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

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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 study its real implications. In a future with higher dollar interest rates, arbitrageurs' dollar funding and U.S. monetary policy spillovers are dampened.

JEL codes: E44, E52, F31, G15

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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 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. We then 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 in each destination market subject to adjustment costs, so monetary policy is non-neutral. 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 face adjustment costs in trading 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 can 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, if households' intertemporal elasticity of substitution (IES) differs from one, 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 will 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 15 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 then 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 in generating cross-country variation in response to U.S. policy surprises. The heterogeneous effects on risk premia are reflected in heterogeneous spillovers to the real economy: emerging markets experience larger inflation, a larger increase in policy rates, and a larger reduction in consumption than other G10 economies. Consistent with prior research on sudden stops, such economies can mitigate the reduction in consumption by employing a monetary policy response that permits the local currency to depreciate by more, or by selling dollar bonds through foreign exchange intervention. The latter response is particularly effective because it addresses the change in risk premia directly. Having disciplined the price impact of flows using the evidence from quantitative easing announcements, we find that a sale of short-term dollar bonds by 20% of annual GDP by Brazil, a representative emerging market in our analysis, would appreciate the Brazilian real by roughly 1*pp*, reduce its 10-year yield by roughly 0.1*pp*, and stimulate consumption by roughly 0.025*pp*.

Looking forward, the model implies that a future with 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 second moments unchanged, the mean-variance efficient portfolio features a smaller long position in non-dollar currencies. Quantitatively, a 0.5*pp* higher dollar interest rate (modeled as arising from lower U.S. household patience) reduces arbitrageurs' portfolio share in non-dollar bonds by 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 smaller domestic inflation and 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 (2025)), 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 the global price of risk.

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 (2015), Miranda-Agrippino and Rey (2020), 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 New Keynesian models (e.g.,

[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.

Aghion, Bacchetta, and Banerjee (2007), Gertler, Gilchrist, and Natalucci (2007), and Aoki, Benigno, and Kiyotaki (2021)). Georgiadis, Mueller, and Schumann (2023) and Akinci and 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.

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)). Our result that arbitrageurs will be dollar funded because the excess return on dollar-funded positions is high relative to its risk complements CIP-based measures of dollar convenience (Jiang, Krishnamurthy, and Lustig (2021), Du, Keerati, and Schreger (2025)). 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. That said, we believe our approach can be fruitfully combined with a growing literature estimating international portfolios using granular data.[2] 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 (2024) on the one hand, and Gourinchas, Ray, and Vayanos (2024) 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.[3] A key difference from all of these papers is that the risk-bearing capacity of

[2]See, for instance, Maggiori, Neiman, and Schreger (2020), Koijen and Yogo (2022), Du, Fontana, Jakubik, Koijen, and Shin (2023), Hacioglu-Hoke, Ostry, Rey, Rousset Planat, Stavrakeva, and Tang (2024), Rey and Stavrakeva (2024), and Dao, Gourinchas, and Itskhoki (2025), among others.

[3]In contemporaneous and complementary work, Caballero and Simsek (2026) also study an open economy general equilibrium model in which arbitrageurs trade multiple risky assets. They focus on

arbitrageurs is endogenous, owing to the endogenous evolution of their wealth. The important role of arbitrageur wealth echoes themes from Brunnermeier and Sannikov (2012) and He and Krishnamurthy (2013), and is at the core of our explanation for why U.S. monetary policy affects the global price of risk.[4]

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 (2025)). 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 (2024)). 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 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.

how the monetary policy rule can affect the propagation of exogenous risk premium shocks, while we study how monetary policy (particularly in the U.S.) propagates through risk premia.

[4] In this sense, our model 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.

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

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

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.[5] A one-period bond in currency k pays $1 + i_{k,t}$ at $t + 1$. The consol in currency k trades at price $P_{k,t}^{\delta}$ 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

[5] The effects of monetary policy on risk pricing would be evident with trade only in one-period bonds. But we allow for trade in consols as well, anticipating that we will jointly seek to confront the model against data on exchange rates and yield curves.

