$$c_{jk,t} = \eta_{jk}\left(p_{jk,t}\right)^{-\varsigma} c_{j,t}, \ \ \forall j, k, \tag{15}$$

$$c^m_{jk,t} = \left(\frac{p^m_{jk,t}}{p_{jk,t}}\right)^{-\varepsilon} c_{jk,t}, \ \ \forall j, k, m. \tag{16}$$

The resource constraints and evolution of net foreign assets of all countries besides the U.S., which also makes use of firms' profits, goods market clearing, and the

equilibrium lump-sum transfers, are:

$$c_{j,t} + \sum_{k=1}^{n} q_{j,t} q_{k,t}^{-1} \left(b_{jk,t} + b_{jk,t}^{\delta} \right) = \zeta_j^{-1} q_{j,t} \sum_{k=1}^{n} \zeta_k q_{k,t}^{-1} p_{kj,t} c_{kj,t} + a_{j,t}, \quad \forall j > 1, \tag{17}$$

$$a_{j,t+1} = \sum_{k=1}^{n} \exp(\Delta \log q_{j,t+1} - \Delta \log q_{k,t+1} - \pi_{k,t+1}) \times \left((1 + i_{k,t}) q_{j,t} q_{k,t}^{-1} b_{jk,t} + \left(\frac{1 + \delta P_{k,t+1}^{\delta}}{P_{k,t}^{\delta}} \right) q_{j,t} q_{k,t}^{-1} b_{jk,t}^{\delta} \right), \quad \forall j > 1. \tag{18}$$

The optimal price-setting conditions of firms in each origin, destination market, and invoicing currency, which also makes use of the symmetry of firms as well as goods market clearing, are:

$$\begin{aligned} \left(\exp(\pi_{kj,t}^{m}) - 1 \right) \exp(\pi_{kj,t}^{m}) = {} & \frac{\varepsilon - 1}{\chi_p} \left[\frac{\varepsilon}{\varepsilon - 1} \frac{w_{j,t}}{q_{j,t} q_{k,t}^{-1} p_{kj,t}^{m}} - 1 \right] \\ & + E_t \beta_j \left(\frac{c_{j,t+1}}{c_{j,t}} \right)^{-1/\psi} \exp(\Delta \log q_{j,t+1} - \Delta \log q_{k,t+1}) \frac{p_{kj,t+1}^{m} c_{kj,t+1}^{m}}{p_{kj,t}^{m} c_{kj,t}^{m}} \times \\ & \left(\exp(\pi_{kj,t+1}^{m}) - 1 \right) \exp(\pi_{kj,t+1}^{m}), \quad \forall j, k, m. \end{aligned} \tag{19}$$

The definition of excess returns, optimality conditions of arbitrageurs, and resource constraints of arbitrageurs are:

$$ex_{j,t+1} = \frac{E_{j,t}}{E_{j,t+1}} (1 + i_{j,t}) - (1 + i_{1,t}), \quad \forall j > 1, \tag{20}$$

$$ex_{j,t+1}^{\delta} \equiv \frac{E_{j,t}}{E_{j,t+1}} \frac{1 + \delta P_{j,t+1}^{\delta}}{P_{j,t}^{\delta}} - (1 + i_{1,t}), \quad \forall j, \tag{21}$$

$$\begin{aligned} E_t ex_{j,t+1} = {} & \frac{\gamma}{w_{a,t}} \left[\sum_{k=2}^{n} Cov_t \left(ex_{j,t+1}, ex_{k,t+1} \right) q_{k,t}^{-1} b_{ak,t} \right. \\ & \left. + \sum_{k=1}^{n} Cov_t \left(ex_{j,t+1}, ex_{k,t+1}^{\delta} \right) q_{k,t}^{-1} b_{ak,t}^{\delta} \right], \quad \forall j > 1, \end{aligned} \tag{22}$$

$$\begin{aligned} E_t ex_{j,t+1}^{\delta} = {} & \frac{\gamma}{w_{a,t}} \left[\sum_{k=2}^{n} Cov_t \left(ex_{j,t+1}^{\delta}, ex_{k,t+1} \right) q_{k,t}^{-1} b_{ak,t} \right. \\ & \left. + \sum_{k=1}^{n} Cov_t \left(ex_{j,t+1}^{\delta}, ex_{k,t+1}^{\delta} \right) q_{k,t}^{-1} b_{ak,t}^{\delta} \right], \quad \forall j, \end{aligned} \tag{23}$$

$$\sum_{j=1}^{n} q_{j,t}^{-1}\left[b_{aj,t}+b_{aj,t}^{\delta}\right]=w_{a,t}, \tag{24}$$

$$w_{a,t+1}=\xi\omega_a+(1-\xi)\exp(-\pi_{1,t+1})\Bigg[(1+i_{1,t})w_{a,t} + \sum_{j=2}^{n} ex_{j,t+1}q_{j,t}^{-1}b_{aj,t}+\sum_{j=1}^{n} ex_{j,t+1}^{\delta}q_{j,t}^{-1}b_{aj,t}^{\delta}\Bigg]. \tag{25}$$

The definitions of CPIs and mapping between prices are:

$$1=\left(\sum_{k=1}^{n}\eta_{jk}\left(p_{jk,t}\right)^{1-\varsigma}\right)^{\frac{1}{1-\varsigma}},\ \ \forall j, \tag{26}$$

$$p_{kj,t}=\left(\sum_{m=1}^{n}\theta_{kj}^{m}\left[p_{kj,t}^{m}\right]^{1-\varepsilon}\right)^{\frac{1}{1-\varepsilon}},\ \ \forall j,k, \tag{27}$$

$$\pi_{kj,t}=\Delta\log p_{kj,t}+\pi_{k,t},\ \ \forall j,k, \tag{28}$$

$$\pi_{kj,t}^{m}=\Delta\log p_{kj,t}^{m}-\Delta\log q_{k,t}+\Delta\log q_{m,t}+\pi_{m,t},\ \ \forall j,k,m, \tag{29}$$

$$\Delta\log q_{j,t}=\Delta\log E_{j,t}+\pi_{1,t}-\pi_{j,t},\ \ \forall j>1. \tag{30}$$

The market clearing conditions, also making use of firms' technologies, are:

$$\sum_{k=1}^{n}\zeta_k\sum_{m=1}^{n}\theta_{kj}^{m}c_{kj,t}^{m}=\zeta_j\ell_{j,t},\ \ \forall j, \tag{31}$$

$$\sum_{k=1}^{n}\zeta_k b_{kj,t}+b_{aj,t}=0,\ \ \forall j, \tag{32}$$

$$\sum_{k=1}^{n}\zeta_k b_{kj,t}^{\delta}+b_{aj,t}^{\delta}=0,\ \ \forall j. \tag{33}$$

The specification of policy is finally:

$$\log(1+i_{j,t})=\log(1+i_j)+\rho_i\left(\log(1+i_{j,t-1})-\log(1+i_j)\right)+(1-\rho_i)\phi_\pi\pi_{j,t}+\nu_{j,t},\ \ \forall j. \tag{34}$$

By Walras' Law, we have that the U.S. aggregate resource constraint

$$c_{1,t} + \sum_{k=1}^{n} q_{k,t}^{-1} \left(b_{1k,t} + b_{ak,t} + b_{ak,t}^{\delta} + b_{1k,t}^{\delta} \right) = \sum_{k=1}^{n} \zeta_k q_{k,t}^{-1} p_{k1,t} c_{k1,t}$$

$$+ \sum_{k=1}^{n} q_{k,t}^{-1} \exp(-\pi_{k,t}) \left((1 + i_{k,t-1})(b_{1k,t-1} + b_{ak,t-1}) + \left(\frac{1 + \delta P_{k,t}^{\delta}}{P_{k,t-1}^{\delta}} \right) \left(b_{1k,t-1}^{\delta} + b_{ak,t-1}^{\delta} \right) \right)$$

is satisfied as well.

(11)-(34) define a dynamical system of $3n^3 + 5n^2 + 13n - 3$ equations in $3n^3 + 5n^2 + 13n - 3$ unknowns

$$\{\{\{\{c_{jk,t}^{m}, p_{jk,t}^{m}, \pi_{jk,t}^{m}\}, c_{jk,t}, b_{jk,t}, b_{jk,t}^{\delta}, p_{jk,t}, \pi_{jk,t}\}, c_{j,t}, \ell_{j,t}, b_{aj,t}, b_{aj,t}^{\delta}, w_{j,t}, \pi_{j,t}, i_{j,t}, P_{j,t}^{\delta}, ex_{j,t+1}^{\delta}\},$$
$$\{q_{j,t}, E_{j,t}, ex_{j,t+1}, a_{j,t+1}\}_{j>1}, w_{a,t+1}\},$$

given the endogenous state variables $\{\{\{p_{jk,t-1}^{m}\}, p_{jk,t-1}\}, \{q_{j,t-1}, E_{j,t-1}, a_{j,t}\}_{j>1}, w_{a,t}\}$, exogenous state variables $\{\nu_{j,t}\}$ and any other driving forces which can be added to the model, and evolution of exogenous state variables.

A.2 First-order approximation with risk premia

We now characterize the first-order approximation studied in the paper. Since the approximation for all equilibrium conditions except arbitrageurs' portfolio choice (22)-(23) is standard, we focus on the latter. The approach and exposition closely follows that in Kekre and Lenel (2024a), which in turn follows Itskhoki and Mukhin (2021) and Borovicka et al. (2023).

Let us collect the state variables of the model in a vector $\boldsymbol{\theta}_t$, and let us scale the volatility of all innovations by the perturbation parameter σ. We will approximate the equilibrium around the point of approximation $\sigma \to 0$ and $\boldsymbol{\theta}_t = \boldsymbol{\theta}$.

Around that point, the first-order evolution of state variables is given by

$$\boldsymbol{\theta}_{t+1} = \boldsymbol{\theta} + \boldsymbol{\Delta}_{\theta}^{\theta} (\boldsymbol{\theta}_t - \boldsymbol{\theta}) + \boldsymbol{\Delta}_{\epsilon}^{\theta} \sigma \boldsymbol{\epsilon}_{t+1},$$

where $\boldsymbol{\Delta}_{\theta}^{\theta}$ is a matrix of coefficients on lagged state variables, $\boldsymbol{\Delta}_{\epsilon}^{\theta}$ is a matrix of coefficients on innovations to driving forces, and $\boldsymbol{\epsilon}_{t+1}$ is a vector of these innovations. We ignore coefficients on the perturbation parameter σ anticipating that these will be zero around the point of approximation. Similarly, a first-order approximation for $\log q_{k,t}$, $b_{ak,t}$, $b_{ak,t}^{\delta}$, $\log w_{a,t}$, $ex_{j,t+1}$, and $ex_{j,t+1}^{\delta}$ around the point of approximation is

given by

$$\begin{aligned}
\log q_{k,t} &= \log q_k + \boldsymbol{\delta}_\theta^{q_k} (\boldsymbol{\theta}_t - \boldsymbol{\theta}), \\
b_{ak,t} &= b_{ak} + \boldsymbol{\delta}_\theta^{b_{ak}} (\boldsymbol{\theta}_t - \boldsymbol{\theta}), \\
b_{ak,t}^\delta &= b_{ak}^\delta + \boldsymbol{\delta}_\theta^{b_{ak}^\delta} (\boldsymbol{\theta}_t - \boldsymbol{\theta}), \\
\log w_{a,t} &= \log w_a + \boldsymbol{\delta}_\theta^{w_a} (\boldsymbol{\theta}_t - \boldsymbol{\theta}), \\
ex_{j,t+1} &= ex_j + \boldsymbol{\delta}_\theta^{ex_j} (\boldsymbol{\theta}_t - \boldsymbol{\theta}) + \boldsymbol{\delta}_\epsilon^{ex_j} \sigma \boldsymbol{\epsilon}_{t+1}, \\
ex_{j,t+1}^\delta &= ex_j^\delta + \boldsymbol{\delta}_\theta^{ex_j^\delta} (\boldsymbol{\theta}_t - \boldsymbol{\theta}) + \boldsymbol{\delta}_\epsilon^{ex_j^\delta} \sigma \boldsymbol{\epsilon}_{t+1},
\end{aligned}$$

where $\boldsymbol{\delta}_\theta^x$ is a row vector of coefficients on the state variables and $\boldsymbol{\delta}_\epsilon^{ex_j}$, $\boldsymbol{\delta}_\epsilon^{ex_j^\delta}$ are row vectors of coefficients on innovations.

Now using the above first-order approximations, we have that

$$\begin{aligned}
Cov_t (ex_{j,t+1}, ex_{k,t+1}) &= Cov_t \left(ex_j + \boldsymbol{\delta}_\theta^{ex_j} (\boldsymbol{\theta}_t - \boldsymbol{\theta}) + \boldsymbol{\delta}_\epsilon^{ex_j} \sigma \boldsymbol{\epsilon}_{t+1}, ex_k + \boldsymbol{\delta}_\theta^{ex_k} (\boldsymbol{\theta}_t - \boldsymbol{\theta}) + \boldsymbol{\delta}_\epsilon^{ex_k} \sigma \boldsymbol{\epsilon}_{t+1} \right), \\
&= \boldsymbol{\delta}_\epsilon^{ex_j} \Sigma^2 \left(\boldsymbol{\delta}_\epsilon^{ex_k} \right)' \sigma^2, \\
&\equiv Cov(ex_{j,+1}, ex_{k,+1}) \sigma^2,
\end{aligned}$$

where Σ is the covariance matrix of the innovations, scaled appropriately by σ^2. An analogous argument holds for $Cov_t \left(ex_{j,t+1}, ex_{k,t+1}^\delta \right)$ and $Cov_t \left(ex_{j,t+1}^\delta, ex_{k,t+1}^\delta \right)$. Thus, substituting the first-order approximations for all variables into (22) yields

$$\begin{aligned}
& ex_j + \boldsymbol{\delta}_\theta^{ex_j} (\boldsymbol{\theta}_t - \boldsymbol{\theta}) \\
&= \frac{\gamma(\sigma)\sigma^2}{w_a + \boldsymbol{\delta}_\theta^{w_a} (\boldsymbol{\theta}_t - \boldsymbol{\theta})} \left[\sum_{k=2}^{n} Cov(ex_{j,+1}, ex_{k,+1}) \exp \left(- \log q_k - \boldsymbol{\delta}_\theta^{q_k} (\boldsymbol{\theta}_t - \boldsymbol{\theta}) \right) \left(b_{ak} + \boldsymbol{\delta}_\theta^{b_{ak}} (\boldsymbol{\theta}_t - \boldsymbol{\theta}) \right) \right. \\
&\qquad \left. + \sum_{k=1}^{n} Cov(ex_{j,+1}, ex_{k,+1}^\delta) \exp \left(- \log q_k - \boldsymbol{\delta}_\theta^{q_k} (\boldsymbol{\theta}_t - \boldsymbol{\theta}) \right) \left(b_{ak}^\delta + \boldsymbol{\delta}_\theta^{b_{ak}^\delta} (\boldsymbol{\theta}_t - \boldsymbol{\theta}) \right) \right].
\end{aligned}$$

Hence, at the point of approximation, under the assumption $\gamma(\sigma)\sigma^2 \to \Gamma$ as $\sigma \to 0$, we have that

$$ex_j = \frac{\Gamma}{w_a} \left[\sum_{k=2}^{n} Cov(ex_{j,+1}, ex_{k,+1}) q_k^{-1} b_{ak} + \sum_{k=1}^{n} Cov(ex_{j,+1}, ex_{k,+1}^\delta) q_k^{-1} b_{ak}^\delta \right].$$

Differentiating the above condition with respect to each state variable and evaluating

at the point of approximation, the coefficients will solve

$$E_t\hat{ex}_{j,t+1} = -ex_j\hat{w}_{a,t} + \frac{\Gamma}{w_a}\left[\sum_{k=2}^{n} Cov(ex_{j,+1}, ex_{k,+1})\left(-q_k^{-1}b_{ak}\hat{q}_{k,t} + q_k^{-1}\hat{b}_{ak,t}\right)\right.$$
$$\left. + \sum_{k=1}^{n} Cov(ex_{j,+1}, ex^{\delta}_{k,+1})\left(-q_k^{-1}b^{\delta}_{ak}\hat{q}_{k,t} + q_k^{-1}\hat{b}^{\delta}_{ak,t}\right)\right].$$

It is straightforward to show that a similar argument applies for consol pricing (23). At the point of approximation,

$$ex_j = \frac{\Gamma}{w_a}\left[\sum_{k=2}^{n} Cov(ex^{\delta}_{j,+1}, ex_{k,+1})q_k^{-1}b_{ak} + \sum_{k=1}^{n} Cov(ex^{\delta}_{j,+1}, ex^{\delta}_{k,+1})q_k^{-1}b^{\delta}_{ak}\right],$$

and to first order around the point of approximation,

$$E_t\hat{ex}^{\delta}_{j,t+1} = -ex^{\delta}_j\hat{w}_{a,t} + \frac{\Gamma}{w_a}\left[\sum_{k=2}^{n} Cov(ex^{\delta}_{j,+1}, ex_{k,+1})\left(-q_k^{-1}b_{ak}\hat{q}_{k,t} + q_k^{-1}\hat{b}_{ak,t}\right)\right.$$
$$\left. + \sum_{k=1}^{n} Cov(ex^{\delta}_{j,+1}, ex^{\delta}_{k,+1})\left(-q_k^{-1}b^{\delta}_{ak}\hat{q}_{k,t} + q_k^{-1}\hat{b}^{\delta}_{ak,t}\right)\right].$$

A.3 First-order aggregation of price-setting

We finally describe how the first-order approximation of firms' price-setting conditions can be conveniently aggregated to simplify the state space of the first-order solution.