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 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

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

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).[6] Financial wealth evolves according to

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

Firms A unit mass of monopolistically competitive firms in j hire labor and produce differentiated varieties which are then sold globally. Firm $s \in [0,1]$ faces Rotemberg (1982) adjustment costs when setting prices in each destination market,[7]

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

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

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

where χ_p controls the magnitude of price adjustment costs, $P_{kj,t}(s)$ is the firm's price in destination k, and $P_{kj,t}$ and $y_{kj,t}$ are aggregates that the firm takes as given. We assume local currency pricing (LCP) so that exchange rate pass-through is incomplete, as in the data, and we leave the analysis of alternatives (such as dollar pricing) for future work. The firm's local currency profits from global sales are thus

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

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

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

where

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

Finally, the firm uses a linear technology, so that

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

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

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

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

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

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

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

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. It can also be interpreted as capturing balance sheet constraints on arbitrageurs that depend on the variance of investment payoffs, and where the tightness of the constraint is rising in $\gamma/W_{a,t}$.

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

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

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

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

where $\nu_{ij,t}$ follows an AR(1) process with persistence ρ_{ij}. The U.S. monetary shock,

$\nu_{i1,t}$, 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

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

where we use that all firms in j will, in equilibrium, choose symmetric prices in each destination market k.

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 $p_{jk,t-1} \equiv P_{jk,t-1}/P_{j,t-1}$, 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 full equilibrium conditions of the model.

We build on Itskhoki and Mukhin (2021), Borovicka, Hansen, and Sargent (2023), and Kekre and Lenel (2024) 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 in currency j

$$ex_{j,t+1} \equiv \frac{E_{j,t}}{E_{j,t+1}}(1+i_{j,t}) - (1+i_{1,t})$$

and consol in currency j

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

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

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

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 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

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

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^{δ}_j and $E_t \hat{ex}^{\delta}_{j,t+1}$ for all j. In this way, currency and bond risk premia are time-varying even using a first-order approximation.

In equilibrium, the covariances of excess returns evaluated around steady-state need to be consistent with the first-order dynamics of the model. 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.[8] The implicit assumption we are making 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 covariances along the cross-section of currency and bond excess returns.

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 ($\chi_p \to \infty$). 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 $\{b^{\delta}_{jk}\}$, 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 $\{ex_j, ex^{\delta}_j\}$, one-period bond positions $\{b_{jk}\}$, and effective risk aversion Γ/w_a are small. These assumptions simplify the

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

mathematics without affecting the key economics.

We obtain:

Proposition 1. *Assume $\chi_p \to \infty$, zero $\{b^{\delta}_{jk}\}$ and ρ_i, and small $\{ex_j, ex^{\delta}_j, b_{jk}\}$ and Γ/w_a. Then given a U.S. monetary shock $d\nu_{i1,t}$,*

$$\hat{E}_{j,t} = \hat{q}_{j,t} = \Bigg(\left[\beta^{-1} - \rho_{i1}\right]^{-1} \left[1 + (\beta^{-1} - 1)(1 - \rho_{i1})^{-1}\psi\right]$$
$$+ \left[\beta^{-1} - \rho_{i1}\right]^{-1} \left[1 + (\beta^{-1} - 1)(1 - \rho_{i1})^{-1}\psi\right] \beta \frac{1 - \xi}{\xi} \frac{\sum_{k=2}^{n} q_k^{-1} b_{ak}}{w_a} ex_j$$
$$- \left[\beta^{-1} - \rho_{i1}\right]^{-2} \left[\beta^{-1} - 1\right]^{-1} \left[1 - \psi\right] \frac{\Gamma}{w_a} Cov\left(ex_{j,+1}, \sum_{k=2}^{n} \eta_{1k} ex_{k,+1}\right) \Bigg) d\nu_{i1,t}. \quad (3)$$

The first line of (3) 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_{i1}, and becomes arbitrarily large as the discount factor β and persistence ρ_{i1} 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 (3) 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 will reduce their wealth and thus raise their effective risk aversion. As a result, "risky" currencies ($ex_j > 0$) will 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.