The first-order approximation of (27) implies

$$\hat{p}_{kj,t} = \sum_{m=1}^{n} \theta^m_{kj}\left(\frac{p^m_{kj}}{p_{kj}}\right)^{1-\varepsilon}\hat{p}^m_{kj,t}$$

Now in steady-state, (19) and (27) imply $p^m_{kj} = p_{kj}$. Hence, the above simplifies to

$$\hat{p}_{kj,t} = \sum_{m=1}^{n} \theta^m_{kj}\hat{p}^m_{kj,t}. \tag{35}$$

The standard first-order approximation of (19) around the (zero-inflation) steady-

state is

$$\pi^m_{kj,t} = \frac{\varepsilon - 1}{\chi_p}\left(\hat{w}_{j,t} - \hat{q}_{j,t} + \hat{q}_{k,t} - \hat{p}^m_{kj,t}\right) + \beta_j E_t \pi^m_{kj,t+1},$$

where we do not use the $\hat{\cdot}$ notation for inflation, as is standard. Combining results implies

$$\sum_{m=1}^{n} \theta^m_{kj} \pi^m_{kj,t} = \frac{\epsilon - 1}{\chi_p}\left(\hat{w}_{j,t} - \hat{q}_{j,t} + \hat{q}_{k,t} - \hat{p}_{kj,t}\right) + \beta_j E_t \sum_{m=1}^{n} \theta^m_{kj} \pi^m_{kj,t+1}. \tag{36}$$

Moreover, substituting (35) into (28)-(29) implies

$$\sum_{m=1}^{n} \theta^m_{kj} \pi^m_{kj,t} = \pi_{kj,t} - \pi_{k,t} - \Delta\hat{q}_{k,t} + \sum_{m=1}^{n} \theta^m_{kj}\left(\pi_{m,t} + \Delta\hat{q}_{m,t}\right).$$

Substituting this into (36) implies

$$\begin{aligned}
\pi_{kj,t} - \pi_{k,t} - \Delta\hat{q}_{k,t} + \sum_{m=1}^{n} \theta^m_{kj}\left(\pi_{m,t} + \Delta\hat{q}_{m,t}\right) & \\
= \frac{\epsilon - 1}{\chi_p}\left(\hat{w}_{j,t} - \hat{q}_{j,t} + \hat{q}_{k,t} - \hat{p}_{kj,t}\right) & \\
+ \beta_j E_t\left[\pi_{kj,t+1} - \pi_{k,t+1} - \Delta\hat{q}_{k,t+1} + \sum_{m=1}^{n} \theta^m_{kj}\left(\pi_{m,t+1} + \Delta\hat{q}_{m,t+1}\right)\right]. &
\end{aligned}$$

Finally, note that the first-order approximation of (16) and (31), using as well the above results and that $c^m_{kj} = c_{kj}$ in steady-state, imply

$$\sum_{k=1}^{n} \frac{\zeta_k c_{kj}}{\zeta_j \ell_j} \hat{c}_{kj,t} = \hat{\ell}_{j,t}.$$

It follows that we can solve for the first-order dynamics without reference to the invoicing-specific prices/inflation/consumption. The relevant dynamics of prices in this system are simply

$$\begin{aligned}
\pi_{kj,t} - \pi_{k,t} - \Delta\hat{q}_{k,t} + \sum_{m=1}^{n} \theta^m_{kj}\left(\pi_{m,t} + \Delta\hat{q}_{m,t}\right) & \\
= \frac{\epsilon - 1}{\chi_p}\left(\hat{w}_{j,t} - \hat{q}_{j,t} + \hat{q}_{k,t} - \hat{p}_{kj,t}\right) &
\end{aligned}$$

$$+\beta_j E_t\left[\pi_{kj,t+1}-\pi_{k,t+1}-\Delta\hat{q}_{k,t+1}+\sum_{m=1}^{n}\theta_{kj}^{m}\left(\pi_{m,t+1}+\Delta\hat{q}_{m,t+1}\right)\right],\ \ \forall j,k,$$

$$\pi_{kj,t}=\Delta\hat{p}_{kj,t}+\pi_{k,t},\ \ \forall j,k,$$

and $\{\{\{\hat{p}^{m}_{jk,t-1}\}\}\}$ are not state variables of this system.

We can then solve for the invoicing-specific dynamics using the system

$$\begin{aligned}
\pi^{m}_{kj,t}&=\frac{\epsilon-1}{\chi_p}\left(\hat{w}_{j,t}-\hat{q}_{j,t}+\hat{q}_{k,t}-\hat{p}^{m}_{kj,t}\right)+\beta_j E_t\pi^{m}_{kj,t+1}\ \ \forall j,k,m,\\
\pi^{m}_{kj,t}&=\Delta\hat{p}^{m}_{kj,t}-\Delta\hat{q}_{k,t}+\Delta\hat{q}_{m,t}+\pi_{m,t},\ \ \forall j,k,m,\\
\hat{c}^{m}_{kj,t}&=-\varepsilon\left(\hat{p}^{m}_{kj,t}-\hat{p}_{kj,t}\right)+\hat{c}_{kj,t},\ \ \forall j,k,m.
\end{aligned}$$

$\{\{\{\hat{p}^{m}_{jk,t-1}\}\}\}$ are additional state variables of this auxiliary system.

B Proofs of analytical results

In this appendix we provide proofs of our analytical results in section 3.

B.1 Proposition 1

Proof. We first characterize the first-order approximation of the equilibrium (11)-(34) around an arbitrary steady-state, but already using that $\chi_p\to\infty$ and $\theta^{k}_{kj}=1$ for all j,k, which imply

$$\begin{aligned}
\hat{p}_{jk,t}&=0,\\
\pi_{jk,t}&=0,\\
\hat{q}_{j,t}&=\hat{E}_{j,t}.
\end{aligned}$$

We also use that, since households' positions in foreign one-period bonds and consols are exogenous and fixed,

$$\begin{aligned}
\hat{b}_{jk,t}&=0,\ \ \forall j,k\neq j,\\
\hat{b}^{\delta}_{jk,t}&=0,\ \ \forall j,k.
\end{aligned}$$

We finally also use that there is no interest rate inertia ($\rho_i=0$).

The optimality conditions of the representative households in each country are:

$$0 = -\frac{1}{\psi} E_t \Delta \hat{c}_{j,t+1} + \hat{i}_{j,t}, \ \ \forall j,$$
$$\hat{c}_{jk,t} = \hat{c}_{j,t}, \ \ \forall j, k.$$

The labor-leisure condition is only relevant to characterize the behavior of real wages, which are not of interest, so we ignore it.

The resource constraints and evolution of net foreign assets of all countries besides the U.S. are:

$$\begin{aligned}
&c_j \hat{c}_{j,t} + \hat{b}_{jj,t} + \sum_{k=1}^{n} q_j q_k^{-1} (b_{jk} + b_{jk}^{\delta}) \left(\hat{q}_{j,t} - \hat{q}_{k,t}\right) \\
&\quad = \zeta_j^{-1} q_j \sum_{k=1}^{n} \zeta_k q_k^{-1} p_{kj} c_{kj} \left(\hat{q}_{j,t} - \hat{q}_{k,t} + \hat{c}_{k,t}\right) + \hat{a}_{j,t}, \ \ \forall j > 1, \\
&\hat{a}_{j,t+1} = (1 + i_j) \hat{b}_{jj,t} + \sum_{k=1}^{n} \Bigg[\left((1 + i_k) q_j q_k^{-1} b_{jk} + \left(\frac{1 + \delta P_k^{\delta}}{P_k^{\delta}} \right) q_j q_k^{-1} b_{jk}^{\delta} \right) \left(\hat{q}_{j,t+1} - \hat{q}_{k,t+1}\right) \\
&\quad + (1 + i_k) q_j q_k^{-1} b_{jk} \hat{i}_{k,t} + \left(\frac{1 + \delta P_k^{\delta}}{P_k^{\delta}} \right) q_j q_k^{-1} b_{jk}^{\delta} \left(\frac{\delta P_k^{\delta}}{1 + \delta P_k^{\delta}} \hat{P}_{k,t+1}^{\delta} - \hat{P}_{k,t}^{\delta} \right) \Bigg], \ \ \forall j > 1.
\end{aligned}$$

The definition of excess returns, optimality conditions of arbitrageurs, and resource constraints of arbitrageurs are:

$$\begin{aligned}
&\hat{ex}_{j,t+1} = (1 + i_j) \left(-\Delta \hat{q}_{j,t+1} + \hat{i}_{j,t} \right) - (1 + i_1) \hat{i}_{1,t}, \ \ \forall j > 1, \\
&\hat{ex}_{j,t+1}^{\delta} = \left(\frac{1 + \delta P_j^{\delta}}{P_j^{\delta}} \right) \left(-\Delta \hat{q}_{j,t+1} + \frac{\delta P_j^{\delta}}{1 + \delta P_j^{\delta}} \hat{P}_{j,t+1}^{\delta} - \hat{P}_{j,t}^{\delta} \right) - (1 + i_1) \hat{i}_{1,t}, \ \ \forall j, \\
&E_t \hat{ex}_{j,t+1} = -ex_j \hat{w}_{a,t} + \frac{\Gamma}{w_a} \Bigg[\sum_{k=2}^{n} Cov\left(ex_{j,+1}, ex_{k,+1}\right) \left(q_k^{-1} \hat{b}_{ak,t} - q_k^{-1} b_{ak} \hat{q}_{k,t} \right) \\
&\quad + \sum_{k=1}^{n} Cov\left(ex_{j,+1}, ex_{k,+1}^{\delta}\right) \left(q_k^{-1} \hat{b}_{ak,t}^{\delta} - q_k^{-1} b_{ak}^{\delta} \hat{q}_{k,t} \right) \Bigg], \ \ \forall j > 1, \\
&E_t \hat{ex}_{j,t+1}^{\delta} = -ex_j^{\delta} \hat{w}_{a,t} + \frac{\Gamma}{w_a} \Bigg[\sum_{k=2}^{n} Cov\left(ex_{j,+1}^{\delta}, ex_{k,+1}\right) \left(q_k^{-1} \hat{b}_{ak,t} - q_k^{-1} b_{ak} \hat{q}_{k,t} \right) \\
&\quad + \sum_{k=1}^{n} Cov\left(ex_{j,+1}^{\delta}, ex_{k,+1}^{\delta}\right) \left(q_k^{-1} \hat{b}_{ak,t}^{\delta} - q_k^{-1} b_{ak}^{\delta} \hat{q}_{k,t} \right) \Bigg], \ \ \forall j,
\end{aligned}$$

$$\sum_{j=1}^{n} q_j^{-1} \left(\hat{b}_{aj,t} + \hat{b}^{\delta}_{aj,t} - \left(b_{aj} + b^{\delta}_{aj} \right) \hat{q}_{j,t} \right) = w_a \hat{w}_{a,t},$$

$$\begin{aligned} w_a \hat{w}_{a,t+1} = (1-\xi) \Bigg[& (1+i_1) w_a \hat{i}_{1,t} + (1+i_1) w_a \hat{w}_{a,t} \\ & + \sum_{j=2}^{n} q_j^{-1} \left(b_{aj} \hat{ex}_{j,t+1} - ex_j b_{aj} \hat{q}_{j,t} + ex_j \hat{b}_{aj,t} \right) \\ & + \sum_{j=1}^{n} q_j^{-1} \left(b^{\delta}_{aj} \hat{ex}^{\delta}_{j,t+1} - ex^{\delta}_j b^{\delta}_{aj} \hat{q}_{j,t} + ex^{\delta}_j \hat{b}^{\delta}_{aj,t} \right) \Bigg]. \end{aligned}$$

The bond market clearing conditions are:

$$\sum_{k=1}^{n} \zeta_k \hat{b}_{kj,t} + \hat{b}_{aj,t} = 0, \ \ \forall j,$$
$$\sum_{k=1}^{n} \zeta_k \hat{b}^{\delta}_{kj,t} + \hat{b}^{\delta}_{aj,t} = 0, \ \ \forall j.$$

The goods market clearing condition is only relevant to characterize the implied behavior of labor, which is again not of interest, so we ignore it.

The specification of policy is finally:

$$\hat{i}_{j,t} = \nu_{j,t}, \ \ \forall j.$$

As stated in the claim, we will now characterize the effects of a U.S. monetary shock $\nu_{1,t}$ around a steady-state with small excess returns, small one-period bond positions, and zero consol positions. We first formalize what this condition means.

Suppose we parameterize primitives as:

$$\begin{aligned} \bar{b}_{jk} &= \kappa \tilde{\bar{b}}_{jk}, \ \ \forall j, k \neq j, \\ \bar{b}^{\delta}_{jk} &= 0, \ \ \forall j, k, \\ \beta_j &= \beta + \kappa \left(\tilde{\beta}_j - \beta \right), \ \ \forall j, \\ \Gamma &= \tilde{\Gamma} \kappa^2, \\ \omega_a &= \tilde{\omega}_a \kappa^2, \end{aligned}$$

for some $\kappa > 0$. Let us further define the "tilde" allocation

$$\begin{aligned}
\tilde{i}_1 &\equiv (i_1 - (\beta^{-1} - 1))/\kappa, \\
\tilde{ex}_j &\equiv ex_j/\kappa, \ \ \forall j > 1, \\
\tilde{ex}_j^\delta &\equiv ex_j^\delta/\kappa, \ \ \forall j, \\
\tilde{w}_a &\equiv w_a/\kappa^2, \\
\tilde{b}_{aj} &\equiv b_{aj}/\kappa, \ \ \forall j > 1, \\
\tilde{b}_{jk} &\equiv b_{jk}/\kappa, \ \ \forall j, k \neq j.
\end{aligned}$$

Now in any steady-state with zero household consol positions and thus zero arbitrageur consol positions, the equilibrium conditions from appendix A imply that $\{\{i_j, ex_j^\delta, q_j^{-1} b_{aj}\}, w_a\}$ must solve

$$\begin{aligned}
1 &= \beta_j (1 + i_j), \ \ \forall j, \\
i_j - i_1 &= \frac{\Gamma}{w_a} \sum_{k=2}^{n} Cov\left(ex_{j,+1}, ex_{k,+1}\right) q_k^{-1} b_{ak}, \ \ \forall j > 1, \\
ex_j^\delta &= \frac{\Gamma}{w_a} \sum_{k=2}^{n} Cov\left(ex_{j,+1}^\delta, ex_{k,+1}\right) q_k^{-1} b_{ak}, \ \ \forall j, \\
\sum_{j=1}^{n} q_j^{-1} b_{aj} &= w_a, \\
w_a &= \xi \omega_a + (1 - \xi) \left[(1 + i_1) w_a + \sum_{j=2}^{n} (i_j - i_1) q_j^{-1} b_{aj} \right].
\end{aligned}$$

Then substituting in for the "tilde" allocation above and evaluating at the limit $\kappa \to 0$ implies that $\{\{\tilde{i}_j, \tilde{ex}_j^\delta, q_j^{-1} \tilde{b}_{aj}\}, \tilde{w}_a\}$ solve

$$\begin{aligned}
0 &= \beta^{-1} \left(\tilde{\beta}_j - \beta \right) + \beta \tilde{i}_1, \ \ \forall j, \\
\tilde{ex}_j &= \frac{\tilde{\Gamma}}{\tilde{w}_a} \sum_{k=2}^{n} Cov\left(ex_{j,+1}, ex_{k,+1}\right) q_k^{-1} \tilde{b}_{ak}, \ \ \forall j > 1, \\
\tilde{ex}_j^\delta &= \frac{\tilde{\Gamma}}{\tilde{w}_a} \sum_{k=2}^{n} Cov\left(ex_{j,+1}^\delta, ex_{k,+1}\right) q_k^{-1} \tilde{b}_{ak}, \ \ \forall j, \\
\sum_{j=1}^{n} q_j^{-1} \tilde{b}_{aj} &= 0,
\end{aligned}$$

$$\tilde{w}_a = \xi\tilde{\omega}_a + (1-\xi)\left[\beta^{-1}\tilde{w}_a + \sum_{j=2}^{n} e\tilde{x}_j q_j^{-1}\tilde{b}_{aj}\right].$$

This is a convenient limit because it means that as $\kappa \to 0$, all valuation effects in the household budget constraints vanish in the first-order approximation (since $b_{jk} \to 0$ for all j, k). At the same time, arbitrageurs' effective risk aversion Γ/w_a remains nonzero even in the limit, and the valuation effects of shocks on arbitrageur balance sheets, and the resulting endogenous price of risk, also do not vanish. In particular, note that

$$\frac{ex_j q_j^{-1} b_{aj}}{w_a} \to \frac{e\tilde{x}_j q_j^{-1}\tilde{b}_{aj}}{\tilde{w}_a} \;\; \forall j,$$

which will not in general be zero.