The third line of (3) 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 $d\nu_{i1,t}$,*

$$\begin{aligned}
\hat{P}^{\delta}_{j,t} = \Bigg(& -\left[1-\delta\beta\rho_{i1}\right]^{-1} 1\{j=1\} \\
& -\left[\beta^{-1}-\rho_{i1}\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_{i1})^{-1}\psi\right]\frac{1-\xi}{1-\delta(1-\xi)}\frac{\sum_{k=2}^{n} q_k^{-1} b_{ak}}{w_a}\left(ex^{\delta}_j - ex_j\right) \\
& +\left[\beta^{-1}-\rho_{i1}\right]^{-1}\left[\beta^{-1}-\delta\rho_{i1}\right]^{-1}\left[1-\psi\right]\frac{\Gamma}{w_a} Cov\left(ex^{\delta}_{j,+1}-ex_{j,+1}, \sum_{k=2}^{n}\eta_{1k} ex_{k,+1}\right)\Bigg) d\nu_{i1,t}.
\end{aligned} \tag{4}$$

The first line of (4) 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 according to the expectations hypothesis, with a bigger fall in the price the higher is δ (which controls duration), discount factor β, and persistence ρ_{i1}. With fully sticky prices, 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 (4) 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.[9] 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 (4) 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

[9] In a closed economy, the revaluation of arbitrageur wealth through duration is the focus of Kekre et al. (2025).

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 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, (2) and the analog for ex_j^δ, as

$$\frac{1}{w_a}(\boldsymbol{q}^{-1}\boldsymbol{b}_a) = \frac{1}{\Gamma}\boldsymbol{C}^{-1}\boldsymbol{ex}, \tag{5}$$

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.[10] 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 to 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 15 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), Chile (CLP), Colombia (COP), Hungary (HUF), India (INR), Indonesia (IDR), Israel (ILS), 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

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

	CHF, EUR, JPY	Other non-U.S. G10	EM	Sum
$\Gamma = 1$	-1.9	-2.4	8.8	4.5
$\Gamma = 1$, G10 avg as base	-4.7	-1.4	14.2	8.1
$\Gamma = 1$, G10 IBOR yields	-2.9	-0.4	5.7	2.4
$\Gamma = 1$, no 10-year bonds	-0.6	-2.9	9.2	5.7
$\Gamma = 1$, 2008-2024	-3.6	-2.2	9.5	3.7
$\Gamma = 1$, 0.5*pp* higher dollar interest rate	-2.8	-2.7	6.4	0.9

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.

Bloomberg from April 1991 through February 2024.[11,12] We use government bond yields as our baseline measure of yields, but explore the consequences of using IBOR-based yields for the G10 economies as well. Since Bloomberg reports government par yields, we bootstrap zero coupon yields using this data.[13]

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,[14] 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.

[11] We choose this sample period because a consistent set of Bloomberg tickers for government bond yields 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.

[12] 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 yields and exchange rates used in the analysis.

[13] 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.

[14] 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.

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.[15]

Table 1 reports the implied mean-variance efficient portfolio shares. The portfolio shares in each currency include three-month and ten-year bonds. The holding period is assumed to be one quarter, so that the three-month bonds are "one-period" instruments.[16]

The first row reports that, with $\Gamma = 1$, a mean-variance efficient investor over the sample period has been on average short G10 currencies, long emerging market currencies, and net long all non-dollar currencies. Quantitatively, the investor would borrow 3.5 times their wealth in dollars; borrow 1.9 times their wealth in Swiss franc, euros, and yen; borrow an additional 2.4 times their wealth in other G10 currencies; and invest 8.8 times their wealth in emerging markets. In total, the investor would hold a leveraged position of 4.5 times their wealth in non-dollar currencies.[17] Loosely, 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 of the representative arbitrageur 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

[15] We use the subsample beginning in July 1998 to estimate the covariance matrix. While we still compute pairwise covariances to deal with the fact that we do not have a balanced panel, starting in July 1998 means that we have a balanced panel for the G10 economies at least, helping to ensure the resulting covariance matrix is positive semi-definite.

[16] We make this assumption because 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.