We also make only for expositional simplicity the assumptions that

$$\sum_{k=1}^{n} \zeta_k \eta_{kj} = \zeta_j, \;\; \forall j,$$
$$\bar{\varphi} = \frac{\varepsilon - 1}{\varepsilon},$$

which implies that the limiting steady-state features:

$$c_j = 1, \;\; \forall j,$$
$$\ell_j = 1, \;\; \forall j,$$
$$c_{jk} = \eta_{jk}, \;\; \forall j, k,$$
$$p_{jk} = 1, \;\; \forall j, k,$$
$$q_j = 1, \;\; \forall j.$$

Given these assumptions on the steady-state, we can write the first-order system as follows. We do so in matrix notation for brevity. The local currency bond optimality conditions together with the monetary policy rules imply

$$\mathbf{0}_n = -\frac{1}{\psi} E_t \Delta\hat{\boldsymbol{c}}_{t+1} + \left(\mathbf{1}_n - \boldsymbol{J}'_{n-1}\mathbf{1}_{n-1}\right)\nu_{1,t}, \tag{37}$$

where $\mathbf{0}_n$ is a $n \times 1$ vector of zeros; $\hat{\boldsymbol{c}}_t$ is a $n \times 1$ vector of $\hat{c}_{j,t}$; $\mathbf{1}_n$ is a $n \times 1$ vector of ones; and $\boldsymbol{J}_{n-1}$ is a $(n-1) \times n$ matrix of zeros in the first column, and $\boldsymbol{I}_{n-1}$ in the

remaining columns.

The non-U.S. resource constraints, together with the evolution of net worth and the assumption $\sum_{k=1}^{n} \zeta_k \eta_{kj} = \zeta_j$, imply

$$\boldsymbol{J}_{n-1}\hat{\boldsymbol{c}}_t + \beta\hat{\boldsymbol{a}}_t = \hat{\boldsymbol{q}}_t - \boldsymbol{J}_{n-1}\boldsymbol{\zeta}^{-1}(\boldsymbol{\zeta}\boldsymbol{\eta})'\boldsymbol{J}'_{n-1}\hat{\boldsymbol{q}}_t + \boldsymbol{J}_{n-1}\boldsymbol{\zeta}^{-1}(\boldsymbol{\zeta}\boldsymbol{\eta})'\hat{\boldsymbol{c}}_t + \hat{\boldsymbol{a}}_{t-1}, \tag{38}$$

where $\hat{\boldsymbol{a}}_t$ is a $(n-1)\times 1$ vector of $\hat{a}_{j,t}$; $\hat{\boldsymbol{q}}_t$ is a $(n-1)\times 1$ vector of $\hat{q}_{j,t}$; $\boldsymbol{\zeta}$ is a $n\times n$ matrix of ζ_j on the diagonals, and zero on off-diagonals; and $\boldsymbol{\eta}$ is a $n \times n$ matrix of η_{jk}. Note that we have changed the timing convention for $\hat{a}_{j,t}$ given the absence of valuation effects on the household budget constraints around the assumed steady-state.

The definition of excess returns, together with the monetary policy rules, implies

$$\hat{\boldsymbol{ex}}_{t+1} = \beta^{-1}\left(-\Delta\hat{\boldsymbol{q}}_{t+1} - \mathbf{1}_{n-1}\nu_{1,t}\right). \tag{39}$$

The arbitrageur optimality conditions for one-period bonds are

$$E_t\hat{\boldsymbol{ex}}_{t+1} = -\boldsymbol{ex}\hat{w}_{a,t} - \beta\frac{\Gamma}{w_a}\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}'_{n-1}\hat{\boldsymbol{a}}_t, \tag{40}$$

where $\boldsymbol{C}$ is a $(n-1)\times(n-1)$ matrix of $Cov\left(ex_{j,+1}, ex_{k,+1}\right)$ (in an abuse of the notation for $\boldsymbol{C}$ used in the main text, which includes covariances with consol bond returns).

Finally, the evolution of arbitrageur net worth is

$$\begin{aligned}\hat{w}_{a,t+1} = (1-\xi)\Bigg[&\beta^{-1}\nu_{1,t} + \beta^{-1}\hat{w}_{a,t} \\ &+ \frac{1}{w_a}\boldsymbol{b}'_a\hat{\boldsymbol{ex}}_{t+1} - \frac{1}{w_a}\boldsymbol{ex}'\mathrm{diag}(\boldsymbol{b}_a)\hat{\boldsymbol{q}}_t - \frac{1}{w_a}\boldsymbol{ex}'\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}'_{n-1}\hat{\boldsymbol{a}}_t\Bigg],\end{aligned} \tag{41}$$

where $\boldsymbol{ex}$ is a $(n-1)\times 1$ vector of ex_j; $\boldsymbol{b}_a$ is a $(n-1)\times 1$ vector of b_{aj} for $j>1$; and $\mathrm{diag}(\boldsymbol{b}_a)$ is the diagonal matrix constructed from elements of $\boldsymbol{b}_a$.

Then, (37)-(41) is a linear system of expectational difference equations in $\{\hat{\boldsymbol{c}}_t, \hat{\boldsymbol{q}}_t, \hat{\boldsymbol{a}}_t, \hat{w}_{a,t+1}, \hat{\boldsymbol{ex}}_{t+1}\}$, given the state variables $\{\hat{\boldsymbol{a}}_{t-1}, \hat{w}_{a,t}, \nu_{1,t}\}$ and the evolution of the exogenous state variable $\nu_{1,t}$. We conjecture a solution of the form

$$\begin{aligned}\hat{\boldsymbol{c}}_t &= \boldsymbol{\Delta}^c_a\hat{\boldsymbol{a}}_{t-1} + \boldsymbol{\delta}^c_w\hat{w}_{a,t} + \boldsymbol{\delta}^c_\nu\nu_{1,t},\\ \hat{\boldsymbol{q}}_t &= \boldsymbol{\Delta}^q_a\hat{\boldsymbol{a}}_{t-1} + \boldsymbol{\delta}^q_w\hat{w}_{a,t} + \boldsymbol{\delta}^q_\nu\nu_{1,t},\\ \hat{\boldsymbol{a}}_t &= \boldsymbol{\Delta}^a_a\hat{\boldsymbol{a}}_{t-1} + \boldsymbol{\delta}^a_w\hat{w}_{a,t} + \boldsymbol{\delta}^a_\nu\nu_{1,t},\end{aligned}$$

$$\hat{w}_{a,t+1} = \boldsymbol{\delta}_a^{Ew} \hat{\boldsymbol{a}}_{t-1} + \delta_w^{Ew} \hat{w}_{a,t} + \delta_\nu^{Ew} \nu_{1,t} + \delta_\epsilon^w \epsilon_{1,t+1},$$
$$\hat{\boldsymbol{ex}}_{t+1} = \boldsymbol{\Delta}_a^{Eex} \hat{\boldsymbol{ex}}_{t-1} + \boldsymbol{\delta}_w^{Eex} \hat{w}_{a,t} + \boldsymbol{\delta}_\nu^{Eex} \nu_{1,t} + \boldsymbol{\delta}_\epsilon^{ex} \epsilon_{1,t+1},$$

where $\boldsymbol{\Delta}_\cdot^\cdot$, $\boldsymbol{\delta}_\cdot^\cdot$, and $\delta_\cdot^\cdot$ are matrices, vectors, and scalars of coefficients, respectively. We use the method of undetermined coefficients to characterize the solution.

Our main object of interest in the proposition is the impact effect of $\epsilon_{1,t}$ on exchange rates. This is given by

$$\boldsymbol{\delta}_w^q \delta_\epsilon^w + \boldsymbol{\delta}_\nu^q.$$

By (38), we have that

$$\hat{\boldsymbol{q}}_t = \boldsymbol{Q}_{n-1} \left[\boldsymbol{J}_{n-1} \left(\boldsymbol{I}_n - \boldsymbol{\zeta}^{-1} (\boldsymbol{\zeta}\boldsymbol{\eta})' \right) \hat{\boldsymbol{c}}_t + \beta \hat{\boldsymbol{a}}_t - \hat{\boldsymbol{a}}_{t-1} \right]$$

where

$$\boldsymbol{Q}_{n-1} \equiv \left(\boldsymbol{I}_{n-1} - \boldsymbol{J}_{n-1} \boldsymbol{\zeta}^{-1} (\boldsymbol{\zeta}\boldsymbol{\eta})' \boldsymbol{J}'_{n-1} \right)^{-1},$$

assuming the expression in parenthesis is invertible. It follows that

$$\boldsymbol{\delta}_w^q = \boldsymbol{Q}_{n-1} \left[\boldsymbol{J}_{n-1} \left(\boldsymbol{I}_n - \boldsymbol{\zeta}^{-1} (\boldsymbol{\zeta}\boldsymbol{\eta})' \right) \boldsymbol{\delta}_w^c + \beta \boldsymbol{\delta}_w^a \right],$$
$$\boldsymbol{\delta}_\nu^q = \boldsymbol{Q}_{n-1} \left[\boldsymbol{J}_{n-1} \left(\boldsymbol{I}_n - \boldsymbol{\zeta}^{-1} (\boldsymbol{\zeta}\boldsymbol{\eta})' \right) \boldsymbol{\delta}_\nu^c + \beta \boldsymbol{\delta}_\nu^a \right].$$

We thus characterize $\{\boldsymbol{\delta}_w^c, \boldsymbol{\delta}_w^a, \delta_\epsilon^w, \boldsymbol{\delta}_\nu^c, \boldsymbol{\delta}_\nu^a\}$ before combining them using the above formulas.

$\boldsymbol{\delta}_w^c$ Using the method of undetermined coefficients, we can show

$$\boldsymbol{\delta}_w^c = \boldsymbol{0}_n,$$

a $n \times 1$ vector of zeros. Intuitively, with fully sticky prices in local currency and central banks which target inflation, changes in arbitrageur wealth and thus risk pricing do not affect the consumption decisions of households, which are pinned down by the path of monetary policy alone.

$\boldsymbol{\delta}_w^a$ Using the method of undetermined coefficients, we can show that $\boldsymbol{\delta}_w^a$ implicitly solves

$$\left[\left(1-\beta(1-\xi)\left[\left(\beta^{-1}-\frac{1}{w_a}\boldsymbol{b}_a'\boldsymbol{ex}\right)\right.\right.\right.$$
$$\left.\left.-\frac{1}{w_a}\left[\left(\boldsymbol{ex}'+\beta\frac{\Gamma}{w_a}\boldsymbol{b}_a'\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}'+\beta\boldsymbol{ex}'\text{diag}(\boldsymbol{b}_a)\boldsymbol{Q}_{n-1}\right]\boldsymbol{\delta}_w^a\right]\right)\boldsymbol{I}_{n-1}$$
$$\left.+\beta\left(\boldsymbol{I}_{n-1}-\boldsymbol{\Delta}_a^a\right)+\beta^2\frac{\Gamma}{w_a}\boldsymbol{Q}_{n-1}^{-1}\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}'\right]\boldsymbol{\delta}_w^a=-\beta\boldsymbol{Q}_{n-1}^{-1}\boldsymbol{ex}. \quad (42)$$

We can also show that $\boldsymbol{\Delta}_a^a$ implicitly solves

$$\left(\beta\boldsymbol{\Delta}_a^a-\boldsymbol{I}_{n-1}\right)\left(\boldsymbol{\Delta}_a^a-\boldsymbol{I}_{n-1}\right)$$
$$=\beta(1-\xi)\boldsymbol{\delta}_w^a\left[\frac{1}{w_a}\left(\boldsymbol{ex}'+\beta\frac{\Gamma}{w_a}\boldsymbol{b}_a'\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}'\boldsymbol{\Delta}_a^a-\frac{1}{w_a}\boldsymbol{ex}'\text{diag}(\boldsymbol{b}_a)\boldsymbol{Q}_{n-1}\left(\boldsymbol{I}_{n-1}-\beta\boldsymbol{\Delta}_a^a\right)\right]$$
$$+\beta^2\frac{\Gamma}{w_a}\boldsymbol{Q}_{n-1}^{-1}\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}'\boldsymbol{\Delta}_a^a. \quad (43)$$

As is evident, in general we cannot solve (42) and (43) in closed form. However, we can characterize them in closed form in a limiting case, and we can approximate them away from that case, as we further describe below.

$\boldsymbol{\delta}_\epsilon^w$ Using the method of undetermined coefficients, we can show

$$\delta_\epsilon^w=-\left(w_a+(1-\xi)\boldsymbol{b}_a'\beta^{-1}\boldsymbol{Q}_{n-1}\left[\boldsymbol{J}_{n-1}\left(\boldsymbol{I}_n-\boldsymbol{\zeta}^{-1}(\boldsymbol{\zeta}\boldsymbol{\eta})'\right)\boldsymbol{\delta}_w^c+\beta\boldsymbol{\delta}_w^a\right]\right)^{-1}\times$$
$$(1-\xi)\boldsymbol{b}_a'\beta^{-1}\boldsymbol{Q}_{n-1}\left[\boldsymbol{J}_{n-1}\left(\boldsymbol{I}_n-\boldsymbol{\zeta}^{-1}(\boldsymbol{\zeta}\boldsymbol{\eta})'\right)\boldsymbol{\delta}_\nu^c+\beta\boldsymbol{\delta}_\nu^a\right].$$

$\boldsymbol{\delta}_\nu^c$ Using the method of undetermined coefficients, we can show

$$\boldsymbol{\delta}_\nu^c=-(1-\rho_1)^{-1}\psi\left(\boldsymbol{1}_n-\boldsymbol{J}_{n-1}'\boldsymbol{1}_{n-1}\right).$$

That is, a U.S. monetary tightening lowers U.S. consumption and leaves consumption in the rest of the world unchanged. This is again a consequence of fully sticky prices in local currency and central banks which target inflation.

$\boldsymbol{\delta}_\nu^a$ Using the method of undetermined coefficients, we can show

$$\boldsymbol{\delta}_\nu^a=\Big[\left(1-\beta\rho_1\right)\boldsymbol{I}_{n-1}+\beta\left(\boldsymbol{I}_{n-1}-\boldsymbol{\Delta}_a^a\right)$$

$$+\beta(1-\xi)\frac{1}{w_a}\boldsymbol{\delta}_w^a\left[\left(\boldsymbol{ex}'+\beta\frac{\Gamma}{w_a}\boldsymbol{b}_a'\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}'+\beta\boldsymbol{ex}'\text{diag}(\boldsymbol{b}_a)\boldsymbol{Q}_{n-1}\right]+\beta^2\frac{\Gamma}{w_a}\boldsymbol{Q}_{n-1}^{-1}\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}'\Bigg]^{-1}\times$$
$$\left[\beta(1-\xi)\boldsymbol{\delta}_w^a\left[\beta^{-1}-(1-\rho_1)^{-1}\psi\frac{1}{w_a}\boldsymbol{ex}'\text{diag}(\boldsymbol{b}_a)\mathbf{1}_{n-1}\right]+(1-\psi)\boldsymbol{Q}_{n-1}^{-1}\mathbf{1}_{n-1}\right]. \quad (44)$$

Combining these We can finally combine the above results to obtain the impact effect of a U.S. monetary tightening on exchange rates

$$\boldsymbol{\delta}_w^q\delta_\epsilon^w+\boldsymbol{\delta}_\nu^q,$$

again using

$$\boldsymbol{\delta}_w^q=\boldsymbol{Q}_{n-1}\left[\boldsymbol{J}_{n-1}\left(\boldsymbol{I}_n-\boldsymbol{\zeta}^{-1}(\boldsymbol{\zeta}\boldsymbol{\eta})'\right)\boldsymbol{\delta}_w^c+\beta\boldsymbol{\delta}_w^a\right],$$
$$\boldsymbol{\delta}_\nu^q=\boldsymbol{Q}_{n-1}\left[\boldsymbol{J}_{n-1}\left(\boldsymbol{I}_n-\boldsymbol{\zeta}^{-1}(\boldsymbol{\zeta}\boldsymbol{\eta})'\right)\boldsymbol{\delta}_\nu^c+\beta\boldsymbol{\delta}_\nu^a\right].$$

Since we cannot in general obtain a closed form for $\boldsymbol{\delta}_w^a$ and $\boldsymbol{\delta}_\nu^a$, it will be convenient to approximate these solutions to first order in $\boldsymbol{ex}$ and Γ/w_a around the point with $\boldsymbol{ex}\to 0$ and $\Gamma/w_a\to 0$. At that point of approximation, it is straightforward to show that the unique stable solution to (42)-(44) is

$$\lim_{\boldsymbol{ex},\Gamma\to 0}\boldsymbol{\Delta}_a^a=\boldsymbol{I}_{n-1},$$
$$\lim_{\boldsymbol{ex},\Gamma\to 0}\boldsymbol{\delta}_w^a=\mathbf{0}_{n-1},$$
$$\lim_{\boldsymbol{ex},\Gamma\to 0}\boldsymbol{\delta}_\nu^a=(1-\beta\rho_1)^{-1}(1-\psi)\boldsymbol{Q}_{n-1}^{-1}\mathbf{1}_{n-1}.$$

That is, in this limit net foreign assets follow a unit root, and respond to U.S. monetary shocks depending on households' intertemporal elasticity of substitution ψ as described in the main text. It follows that

$$\lim_{\boldsymbol{ex},\Gamma\to 0}\left[\boldsymbol{\delta}_w^q\delta_\epsilon^w+\boldsymbol{\delta}_\nu^q\right]=\left[\beta^{-1}-\rho_1\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_1)^{-1}\psi\right]\mathbf{1}_{n-1}.$$

We can then differentiate (42)-(44) with respect to $\boldsymbol{ex}$ and evaluate at the point of approximation, and we obtain

$$\lim_{\boldsymbol{ex},\Gamma\to 0}\frac{d\boldsymbol{\Delta}_a^a}{d\boldsymbol{ex}}=\mathbf{0}_{(n-1)\times(n-1)},$$

$$\lim_{\boldsymbol{ex},\Gamma\to 0} \frac{d\boldsymbol{\delta}_w^a}{d\boldsymbol{ex}} = -\frac{1}{\xi}\beta \boldsymbol{Q}_{n-1}^{-1},$$
$$\lim_{\boldsymbol{ex},\Gamma\to 0} \frac{d\boldsymbol{\delta}_\nu^a}{d\boldsymbol{ex}} = \mathbf{0}_{(n-1)\times(n-1)}.$$

Using as well that

$$\lim_{\boldsymbol{ex},\Gamma\to 0} \delta_\epsilon^w = -(1-\xi)\frac{\sum_{k=2}^n b_{ak}}{w_a}\beta^{-1}\left[\beta^{-1}-\rho_1\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_1)^{-1}\psi\right],$$

it follows that

$$\lim_{\boldsymbol{ex},\Gamma\to 0} \frac{d\left[\boldsymbol{\delta}_w^q\delta_\epsilon^w+\boldsymbol{\delta}_\nu^q\right]}{d\boldsymbol{ex}} = \frac{1-\xi}{\xi}\beta\frac{\sum_{k=2}^n b_{ak}}{w_a}\left[\beta^{-1}-\rho_1\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_1)^{-1}\psi\right]\boldsymbol{I}_{n-1}.$$

Finally, we can differentiate (42)-(44) with respect to Γ/w_a and evaluate at the point of approximation, and we obtain

$$\lim_{\boldsymbol{ex},\Gamma\to 0} \frac{d\boldsymbol{\Delta}_a^a}{d[\Gamma/w_a]} = -(1-\beta)^{-1}\beta^2\boldsymbol{Q}_{n-1}^{-1}\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}'_{n-1},$$
$$\lim_{\boldsymbol{ex},\Gamma\to 0} \frac{d\boldsymbol{\delta}_w^a}{d[\Gamma/w_a]} = \mathbf{0}_{n-1},$$
$$\lim_{\boldsymbol{ex},\Gamma\to 0} \frac{d\boldsymbol{\delta}_\nu^a}{d[\Gamma/w_a]} = -(\beta^{-1}-\rho_1)^{-2}(1-\beta)^{-1}\boldsymbol{Q}_{n-1}^{-1}\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\eta}_1'(1-\psi),$$

where $\boldsymbol{\eta}_1$ is the $1\times n$ vector of $\{\eta_{1k}\}$, the expenditure shares of U.S. households on goods from all origins. It follows that

$$\lim_{\boldsymbol{ex},\Gamma\to 0} \frac{d\left[\boldsymbol{\delta}_w^q\delta_\epsilon^w+\boldsymbol{\delta}_\nu^q\right]}{d[\Gamma/w_a]} = -(\beta^{-1}-\rho_1)^{-2}(1-\beta)^{-1}\beta\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\eta}_1'(1-\psi).$$

Combining these results, we obtain that to first order in $\boldsymbol{ex}$ and Γ/w_a,

$$\begin{aligned}\boldsymbol{\delta}_w^q\delta_\epsilon^w+\boldsymbol{\delta}_\nu^q = {}& \left[\beta^{-1}-\rho_1\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_1)^{-1}\psi\right]\mathbf{1}_{n-1}\\ &+\frac{1-\xi}{\xi}\beta\frac{\sum_{k=2}^n b_{ak}}{w_a}\left[\beta^{-1}-\rho_1\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_1)^{-1}\psi\right]\boldsymbol{ex}\\ &-(\beta^{-1}-\rho_1)^{-2}(1-\beta)^{-1}\beta\frac{\Gamma}{w_a}\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\eta}_1'(1-\psi).\end{aligned}$$

Noting that

$$\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\eta}_1' = \begin{bmatrix} Cov\left(ex_{2,+1}, \sum_{k=2}^{n} \eta_{1k} ex_{k,+1}\right) \\ \vdots \\ Cov\left(ex_{n,+1}, \sum_{k=2}^{n} \eta_{1k} ex_{k,+1}\right) \end{bmatrix},$$

we obtain the claimed result in the proposition. □

B.2 Proposition 2

Proof. The arbitrageur optimality conditions for consols can be written in vector notation as

$$E_t \hat{\boldsymbol{ex}}_{t+1}^{\delta} = -\boldsymbol{ex}^{\delta} \hat{w}_{a,t} - \beta \frac{\Gamma}{w_a} \boldsymbol{C}^{\delta} \boldsymbol{J}_{n-1} \boldsymbol{\zeta} \boldsymbol{J}_{n-1}' \hat{\boldsymbol{a}}_t.$$

where $\hat{\boldsymbol{ex}}_t^{\delta}$ is a $n \times 1$ vector of $\hat{ex}_{j,t}^{\delta}$; and $\boldsymbol{C}^{\delta}$ is a $n \times (n-1)$ matrix of $Cov\left(ex_{j,+1}^{\delta}, ex_{k,+1}\right)$.

And the definition of excess consol returns, together with the monetary policy rules, implies

$$\hat{\boldsymbol{ex}}_{t+1}^{\delta} = -\beta^{-1} \boldsymbol{J}_{n-1}' \Delta \hat{\boldsymbol{q}}_{t+1} + \delta \hat{\boldsymbol{P}}_{t+1}^{\delta} - \beta^{-1} \hat{\boldsymbol{P}}_t^{\delta} - \beta^{-1} \mathbf{1}_n \nu_{1,t}.$$

Using the definition of excess currency returns, this can be equivalently written

$$\hat{\boldsymbol{ex}}_{t+1}^{\delta} = \boldsymbol{J}_{n-1}' \hat{\boldsymbol{ex}}_{t+1} + \delta \hat{\boldsymbol{P}}_{t+1}^{\delta} - \beta^{-1} \hat{\boldsymbol{P}}_t^{\delta} - \beta^{-1} \left(\mathbf{1}_n - \boldsymbol{J}_{n-1}' \mathbf{1}_{n-1}\right) \nu_{1,t}.$$

Combining these results and using (40) yields

$$\begin{aligned} \delta E_t \hat{\boldsymbol{P}}_{t+1}^{\delta} - \beta^{-1} \hat{\boldsymbol{P}}_t^{\delta} - \beta^{-1} \left(\mathbf{1}_n - \boldsymbol{J}_{n-1}' \mathbf{1}_{n-1}\right) \nu_{1,t} & \\ = -\left(\boldsymbol{ex}^{\delta} - \boldsymbol{J}_{n-1}' \boldsymbol{ex}\right) \hat{w}_{a,t} - \beta \frac{\Gamma}{w_a} \left(\boldsymbol{C}^{\delta} - \boldsymbol{J}_{n-1}' \boldsymbol{C}\right) \boldsymbol{J}_{n-1} \boldsymbol{\zeta} \boldsymbol{J}_{n-1}' \hat{\boldsymbol{a}}_t. & \end{aligned}$$

It is already evident that it is the average local currency bond carry trade returns $\boldsymbol{ex}^{\delta} - \boldsymbol{J}_{n-1}' \boldsymbol{ex}$, and covariances with these returns $\boldsymbol{C}^{\delta} - \boldsymbol{J}_{n-1}' \boldsymbol{C}$, which will govern the response of consol prices to a change in risk pricing. Defining a solution of the form

$$\hat{\boldsymbol{P}}_t^{\delta} = \boldsymbol{\Delta}_a^{P^{\delta}} \hat{\boldsymbol{a}}_{t-1} + \boldsymbol{\delta}_w^{P^{\delta}} \hat{w}_{a,t} + \boldsymbol{\delta}_{\nu}^{P^{\delta}} \nu_{1,t},$$

the method of undetermined coefficients implies

$$\begin{aligned}
\boldsymbol{\Delta}_a^{P^\delta} &= \left[\delta \boldsymbol{\delta}_w^{P^\delta} \boldsymbol{\delta}_a^{Ew} + \beta \frac{\Gamma}{w_a} \left(\boldsymbol{C}^\delta - \boldsymbol{J}'_{n-1}\boldsymbol{C}\right) \boldsymbol{J}_{n-1} \boldsymbol{\zeta} \boldsymbol{J}'_{n-1} \boldsymbol{\Delta}_a^a\right] \left(\beta^{-1}\boldsymbol{I}_{n-1} - \delta \boldsymbol{\Delta}_a^a\right)^{-1}, \\
\boldsymbol{\delta}_w^{P^\delta} &= \left[\delta \boldsymbol{\Delta}_a^{P^\delta} \boldsymbol{\delta}_w^a + \left(\boldsymbol{ex}^\delta - \boldsymbol{J}'_{n-1}\boldsymbol{ex}\right) + \beta \frac{\Gamma}{w_a} \left(\boldsymbol{C}^\delta - \boldsymbol{J}'_{n-1}\boldsymbol{C}\right) \boldsymbol{J}_{n-1} \boldsymbol{\zeta} \boldsymbol{J}'_{n-1} \boldsymbol{\delta}_w^a\right] \left(\beta^{-1} - \delta \delta_w^{Ew}\right)^{-1}, \\
\boldsymbol{\delta}_\nu^{P^\delta} &= \left[\delta \left(\boldsymbol{\Delta}_a^{P^\delta} \boldsymbol{\delta}_\nu^a + \boldsymbol{\delta}_w^{P^\delta} \delta_\nu^{Ew}\right) - \beta^{-1} \left(\mathbf{1}_n - \boldsymbol{J}'_{n-1}\mathbf{1}_{n-1}\right) + \beta \frac{\Gamma}{w_a} \left(\boldsymbol{C}^\delta - \boldsymbol{J}'_{n-1}\boldsymbol{C}\right) \boldsymbol{J}_{n-1} \boldsymbol{\zeta} \boldsymbol{J}'_{n-1} \boldsymbol{\delta}_\nu^a\right] \times \\
&\quad \left(\beta^{-1} - \delta \rho_1\right)^{-1}.
\end{aligned}$$

The impact effect of a U.S. monetary tightening on consol prices is then

$$\boldsymbol{\delta}_w^{P^\delta} \delta_\epsilon^w + \boldsymbol{\delta}_\nu^{P^\delta}.$$

We again approximate this vector to first order in $\boldsymbol{ex}$, $\boldsymbol{ex}^\delta$ and Γ/w_a around the point with $\boldsymbol{ex} \to 0$, $\boldsymbol{ex}^\delta \to 0$, and $\Gamma/w_a \to 0$. At that point of approximation, and using the results derived in the proof of Proposition 1, we have that

$$\begin{aligned}
\lim_{\boldsymbol{ex}, \boldsymbol{ex}^\delta, \Gamma \to 0} \boldsymbol{\Delta}_a^{P^\delta} &= \mathbf{0}_{n \times (n-1)}, \\
\lim_{\boldsymbol{ex}, \boldsymbol{ex}^\delta, \Gamma \to 0} \boldsymbol{\delta}_w^{P^\delta} &= \mathbf{0}_n, \\
\lim_{\boldsymbol{ex}, \boldsymbol{ex}^\delta, \Gamma \to 0} \boldsymbol{\delta}_\nu^{P^\delta} &= -\left(1 - \beta\delta\rho_1\right)^{-1} \left(\mathbf{1}_n - \boldsymbol{J}'_{n-1}\mathbf{1}_{n-1}\right).
\end{aligned}$$

It follows that

$$\lim_{\boldsymbol{ex}, \boldsymbol{ex}^\delta, \Gamma \to 0} \left[\boldsymbol{\delta}_w^{P^\delta} \delta_\epsilon^w + \boldsymbol{\delta}_\nu^{P^\delta}\right] = -\left(1 - \beta\delta\rho_1\right)^{-1} \left(\mathbf{1}_n - \boldsymbol{J}'_{n-1}\mathbf{1}_{n-1}\right).$$

Moreover,

$$\begin{aligned}
\lim_{\boldsymbol{ex}, \boldsymbol{ex}^\delta, \Gamma \to 0} \frac{d\boldsymbol{\Delta}_a^{P^\delta}}{dex_j} &= \mathbf{0}_{n \times (n-1)}, \ \ \forall j, \\
\lim_{\boldsymbol{ex}, \boldsymbol{ex}^\delta, \Gamma \to 0} \frac{d\boldsymbol{\delta}_w^{P^\delta}}{d\boldsymbol{ex}} &= -\beta \left(1 - \delta(1 - \xi)\right)^{-1} \boldsymbol{J}'_{n-1}, \\
\lim_{\boldsymbol{ex}, \boldsymbol{ex}^\delta, \Gamma \to 0} \frac{d\boldsymbol{\delta}_\nu^{P^\delta}}{d\boldsymbol{ex}} &= \mathbf{0}_{n \times (n-1)}.
\end{aligned}$$