[17] While it is not reported in the table, we also note that the investor would hold a position of 1.2 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.

portfolio is invariant to the base currency of the investor.[18] 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 $\Gamma = 1$. In an extension of our model with arbitrageurs caring about payoffs in each possible base currency, these considerations would shape the risk-sharing among arbitrageurs (as in Black (1990)). Our assumption that the representative arbitrageur cares about payoffs in dollars is effectively an assumption that the risk aversion of dollar-based arbitrageurs is much 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 reports the sensitivity of the mean-variance efficient portfolio to alternative ways of estimating $\boldsymbol{ex}$ and $\boldsymbol{C}$. The efficient portfolio has a similar structure 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. The efficient portfolio is also similar using a factor model to construct $\boldsymbol{C}$. In particular, dropping ten-year bonds from the menu of assets (to ease in relating to the prior literature), we use the first two principal components of excess currency returns or the (closely related) dollar and carry factors of Lustig, Roussanov, and Verdelhan (2011)

[18]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.

as factors. One virtue of the factor models is that they can be used to decompose the efficient portfolio into the component arising from priced risk factors, and the component arising from alpha. Both components tend to imply an efficient portfolio that is short the dollar, with their relative role depending on the factors used.

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 5.5.

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 markets. While specific investors along the intermediation chain may be largely hedged, 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 risk premia, 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 in the nearest four quarterly Eurodollar contracts around U.S. monetary policy an-

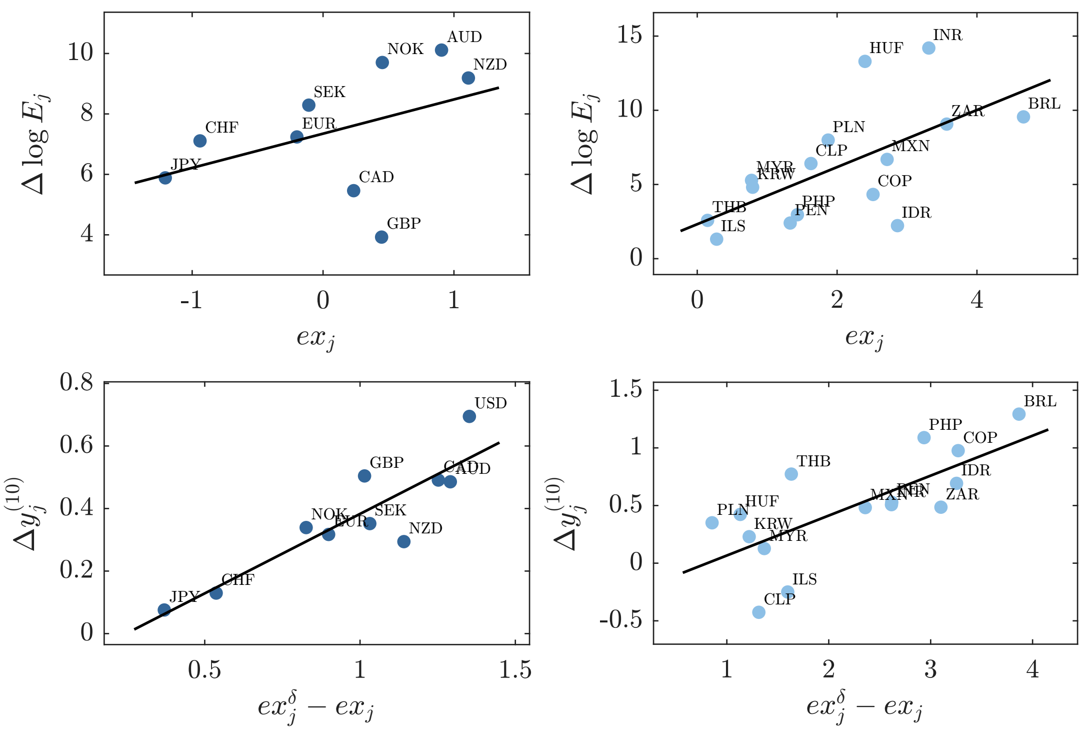

Figure 1: asset price responses to U.S. monetary tightening in data

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 $1pp$. Expected excess returns along x-axis are annualized.

nouncements.[19] 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.

[19]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.