Further using the expression for $\lim_{\boldsymbol{ex}, \Gamma \to 0} \delta_\epsilon^w$ in the proof of Proposition 1, it follows

that

$$\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0}\frac{d\left[\boldsymbol{\delta}_{w}^{P^{\delta}}\delta_{\epsilon}^{w}+\boldsymbol{\delta}_{\nu}^{P^{\delta}}\right]}{d\boldsymbol{ex}}$$
$$=\left(1-\delta(1-\xi)\right)^{-1}(1-\xi)\left(\frac{\sum_{k=2}^{n}b_{ak}}{w_a}\right)\left[\beta^{-1}-\rho_1\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_1)^{-1}\psi\right]\boldsymbol{J}'_{n-1}.$$

Similarly,

$$\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0}\frac{d\boldsymbol{\Delta}_{a}^{P^{\delta}}}{dex_{j}^{\delta}}=\mathbf{0}_{n\times(n-1)},\quad\forall j,$$
$$\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0}\frac{d\boldsymbol{\delta}_{w}^{P^{\delta}}}{d\boldsymbol{ex}^{\delta}}=\beta\left(1-\delta(1-\xi)\right)^{-1}\boldsymbol{I}_n,$$
$$\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0}\frac{d\boldsymbol{\delta}_{\nu}^{P^{\delta}}}{d\boldsymbol{ex}^{\delta}}=\mathbf{0}_{n\times n}.$$

It follows that

$$\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0}\frac{d\left[\boldsymbol{\delta}_{w}^{P^{\delta}}\delta_{\epsilon}^{w}+\boldsymbol{\delta}_{\nu}^{P^{\delta}}\right]}{d\boldsymbol{ex}^{\delta}}$$
$$=-\left(1-\delta(1-\xi)\right)^{-1}(1-\xi)\left(\frac{\sum_{k=2}^{n}b_{ak}}{w_a}\right)\left[\beta^{-1}-\rho_1\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_1)^{-1}\psi\right]\boldsymbol{I}_n.$$

Finally,

$$\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0}\frac{d\boldsymbol{\Delta}_{a}^{P^{\delta}}}{d[\Gamma/w_a]}=(1-\beta\delta)^{-1}\left(\boldsymbol{C}^{\delta}-\boldsymbol{J}'_{n-1}\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}'_{n-1},$$
$$\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0}\frac{d\boldsymbol{\delta}_{w}^{P^{\delta}}}{d[\Gamma/w_a]}=\mathbf{0}_n,$$
$$\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0}\frac{d\boldsymbol{\delta}_{\nu}^{P^{\delta}}}{d[\Gamma/w_a]}=(\beta^{-1}-\rho_1)^{-1}\left(\beta^{-1}-\delta\rho_1\right)^{-1}\left(\boldsymbol{C}^{\delta}-\boldsymbol{J}'_{n-1}\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\eta}'_1(1-\psi).$$

It follows that

$$\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0}\frac{d\left[\boldsymbol{\delta}_{w}^{P^{\delta}}\delta_{\epsilon}^{w}+\boldsymbol{\delta}_{\nu}^{P^{\delta}}\right]}{d[\Gamma/w_a]}=(\beta^{-1}-\rho_1)^{-1}\left(\beta^{-1}-\delta\rho_1\right)^{-1}\left(\boldsymbol{C}^{\delta}-\boldsymbol{J}'_{n-1}\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\eta}'_1(1-\psi).$$

Combining these results, we obtain that to first order in $\boldsymbol{ex}$, $\boldsymbol{ex}^{\delta}$, and Γ/w_a,

$$\begin{aligned}\boldsymbol{\delta}_w^{P^\delta}\delta_\epsilon^w + \boldsymbol{\delta}_\nu^{P^\delta} &= -\left(1-\beta\delta\rho_1\right)^{-1}\left(\mathbf{1}_n - \boldsymbol{J}'_{n-1}\mathbf{1}_{n-1}\right) \\ -(1-\delta(1-\xi))^{-1}(1-\xi)\left(\frac{\sum_{k=2}^n b_{ak}}{w_a}\right)\left[\beta^{-1}-\rho_1\right]^{-1}&\left[1+(\beta^{-1}-1)(1-\rho_1)^{-1}\psi\right]\left(\boldsymbol{ex}^{\delta} - \boldsymbol{J}'_{n-1}\boldsymbol{ex}\right) \\ &+ \left(\beta^{-1}-\rho_1\right)^{-1}\left(\beta^{-1}-\delta\rho_1\right)^{-1}\frac{\Gamma}{w_a}\left(\boldsymbol{C}^{\delta} - \boldsymbol{J}'_{n-1}\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\eta}'_1(1-\psi).\end{aligned}$$

Noting that

$$\mathbf{1}_n - \boldsymbol{J}'_{n-1}\mathbf{1}_{n-1} = \begin{bmatrix} 1 \\ 0 \\ \vdots \\ 0 \end{bmatrix}$$

and again noting that

$$\left(\boldsymbol{C}^{\delta} - \boldsymbol{J}'_{n-1}\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\eta}'_1 = \begin{bmatrix} Cov\left(ex^{\delta}_{1,+1}, \sum_{k=2}^n \eta_{1k} ex_{k,+1}\right) \\ Cov\left(ex^{\delta}_{2,+1} - ex_{2,+1}, \sum_{k=2}^n \eta_{1k} ex_{k,+1}\right) \\ \vdots \\ Cov\left(ex^{\delta}_{n,+1} - ex_{n,+1}, \sum_{k=2}^n \eta_{1k} ex_{k,+1}\right) \end{bmatrix},$$

we obtain the claimed result in the proposition. □

C Additional empirical results

In this appendix we provide additional results accompanying those in section 4.

C.1 Sample periods

Table 4 summarizes the starting months for three-month yields, ten-year yields, and exchange rates used in the empirical analysis. For the G10 economies, the start dates are determined by the latest of April 1991 (when U.S. Treasury yields become available from Bloomberg) and the date the relevant ticker from Bloomberg becomes available. For the emerging markets, the start dates are in some cases pushed further forward to avoid periods of hard pegs, crawling pegs, or "freely falling" exchange rates as classified by Ilzetzki et al. (2019, 2022). We make this choice because such

exchange rate arrangements are not well captured by our model, and to avoid these periods from exerting substantial influence on our estimated comovements.

C.2 Estimating expected excess returns

As described in the main text, we estimate excess currency returns $\{ex_j\}$ by computing average quarterly excess returns for each currency, running a cross-sectional regression of these on the average three-month interest rate differential versus the dollar, and finally using the fitted values as our estimates. We similarly estimate $\{ex_j^\delta\}$ by computing average quarterly excess returns on ten-year bonds in each currency financed by the three-month dollar bond, running a cross-sectional regression of these on the average yield differential between ten-year bonds in that currency and the three-month dollar interest rate, and using the fitted values as our estimates. We use gross returns (not log returns) throughout, consistent with the theory.

We use the fitted values from cross-sectional regressions to reduce the noise in our estimates of excess returns. Figure 11 compares average realized excess returns on currencies and long-term bonds against our estimates of expected excess returns. We see that average interest rate differentials and yield curve slopes indeed do a very good job explaining the cross-section of average excess returns.[57]

C.3 Sensitivity to $\boldsymbol{ex}$ and $\boldsymbol{C}$

In the main text, we demonstrate that a mean-variance efficient portfolio using expected excess returns estimated as above, and the conditional covariance matrix estimated using daily data, is largely a leveraged position in emerging market currencies funded in the dollar for $\Gamma = 1$. We also demonstrate that the conditional responses of exchange rates and long yields to monetary surprises line up with average excess currency and local bond carry returns. Here we provide results investigating the robustness of these results to alternative estimates of $\boldsymbol{ex}$ and $\boldsymbol{C}$.

Expected excess returns In the first row of Table 5 (and in Table 1), we use expected excess returns as constructed in the prior subsection. The second row of

[57] For currencies in particular, this echoes a central message of Hassan and Mano (2019). For a joint analysis of currency and bond risk premia, see Lustig, Stathopoulos, and Verdelhan (2019).

	Government yields		IBOR yields		
	3m	10y	3m	10y	Exchange rate
AUD	Apr-91	Apr-91	Apr-91	Apr-91	Apr-91
CAD	Jun-91	Jun-91	Dec-91	Apr-91	Apr-91
CHF	Feb-94	Feb-94	Apr-91	Apr-91	Apr-91
EUR	Oct-91	Oct-91	Dec-98	Jan-99	Apr-91
GBP	Apr-91	Apr-91	Apr-91	Apr-91	Apr-91
JPY	Sep-92	Sep-92	Apr-91	Apr-91	Apr-91
NOK	Jul-98	Jul-98	Apr-91	Oct-93	Apr-91
NZD	Mar-92	Mar-92	Oct-95	Mar-96	Apr-91
SEK	Feb-94	Feb-94	Apr-91	Apr-91	Apr-91
USD	Apr-91	Apr-91	Apr-91	Apr-91	
BRL*	Mar-07	Mar-07			Oct-99
COP	Aug-03	Oct-04			Aug-92
HUF*	Feb-99	Feb-99			Feb-99
IDR*	Feb-03	Jul-03			Jun-99
INR*	Apr-09	Apr-09			Apr-09
KRW*	Aug-99	Oct-00			Aug-98
MXN*	Sep-02	Feb-03			May-96
MYR*	Aug-05	Aug-05			Aug-05
PEN*	Feb-06	Feb-06			Feb-03
PHP*	Dec-00	Dec-00			Feb-98
PLN*	Apr-98	May-99			Jul-93
THB*	Jun-99	Jun-99			Mar-98
ZAR*	Apr-95	Apr-95			Apr-95

Table 4: starting months for yields and exchange rates used in empirical analysis

Notes: IBOR yields only used in robustness for G10 economies, as described in main text. * denotes emerging markets for which the start dates are pushed forward from the first date of Bloomberg ticker availability to avoid periods of hard pegs, crawling pegs, or "freely falling" exchange rates as classified by Ilzetzki et al. (2019, 2022).

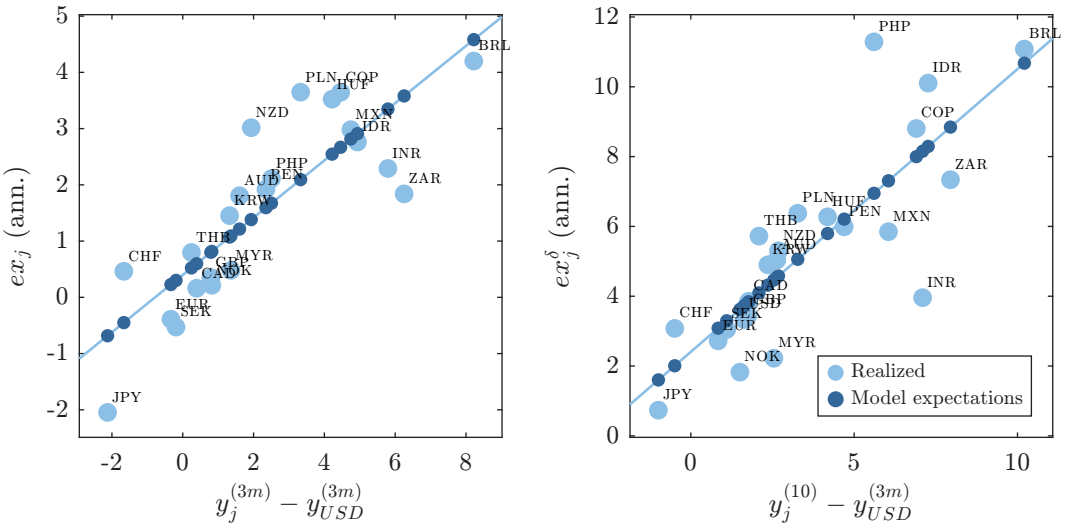


Figure 11: average realized excess returns vs. estimates of expected excess returns

Notes: expectations estimated by running a cross-sectional regression on realized values in each panel and computing fitted values, dropping the constant. Emerging markets depicted with open circles are the smallest six and thus excluded from the quantitative model for computational reasons.

	CHF, EUR, JPY	Other non-U.S. G10	EM	Sum
$\Gamma = 1$	-4.8	-2.2	13.2	6.2
$\Gamma = 1$, realized	-8.7	-3.9	16.8	4.2
$\Gamma = 1$, realized, no 10-year bonds	-4.7	-5.7	16.4	6.0
$\Gamma = 1$, realized, no 10-year bonds, $k = 2$	-4.4	-10.4	23.4	8.7
$\Gamma = 1$, realized, no 10-year bonds, $k = 3$	-5.3	-9.9	22.7	7.6
$\Gamma = 1$, realized, no 10-year bonds, $k = 4$	-6.8	-10.5	25.9	8.6

Table 5: mean-variance efficient portfolio given alternative estimates of $\boldsymbol{ex}$ and $\boldsymbol{C}$

Notes: first row reports baseline results as in Table 1. Second row uses realized average excess returns for $\boldsymbol{ex}$, third row drops 10-year bonds from set of assets, and remaining rows use factor model with k factors for $\boldsymbol{C}$.

Table 5 instead simply uses the realized average excess returns. The basic structure of the mean-variance efficient portfolio remains similar.

Figure 12 shows that the cross-sectional effects of U.S. monetary policy surprises

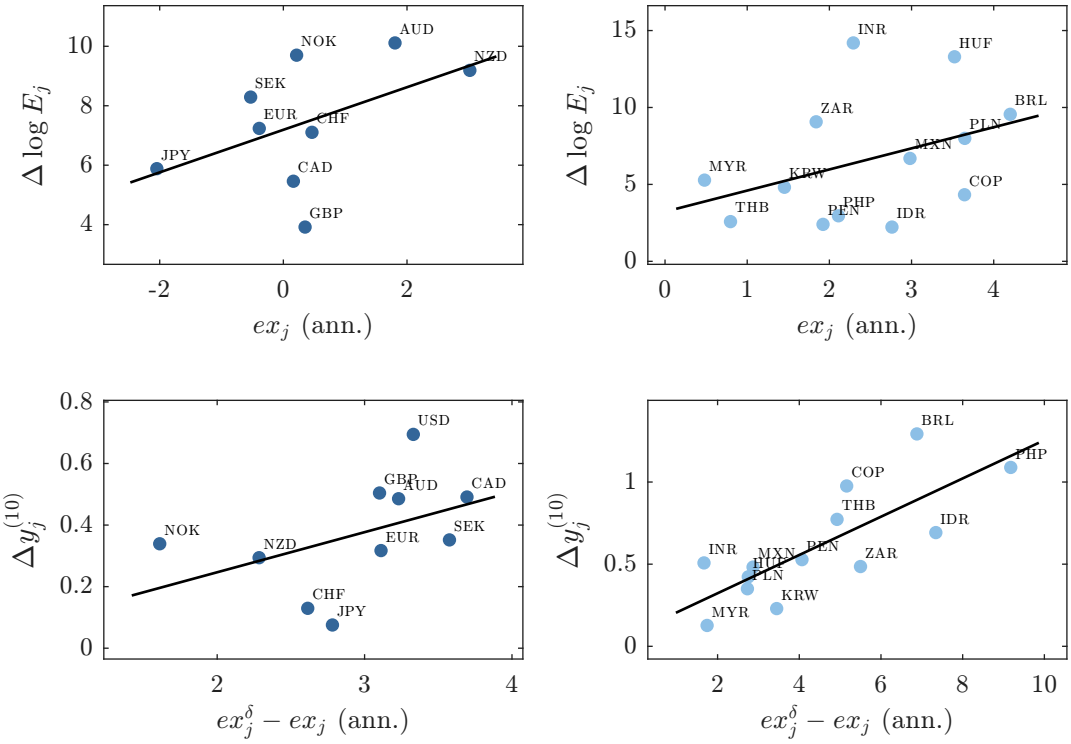


Figure 12: asset price responses to U.S. monetary tightening in data against average realized excess returns

Notes: response to U.S. monetary tightening estimated in data using Bauer and Swanson (2023) measure of monetary surprises and two-day changes in asset prices around surprises. Responses reported for surprise that raises two-year dollar yield by 1*pp*.

are similarly robust when using realized rather than predicted returns along the x-axes. As in Figure 2 in the main text, currencies with higher average realized excess returns depreciate more, and long-dated yields for currencies with higher average realized excess local bond carry returns rise more, following a U.S. tightening. Of course, using realized rather than fitted excess returns implies that the cross-sectional relationships are a bit noisier than in Figure 2.

Factor model In addition to using average realized excess returns, we now revisit the efficient portfolio using a factor-based approach to estimate the covariance matrix of excess returns. We focus for simplicity on the case with no 10-year bonds being traded, so that we can focus on excess currency returns alone. This both simplifies the exposition and allows us to connect to a large literature on risk factors in currency

Cumulative share of $Var(ex_{j,t+1})$ *explained by* k *PCs*					
k	Share				
1	0.450				
2	0.549				
3	0.605				
4	0.659				
Estimated β_j^k					
k	JPY	CHF	NZD	AUD	EM (avg)
1	0.078	0.188	0.279	0.290	0.181
2	-0.235	-0.326	0.016	0.088	0.103
3	-0.131	-0.094	-0.046	-0.017	-0.024
4	-0.381	-0.042	-0.188	-0.142	-0.036

Table 6: principal component analysis of excess currency returns

Notes: last column of second panel reports average of loadings on excess returns for each emerging market currency.

markets (e.g., Lustig, Roussanov, and Verdelhan (2011) and many others).

Suppose excess currency returns load on k factors $\{f_t^1, \ldots, f_t^k\}$ according to

$$ex_{j,t} = \alpha_j + \beta_j^1 f_t^1 + \ldots + \beta_j^k f_t^k + \epsilon_{j,t},$$

where $\epsilon_{j,t}$ is an *iid* Gaussian shock with zero mean and conditional variance σ_j^2. We can stack these to write

$$\boldsymbol{ex}_t = \boldsymbol{\alpha} + \boldsymbol{\beta}\boldsymbol{f}_t + \boldsymbol{\epsilon}_t,$$

where $\boldsymbol{ex}_t$, $\boldsymbol{\alpha}$, and $\boldsymbol{\epsilon}_t$ are $(n-1) \times 1$ vectors, $\boldsymbol{\beta}$ is a $n \times k$ matrix, and $\boldsymbol{f}_t$ is a $k \times 1$ vector. It follows that the conditional covariance matrix of excess returns is

$$\boldsymbol{C} = \boldsymbol{\beta}\boldsymbol{F}\boldsymbol{\beta}' + \boldsymbol{D},$$

where $\boldsymbol{F}$ is conditional covariance matrix of the factors, where we assume constant conditional variances of innovations, and where $\boldsymbol{D} \equiv \text{diag}\left(\sigma_1^2, \ldots, \sigma_{n-1}^2\right)$.

We estimate the covariance matrix in this way using the first k principal components (PCs) of excess currency returns as factors. Table 6 describes the share of

variance in currency returns explained by the PCs as well as the factor loadings β_j^k. As argued in Lustig et al. (2011), the loadings suggest that the first PC essentially captures the average excess return to non-dollar currencies, while the second PC captures the return to the carry trade, with negative loadings on low interest rate currencies like Japan and Switzerland, and positive loadings on high interest rate currencies like Australia, New Zealand, and emerging markets.

The last four rows of Table 5 summarize how this way of estimating the covariance matrix affects the mean-variance efficient portfolio. The third row reports the portfolio shares given average realized excess currency returns and the covariance matrix estimated as described in the main text.[58] The fourth row reports the portfolio shares given average excess returns and the covariance matrix estimated using $k = 2$ factors. Imposing this structure on the covariance matrix leaves the qualitative features of the efficient portfolio unchanged. The remaining rows demonstrate that similar results are obtained using three or four factors as well.

C.4 Evidence against portfolio balance

Figure 13 depicts the same responses of international asset prices to U.S. monetary shocks as in Figure 2, but plots these against covariances of currency and local currency bond carry returns with the (simple) average of currency carry returns that are long foreign currencies, and short the dollar. As demonstrated in Proposition 1, assuming U.S. capital inflows upon a monetary tightening, the portfolio balance channel would imply that currencies with a bigger such covariance should depreciate by less, and the yield on long-dated bonds with a bigger such covariance should rise by less. These predictions clearly do not hold in the data for currencies nor for emerging market bonds.[59] If a U.S. monetary tightening instead implies capital outflows from the U.S., then the response of currencies and emerging market bonds would be consistent with the portfolio balance mechanism. However, the response of G10 bonds would not, as these are essentially unrelated to covariances of local currency bond carry returns with the average currency carry trade.

[58] The only difference from the fourth row of Table 1 is that we use average realized excess currency returns for $\boldsymbol{ex}$, rather than their fitted values on interest rate differentials and yield curve slopes.

[59] Combining the pictures for G10 bonds and EM bonds further undercuts the portfolio balance prediction, since EM bond returns have a higher covariance with the average currency carry trade than G10 bonds, but EM bond yields also rise by more upon a U.S. tightening.

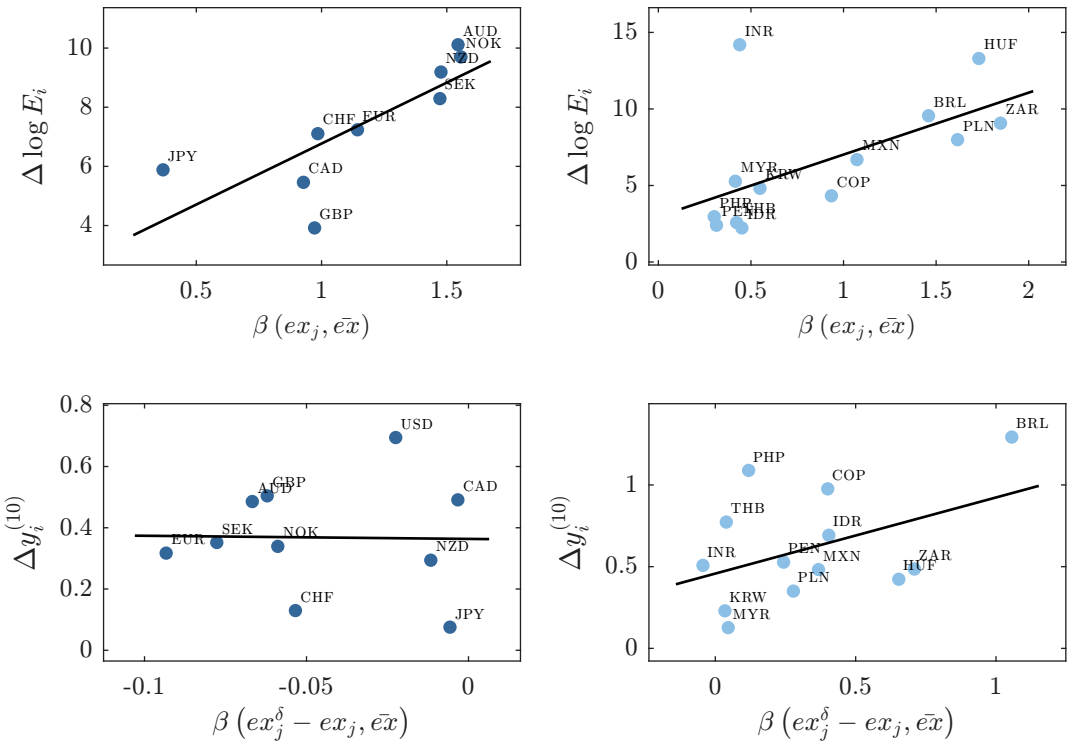


Figure 13: asset price responses to U.S. monetary tightening against covariances with average currency carry return

Notes: response to U.S. monetary tightening estimated in data using Bauer and Swanson (2023) measure of monetary surprises and two-day changes in asset prices around surprises. Responses reported for surprise that raises two-year dollar yield by 1*pp*. $\beta(y,x) \equiv Cov(y,x)/Var(x)$, and $e\bar{x}$ denotes the simple average of currency carry returns.

This evidence also is inconsistent with a portfolio balance channel whereby foreigners' desired position in dollar-denominated bonds $\bar{b}_{j1}$ for $j > 1$ rises with the U.S. nominal interest rate. In that case, arbitrageurs would hold fewer dollar-denominated and more foreign-denominated assets on impact of a U.S. tightening, so that the portfolio balance channel would imply that currencies (long-term bond prices) that covary more with the average currency carry trade return should depreciate by more (fall by more) following a U.S. tightening. This is consistent with the response of currencies and emerging market bonds, but not G10 bonds.

D Additional quantitative results

In this appendix we provide additional results accompanying those in section 5.

D.1 Baseline calibration

The main text described the overall calibration strategy for parameters disciplined after solving the model, as summarized in Table 3. Here we provide detail on these calibrated parameter values and the associated model moments.

The model features the G10 economies and 13 emerging markets featured in our empirical analysis, as well as one "rest of world" (ROW) which is a residual ensuring market clearing given the targets for all other economies.

Table 7 provides the population parameters in each economy, which are targeted to match relative GDP from the IMF International Financial Statistics. The ROW has population such that world GDP is 3.5 times U.S. GDP.

Table 8 provides the discount factors in each economy and consol positions of U.S. households, which are targeted to match the three-month dollar interest rate and excess returns estimated using Bloomberg data. As described in section 5, we further set $\bar{b}_{1j}$ for $j > 1$ so that U.S. households take the opposite side of arbitrageurs in one-period non-dollar bonds, and thus NFAs are all zero. We assume the ROW is in financial autarky so we do not specify its discount factor, and we assume U.S. households have zero positions in its currency.

Table 9 provides the taste parameters governing trade shares, which are targeted to match expenditure shares by origin from the OECD Trade in Value Added dataset. For brevity we focus in this table on taste parameters for imports from and exports to the U.S., though there is an entire matrix of such parameters corresponding to all origins and destinations. We set the ROW's taste parameters so that

$$\sum_{j=1}^{n} \zeta_j \eta_{jk} = \zeta_k,$$

for all k, which implies (together with the above parameters) that all real exchange rates are one in steady-state.

Finally, Table 10 summarizes how the model matches the currency invoicing shares reported in Boz et al. (2025). The available data reports invoicing shares among

		$q_j^{-1} y_j / y_1$	
Country	ζ_j	Target	Achieved
USD	1	normalization	
AUD	0.06	0.06	0.06
CAD	0.09	0.09	0.09
CHF	0.04	0.04	0.04
EUR	0.73	0.73	0.73
GBP	0.16	0.16	0.16
JPY	0.37	0.37	0.37
NOK	0.02	0.02	0.02
NZD	0.01	0.01	0.01
SEK	0.03	0.03	0.03
BRL	0.09	0.09	0.09
COP	0.02	0.02	0.02
HUF	0.01	0.01	0.01
IDR	0.05	0.05	0.05
INR	0.11	0.11	0.11
KRW	0.07	0.07	0.07
MXN	0.07	0.07	0.07
MYR	0.02	0.02	0.02
PEN	0.01	0.01	0.01
PHP	0.01	0.01	0.01
PLN	0.03	0.03	0.03
THB	0.02	0.02	0.02
ZAR	0.02	0.02	0.02
ROW	0.47	0.47	0.47

Table 7: baseline ζ_j parameters

Notes: data targets estimated using IMF International Financial Statistics over 1991 Q2 – 2024 Q1.

		$52 i_j$			$52 \log((1+\delta P_j^\delta)/P_j^\delta)$	
Country	β_j^{52}	Target	Achieved	$\bar{b}_{1j}^\delta/(52 y_1)$	Target	Achieved
USD	0.9901	1.00%	1.00	4.19%	4.73%	4.73%
AUD	0.9781	2.21%	2.21	-0.19%	5.54%	5.54%
CAD	0.9842	1.60%	1.60	1.15%	4.84%	4.84%
CHF	0.9945	0.55%	0.55	2.57%	3.01%	3.01%
EUR	0.9878	1.23%	1.23	1.18%	4.08%	4.08%
GBP	0.9821	1.80%	1.80	-0.65%	4.83%	4.82%
JPY	0.9968	0.32%	0.32	5.45%	2.60%	2.60%
NOK	0.9820	1.82%	1.82	0.27%	4.62%	4.62%
NZD	0.9765	2.38%	2.38	0.57%	5.58%	5.58%
SEK	0.9871	1.30%	1.30	1.19%	4.30%	4.29%
BRL	0.9457	5.58%	5.58	-1.63%	11.68%	11.66%
COP	0.9640	3.67%	3.67	-0.08%	9.00%	8.99%
HUF	0.9652	3.54%	3.54	-2.60%	6.80%	6.79%
IDR	0.9617	3.91%	3.91	0.06%	9.29%	9.28%
INR	0.9575	4.35%	4.35	-1.23%	9.15%	9.15%
KRW	0.9795	2.07%	2.07	0.74%	5.32%	5.31%
MXN	0.9626	3.81%	3.81	-0.35%	8.31%	8.30%
MYR	0.9793	2.09%	2.09	4.85%	5.47%	5.46%
PEN	0.9744	2.59%	2.59	-2.09%	7.21%	7.20%
PHP	0.9736	2.67%	2.67	-0.60%	7.95%	7.94%
PLN	0.9696	3.09%	3.09	-0.51%	6.06%	6.05%
THB	0.9849	1.52%	1.52	3.21%	5.10%	5.09%
ZAR	0.9552	4.58%	4.58	-0.21%	9.84%	9.84%

Table 8: baseline β_j and $\bar{b}_{1j}^\delta$ parameters

Notes: data targets estimated using Bloomberg data over 1991 Q2 – 2024 Q1. Note that for all $j > 1$, we obtain the targeted i_j as $i_1 + ex_j$, where ex_j is estimated as described in section 5. We obtain the targeted $(1+\delta P_j^\delta)/P_j^\delta$ as $1 + i_1 + ex_j^\delta$, where ex_j^δ is estimated as described in section 5. Rest of World is assumed to be in financial autarky so we do not specify β_{ROW} and assume $\bar{b}_{1ROW}^\delta = 0$.

		$P_{j1}c_{j1}/(P_jc_j)$			$P_{1j}c_{1j}/(P_1c_1)$	
Country	η_{j1}	Target	Achieved	η_{1j}	Target	Achieved
USD	90.33%	90.33%	90.33%	90.33%	90.33%	90.33%
AUD	2.19%	2.19%	2.19%	0.09%	0.09%	0.09%
CAD	10.11%	10.11%	10.11%	1.12%	1.12%	1.12%
CHF	2.57%	2.57%	2.57%	0.19%	0.19%	0.19%
EUR	1.89%	1.89%	1.89%	1.71%	1.71%	1.71%
GBP	2.44%	2.44%	2.44%	0.53%	0.53%	0.53%
JPY	1.79%	1.79%	1.79%	0.59%	0.59%	0.59%
NOK	1.78%	1.78%	1.78%	0.06%	0.06%	0.06%
NZD	2.42%	2.42%	2.42%	0.02%	0.02%	0.02%
SEK	1.88%	1.88%	1.88%	0.07%	0.07%	0.07%
BRL	1.90%	1.90%	1.90%	0.15%	0.15%	0.15%
COP	4.26%	4.26%	4.26%	0.07%	0.07%	0.07%
HUF	1.74%	1.74%	1.74%	0.01%	0.01%	0.01%
IDR	1.69%	1.69%	1.69%	0.11%	0.11%	0.11%
INR	1.29%	1.29%	1.29%	0.30%	0.30%	0.30%
KRW	2.84%	2.84%	2.84%	0.24%	0.24%	0.24%
MXN	9.69%	9.69%	9.69%	0.65%	0.65%	0.65%
MYR	3.35%	3.35%	3.35%	0.07%	0.07%	0.07%
PEN	4.28%	4.28%	4.28%	0.03%	0.03%	0.03%
PHP	2.74%	2.74%	2.74%	0.05%	0.05%	0.05%
PLN	1.44%	1.44%	1.44%	0.03%	0.03%	0.03%
THB	3.01%	3.01%	3.01%	0.10%	0.10%	0.10%
ZAR	1.70%	1.70%	1.70%	0.03%	0.03%	0.03%
ROW	9.30%	9.30%	9.30%	3.43%	3.43%	3.43%

Table 9: baseline η_{j1} and η_{1j} parameters

Notes: data targets estimated using OECD Trade in Value Added dataset over 1995 – 2022.

imports in and exports from each country, not each trading pair. Moreover, the data only reports shares in the dollar, euro, home currency, or other currency.[60,61] We thus assume that between destination j and origin k, there are four sets of firms: those that price in local currency ($m = j$), producer currency ($m = k$), dollars ($m = 1$), and euros ($m = e$). We assume that invoicing shares reflect the product of fixed effects for each destination-currency and origin-currency pair:

$$\theta_{jk}^m = \frac{a_j^m b_k^m}{\sum_{r\in\{j,k,1,e\}} a_j^r b_k^r}, \;\; m \in \{j, k, 1, e\},$$

with $\theta_{jk}^m = 0$ for $m \notin \{j, k, 1, e\}$. We estimate $\{a_j^m, b_k^m\}$ (up to scalar normalizations) to minimize the distance between the data and model-implied invoicing shares, weighted by each country's size.[62] For brevity, Table 10 reports the average share of imports in and exports from each country that are invoiced in dollars, both in the data and the model. Dollar pricing is particularly prevalent among emerging markets relative to the G10 economies, and the calibration matches these moments well.

D.2 Identification of key parameters

We now describe in further detail how Γ and ξ in our baseline calibration are identified by the cross-section of asset price responses to a U.S. monetary tightening. Focusing on the G10 asset prices used in estimation, Figure 14 depicts how the response of exchange rates and bond yields would differ under alternative values of these parameters.

The left panels depict how Γ, governing arbitrageurs' leverage and thus revaluation of wealth upon a dollar appreciation, is identified. Consistent with Propositions 1 and 2, at a higher value of Γ and thus lower leverage, the model would imply a much smaller depreciation for currencies offering high average excess returns, and smaller increase in yields for bonds offering high average local currency excess returns, counterfactual to the data.

[60] The data also reports renminbi shares, which we include in "other" since China is in ROW.

[61] For Euro area countries, the euro invoicing share in Boz et al. (2025) includes trade with other Euro area countries. Since we treat the Euro area countries as a single bloc in our calibration, we use the euro share of extra-Euro-area goods trade from ECB (2026).

[62] We assume that in the data, imports to country j invoiced in currencies besides the dollar, euro, or home currency are invoiced in the producer's currency, and that exports from country j invoiced in currencies besides the dollar, euro, or home currency are invoiced in the local currency.

	$\frac{\sum_{k\neq j}\theta^1_{jk}P_{jk}c_{jk}}{\sum_{k\neq j}P_{jk}c_{jk}}$		$\frac{\sum_{k\neq j}\theta^1_{kj}\zeta_k E_k^{-1}P_{kj}c_{kj}}{\sum_{k\neq j}\zeta_k E_k^{-1}P_{kj}c_{kj}}$	
Country	Target	Achieved	Target	Achieved
USD	94.6%	94.6%	95.9%	95.9%
AUD	53.8%	53.8%	80.0%	80.0%
CAD	70.0%	70.0%	70.0%	70.0%
CHF	35.5%	35.5%	20.8%	20.7%
EUR	54.4%	54.4%	38.1%	38.2%
GBP	38.2%	38.2%	30.0%	30.0%
JPY	70.5%	70.4%	50.3%	50.2%
NOK	23.9%	23.9%	52.5%	52.4%
NZD	49.2%	49.2%	57.9%	57.8%
SEK	29.9%	29.9%	26.6%	26.7%
BRL	82.8%	82.8%	94.7%	94.7%
COP	98.3%	98.3%	98.3%	98.3%
HUF	20.3%	20.4%	11.5%	11.5%
IDR	81.1%	81.1%	93.6%	93.6%
INR	86.6%	86.6%	86.7%	86.6%
KRW	80.6%	80.5%	84.8%	84.7%
MXN	91.2%	91.2%	96.8%	96.7%
MYR	81.5%	81.4%	86.0%	86.0%
PEN	92.6%	92.6%	97.3%	97.3%
PHP	91.1%	91.0%	97.6%	97.6%
PLN	28.8%	28.9%	20.3%	20.4%
THB	78.7%	78.7%	80.0%	79.9%
ZAR	42.6%	42.6%	42.6%	42.6%
ROW	58.1%	57.1%	71.0%	70.0%

Table 10: baseline θ^1_{jk} parameters

Notes: data targets estimated using invoicing shares in Boz et al. (2025) averaged over 2000 through 2023. Invoicing shares for Euro area correspond to France, adjusted to reflect invoicing of only trade with non-Euro area counterparts using data in ECB (2026). Given missing data, we set invoicing shares in Mexico to be the same as an average of Brazil, Colombia, and Peru, and we assume invoicing shares for exports in Canada, Colombia and South Africa are same as for imports. Finally, ROW invoicing shares are set to size-weighted average for all countries besides U.S. and Euro area.

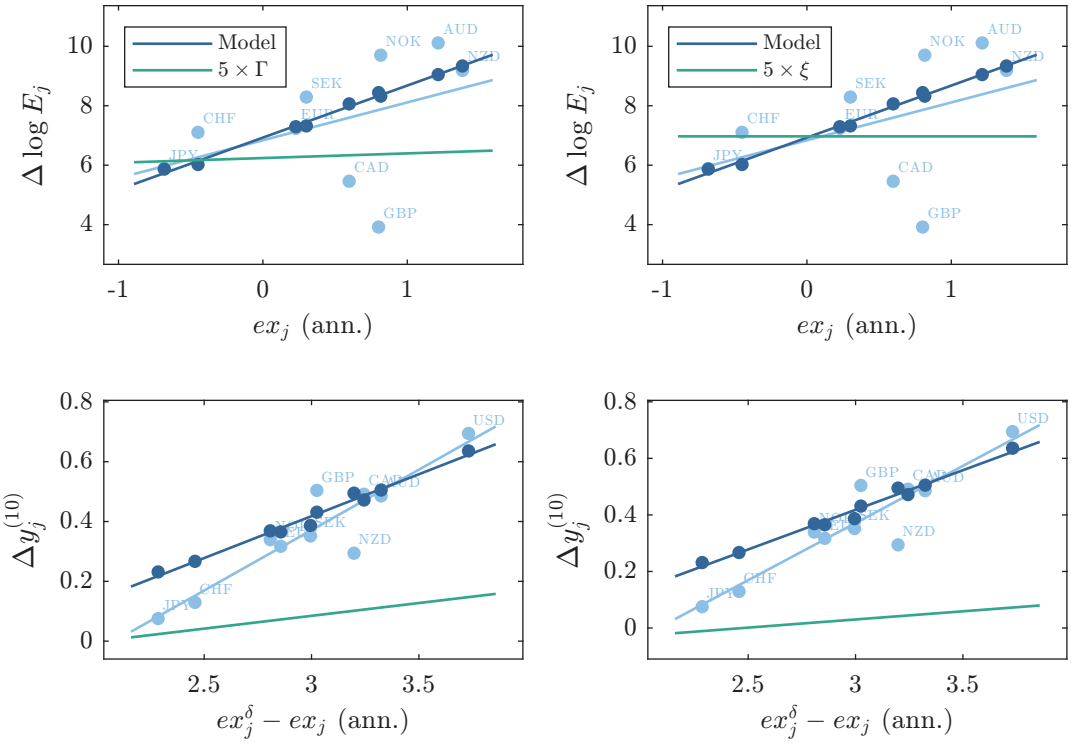


Figure 14: identification of Γ and ξ

Notes: both model and data responses reported for surprise that raises two-year dollar yield by 1*pp*.

The right panels similarly demonstrate how ξ, governing arbitrageurs' exit rate ξ and thus persistence of arbitrageurs' aggregate wealth, is identified. We estimate $\xi = 0.02$, so the expected lifetime of an individual arbitrageur is approximately eight months. Consistent with Propositions 1 and 2, at a higher value of ξ and thus lower persistence of arbitrageur wealth, the model would imply a smaller depreciation for currencies offering high average excess returns and smaller increase in yields for bonds offering high average local currency excess returns. Comparing these propositions, we can also see that ξ is identified from arbitrageur leverage by the relative response of long yields versus exchange rates, provided that $\delta < 1$.

D.3 Alternative calibration matching NFAs

The baseline calibration features zero net foreign asset positions for each country, because the U.S. household takes the other side of the arbitrageur in all assets. This

is a useful benchmark because it eliminates any cross-border wealth effects from asset revaluations. In this section, we study a richer calibration matching net foreign asset positions in the data.

In particular, we assume the U.S. household only takes a non-zero position in the dollar consol, $\bar{b}^{\delta}_{11}$, but zero positions in other currencies' short or consol bonds, $\bar{b}_{1j} = \bar{b}^{\delta}_{1j} = 0$ for $j > 1$. Households in all other countries actively trade their own short bond and hold non-zero positions in their own consol bond and the dollar short bond, $\bar{b}^{\delta}_{jj}$ and $\bar{b}_{j1}$ for $j > 1$, while holding zero positions in other currencies' short or consol bonds. We use $\bar{b}^{\delta}_{jj}$ to match the excess return on long-term bonds in currency j, and $\bar{b}_{j1}$ to match country j's average net foreign asset position over the maintained 1991-2024 period from Milesi-Ferretti (2026).

We keep all other parameters the same as in the baseline calibration (such as Γ and ξ) to facilitate comparison between them. Figure 15 demonstrates that the asset price responses to a U.S. tightening are comparable to those in Figure 4. Similarly, Figure 16 shows that the real effects on G10 and EM countries are similar to those in Figure 8. Taken together, this indicates that the cross-country wealth revaluation induced by a U.S. tightening is second-order relative to the global change in risk pricing operating through arbitrageurs' balance sheet.

To rationalize the fact that arbitrageurs are short the dollar even though the rest of the world has a positive net foreign asset position vis-à-vis the U.S., this calibration requires that the rest of the world holds dollar-denominated assets in excess of their borrowing in local currencies (that is, $\sum_{j>1} \zeta_j \bar{b}_{j1} > 0$). This is consistent with estimates of the currency composition of NFAs (e.g., Allen and Juvenal (2025)). Notably, the magnitude of dollar-denominated NFAs are comparable between model and data: the rest of the world's dollar-denominated NFA is 27% of annual U.S. GDP in the model, as compared to 65% in the data from Allen and Juvenal (2025). This is true even though our calibration approach did not target the magnitude of gross positions: recall that Γ (and thus arbitrageur leverage) was estimated to match the asset price responses to a U.S. monetary tightening, and ω_a (and thus arbitrageurs' wealth) was calibrated to fit the U.S. yield response to QE1. Given the targeted NFAs the calibration could have implied extreme values for countries' dollar vs. non-dollar NFAs, but instead these gross positions are of comparable magnitudes as the data.

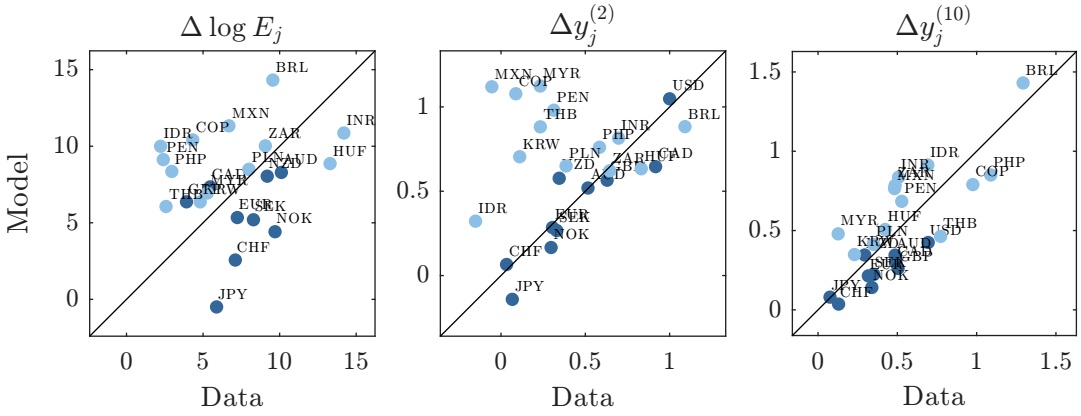


Figure 15: asset price responses to U.S. monetary tightening in calibration targeting NFAs

Notes: data responses reported for surprise that raises two-year dollar yield by 1*pp*, and model responses reported for same $\nu_{1,t}$ innovation depicted in Figure 4.

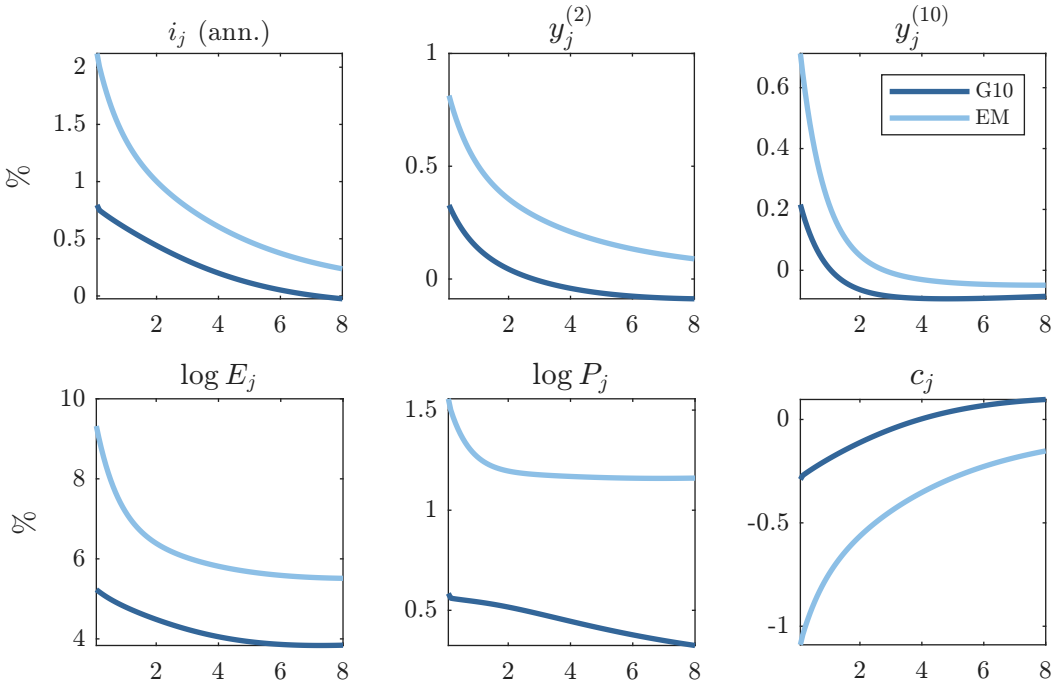


Figure 16: impulse responses to U.S. monetary tightening in calibration targeting NFAs

Notes: responses reported for same $\nu_{1,t}$ innovation depicted in Figure 4. x-axis denotes quarters since the shock.

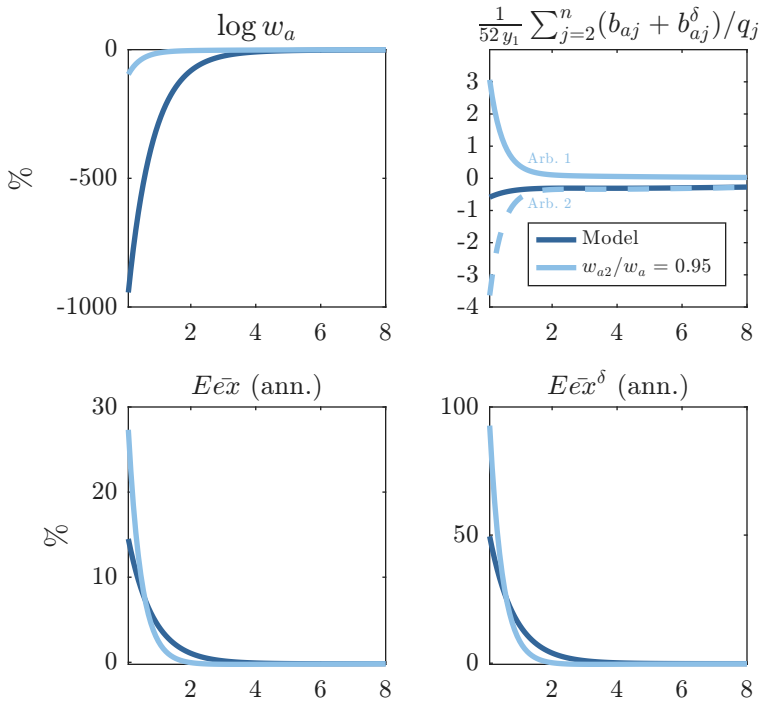


Figure 17: response of arbitrageur wealth and expected excess returns

Notes: responses reported for same $\nu_{1,t}$ innovation depicted in Figure 4. x-axis denotes quarters since the shock.

D.4 Extension with value-at-risk arbitrageurs

As described in the main text, the quick (endogenous) recapitalization of arbitrageurs explains why we need high leverage to match the asset price responses to a U.S. monetary tightening. The first panel of Figure 17 depicts the impulse response of arbitrageur wealth to the same shock depicted in Figure 4. While arbitrageur log-wealth falls by a dramatic 944*pp*, it rapidly rebounds to be only 8*pp* below its steady-state value by one year after the shock.[63] We thus require a large initial decline in wealth — and thus high leverage — to generate a sufficient increase in expected excess returns, and thus response of asset prices on impact.

We now provide more details on the extension of the model with a second set of arbitrageurs that face the value-at-risk constraint (10), which allows us to match the

[63]While these responses are still large, note that we are scaling the U.S. tightening to raise the two-year dollar yield by 1*pp*, an order of magnitude larger than observed surprises in the data.

same asset price responses to a U.S. tightening without requiring such high leverage. We use similar vector notation as in (8) for parsimony.

It is first straightforward to show that the solution to these arbitrageurs' problem (9)-(10), given analogs of (1)-(3), is

$$E_t \boldsymbol{ex}_{t+1} = \left(\frac{1}{2\kappa} E_t \boldsymbol{ex}'_{t+1} \boldsymbol{C}_t^{-1} E_t \boldsymbol{ex}_{t+1} \right)^{1/2} \frac{1}{w_{a_2,t}} \boldsymbol{C}_t \left(\boldsymbol{q}_t^{-1} \boldsymbol{b}_{a_2,t} \right), \tag{45}$$

where $\boldsymbol{q}_t^{-1} \boldsymbol{b}_{a_2,t}$ again denotes the column vector of $\{q_{j,t}^{-1} b_{a_2 j,t}, q_{j,t}^{-1} b_{a_2 j,t}^{\delta}\}$.

Analogous to our assumption that $\gamma(\sigma)\sigma^2 \to \Gamma > 0$ as the perturbation parameter $\sigma \to 0$, we assume that $\kappa(\sigma)/\sigma^2 \to K > 0$. It follows that these arbitrageurs' steady-state portfolio solves

$$\boldsymbol{ex} = \left(\frac{1}{2K} \boldsymbol{ex}' \boldsymbol{C}^{-1} \boldsymbol{ex} \right)^{1/2} \frac{1}{w_{a_2}} \boldsymbol{C} \left(\boldsymbol{q}^{-1} \boldsymbol{b}_{a_2} \right).$$

If K in particular satisfies

$$\left(\frac{1}{2K} \boldsymbol{ex}' \boldsymbol{C}^{-1} \boldsymbol{E}_t \boldsymbol{ex} \right)^{1/2} = \Gamma, \tag{46}$$

these arbitrageurs will choose the same steady-state portfolio shares as the mean-variance arbitrageurs in (8). For simplicity, we make this assumption in what follows.

The first-order approximation of (45) around the steady-state is then

$$E_t \hat{\boldsymbol{ex}}_{t+1} = -\boldsymbol{ex} \hat{w}_{a_2,t} + \boldsymbol{ex} \left(\boldsymbol{ex}' \boldsymbol{C}^{-1} \boldsymbol{ex} \right)^{-1} \boldsymbol{ex}' \boldsymbol{C}^{-1} E_t \hat{\boldsymbol{ex}}_{t+1} + \frac{\Gamma}{w_{a_2}} \boldsymbol{C} \widehat{\boldsymbol{q}_t^{-1} \boldsymbol{b}_{a_2,t}}.$$

This looks the same as for the mean-variance arbitrageurs except for the second term on the right-hand side, which summarizes how these arbitrageurs must additionally adjust their positions to respect the value-at-risk constraint (10). Crucially, note that $\left(\boldsymbol{ex}' \boldsymbol{C}^{-1} \boldsymbol{ex} \right)^{-1} \boldsymbol{ex}' \boldsymbol{C}^{-1} E_t \hat{\boldsymbol{ex}}_{t+1}$ is a scalar, so that assets load heterogeneously on this effect according to $\boldsymbol{ex}$. Intuitively, a tighter value-at-risk constraint is analogous to a higher price of risk.

Lastly, we modify the asset market clearing conditions to include the value-at-risk arbitrageurs along with the standard mean-variance arbitrageurs. It is then straightforward to solve for the first-order approximation of all prices and policies as a function of endogenous and exogenous state variables as before.

Figure 5 in the main text demonstrated how, at higher values of w_{a_2}/w_a (achieved by the arbitrageurs' relative endowments, ω_{a_2} and ω_{a_1}), we require higher Γ and thus lower arbitrageur leverage in non-dollar currencies to match the same asset pricing data conditional on a U.S. monetary shock. As an example for $w_{a_2}/w_a = 0.95$, Figure 18 depicts the asset price responses in this calibration, which are comparable to those in the baseline model (in Figure 4), and thus still comparable to the data. Figure 19 demonstrates the resulting effects of a U.S. tightening on real quantities are also comparable to those in the baseline model (in Figure 8). But, as is evident from Figure 17, the lower leverage of arbitrageurs implies a much smaller decline in their wealth upon the shock. Expected excess returns still rise by a comparable amount as the baseline case — in fact, more on impact, albeit less persistently — because mean-variance arbitrageurs must absorb risky bonds as value-at-risk arbitrageurs shed theirs. The top right panel illustrates this using the dynamic response in arbitrageurs' total position in non-dollar bonds, scaled by annual U.S. output.

D.5 Exchange rate targeting in the monetary policy rule

In the main text, we studied how countries might use foreign exchange intervention to mitigate the effects a U.S. tightening on their external risk premium. Here we instead study a modified conventional policy rule.

Focusing on Brazil as in the main text, Figure 20 adds a term targeting the (real) exchange rate to its monetary policy rule,

$$\begin{aligned}\log(1+i_{BRL,t}) = \log(1+i_{BRL}) + \rho_i \left(\log(1+i_{BRL,t-1}) - \log(1+i_{BRL})\right) \\ + (1-\rho_i)\left(\phi_\pi \pi_{BRL,t} + \phi_{q,BRL}\left(\log q_{BRL,t} - \log q_{BRL}\right)\right).\end{aligned}$$

With a positive value $\phi_{q,BRL}$, Brazil raises its policy rate by more in response to a U.S. tightening.[64] As a result, it experiences a smaller depreciation and smaller inflation, but at the cost of a more severe decline in consumption.

[64] In this exercise, we keep the covariance matrix $\boldsymbol{C}$ unchanged. Jeanne and Rose (2002), Itskhoki and Mukhin (2023), Caballero, Caravello, and Simsek (2025), and Caballero and Simsek (2026) emphasize how changes in the policy rule affect these second moments and thereby risk pricing.

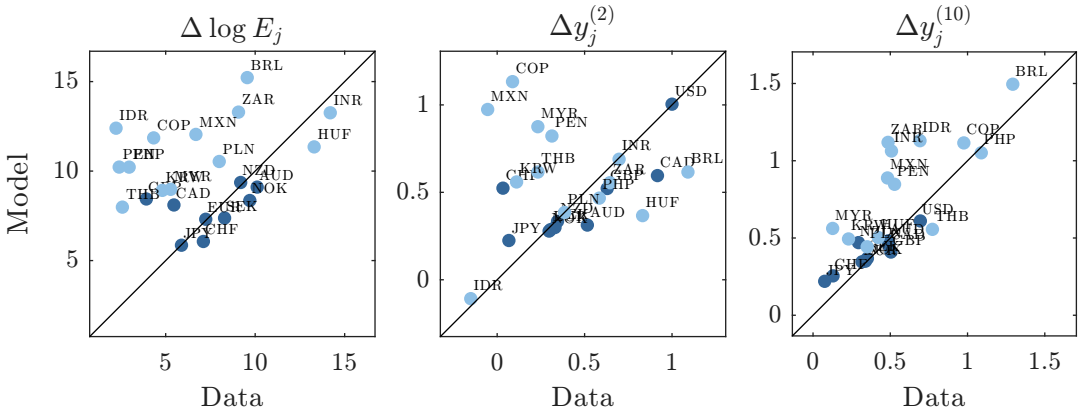


Figure 18: asset price responses to U.S. tightening given $w_{a_2}/w_a = 0.95$

Notes: data responses reported for surprise that raises two-year dollar yield by 1*pp*, and model responses reported for same $\nu_{1,t}$ innovation depicted in Figure 4.

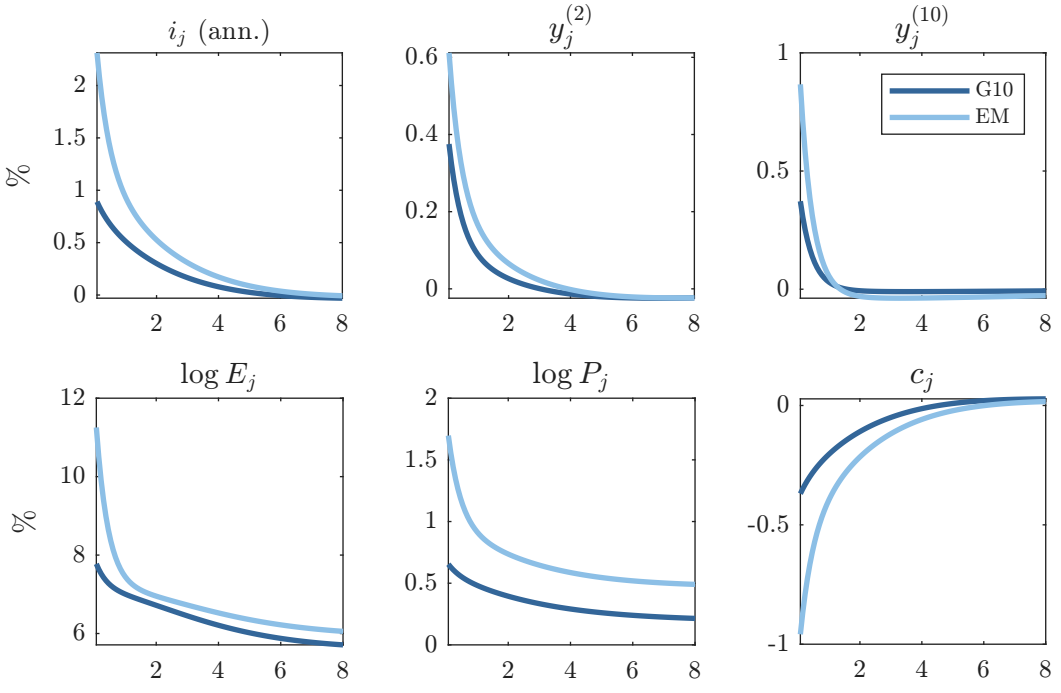


Figure 19: impulse responses to U.S. tightening given $w_{a_2}/w_a = 0.95$

Notes: responses reported for same $\nu_{1,t}$ innovation depicted in Figure 4. x-axis denotes quarters since the shock.

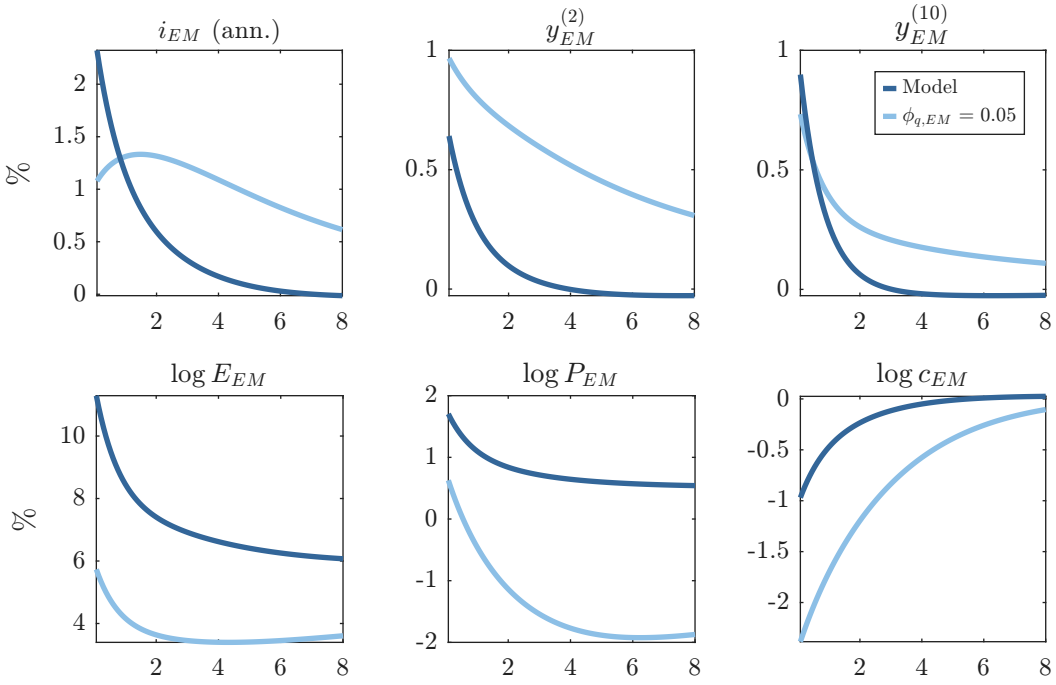


Figure 20: effects of alternative monetary policy response to U.S. monetary tightening

Notes: responses reported for same $\nu_{1,t}$ innovation depicted in Figure 4. x-axis denotes quarters since the shock.

D.6 Price impact of foreign exchange intervention

We finally provide additional results on the price impact of foreign exchange intervention in our model. The results which follow echo ideas emphasized in Kekre and Lenel (2026), which studies the asset price consequences of a global portfolio shift out of dollar-denominated bonds around the April 2025 tariff announcement.

As is evident from (5), there are two key parameters governing the price impact of portfolio flows on arbitrageurs' balance sheet: Γ/w_a, governing the impact on excess returns, and the persistence of the change in portfolios, governing the impact on asset prices. We disciplined Γ (governing arbitrageurs' leverage) to match the asset price responses to U.S. monetary surprises. Given an assumption on the persistence of QE1, we disciplined w_a to match the price impact of QE1 in the data.

Had we assumed a lower persistence of QE1, we would have calibrated a lower w_a to match the same price impact of QE1, and thus higher price impact of Brazil's foreign exchange intervention in Figure 9. Holding fixed w_a and the magnitude of

		$\rho_{\bar{b}_{BRL1}}$ (ann.)		
w_a	half-life QE1	0.06	0.25	0.50
0.06	2 yr	-22%	-13%	-7%
0.11	4 yr	-40%	-22%	-13%
0.14	6 yr	-52%	-28%	-16%

Table 11: required intervention to match $\Delta \log E_{BRL,t}$ around U.S. tightening

Notes: as we vary w_a and $\rho_{\bar{b}_{BRL1}}$, each cell re-estimates $d\bar{b}_{BRL1,t}$ to match $\Delta \log E_{BRL,t}$ around U.S. tightening as in Figure 9. Values in each cell report $d\bar{b}_{BRL1,t}/(52y_{BRL})$.

the foreign exchange intervention, a higher assumed persistence of the intervention would also imply higher price impact. We can also invert these comparative statics: to match a given price impact, a smaller-sized intervention would be needed with lower w_a or a more persistent intervention.

Table 11 quantifies these comparative statics. Each cell reports the size of the required Brazilian foreign exchange intervention to match the same price impact on the exchange rate as depicted in Figure 9. The center cell corresponds to the intervention in that figure. At this range of parameters, selling dollar bonds by 7-52% of annual Brazilian GDP is required to rationalize the Brazilian real response to a U.S. tightening raising the two-year dollar yield by $1pp$.