Figure 1 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 expected average local currency bond carry returns, $ex_j^\delta - ex_j$.[20] 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 the G10 and emerging market economies, are novel.

Appendix C demonstrates that it is difficult to explain the response of these asset prices to a U.S. monetary tightening 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 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*.[21,22]

A similar challenge faces portfolio balance effects arising from other potential capital flows on impact of a U.S. tightening. For instance, one could generalize our model to assume that 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

[20] 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).

[21] 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.

[22] 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 next footnote, we also study quantitative tightening in appendix C).

(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.

Taken together, we conclude that the international asset price responses to a U.S. monetary tightening suggest a dominant role for a higher global price of risk.[23]

5 Quantification

We now discipline the model to rationalize the high-frequency effects of U.S. monetary policy on the cross-section of international asset prices. We show that arbitrageurs' dollar funding is at the core of what makes U.S. monetary policy special. The risk premium spillovers of U.S. monetary policy imply a stagflationary trade-off for foreign central banks, which can in part be managed through FX intervention. In a future of higher dollar interest rates, these spillovers of U.S. monetary policy will be 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.[24] We set the elasticity of substitution across product varieties ϵ to 11 and the price adjustment cost χ_p to $41,600$; together, these imply a Calvo (1983)-equivalent frequency of price adjustment around five quarters. 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 assume that $\delta = 1 - (52 \times 10)^{-1}$ so that the duration of consols is approximately 10 years. Finally, we set the conditional covariance of excess returns

[23] This result is sharpened by contrast with the asset price responses to U.S. quantitative tightening. Appendix C reports the global responses of exchange rates and bond yields using Swanson (2021)'s measure of news about large-scale asset purchases. It demonstrates that these responses *are* consistent with spillovers through portfolio balance. This both validates the substitution elasticities across assets implied by mean-variance optimization, and sharpens the finding that interest rate policy spills over through a distinct mechanism.

[24] 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	41,600
ϕ_π	interest rate coeff. on CPI inflation	1.5
ρ_i	interest rate inertia	$0.4^{1/52}$
δ	consol decay	$1-(52 \times 10)^{-1}$
$Cov(ex_{j,+1}, ex_{k,+1})$	conditional cov. of carry returns	see text

Table 2: baseline parameters set without solving model

Notes: model is solved at a weekly frequency.

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.[25,26] We set the populations ζ_j to target average consumption relative to the U.S. as reported in the IMF national accounts. We set the trade shares η_{jk} to target expenditure shares using the IMF's Direction of Trade Statistics (DOTS) and national accounts data. Finally, we set the discount factors β_j and household positions in foreign bonds and consols $\{\bar{b}_{jk}, \bar{b}^{\delta}_{jk}\}$ 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.[27]

There are (infinitely) many ways to set households' exogenous positions in foreign

[25] 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.

[26] We solve the model for all G10 economies used in our empirical analysis, the nine largest emerging markets from the 15 used in our empirical analysis, and a "rest of world", bringing it to 20 countries. Appendix D describes how we set parameters for the "rest of world".

[27] As described in footnote 16, 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.

	Description	Moments	Targets
ζ_j	populations	$\zeta_j q_j^{-1} c_j / c_1$	relative consumption
β_j	discount factors	i_1, ex_j	dollar rate, excess returns
$\bar{b}^\delta_{1j}$	consol demand	ex^δ_j	excess returns
η_{jk}	trade shares	$P_{jk} c_{jk} / (P_j c_j)$	expend. shares by origin
Γ	arbitrageur risk aversion	resp. to U.S. monetary surprise (see text)	
ξ	arbitrageur exit rate	resp. to U.S. monetary surprise (see text)	
$\rho_{\nu i1}$	persist. U.S. MP shock	resp. to U.S. monetary surprise (see text)	
ω_a	arbitrageur 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 $\{\bar{b}_{jk} = 0, \bar{b}^\delta_{jk} = 0\}$ for $j > 1$.

bonds and consols $\{\bar{b}_{jk}, \bar{b}^\delta_{jk}\}$ 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 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 (2024)).

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 $\rho_{\nu i1}$, 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 observed average excess returns and conditional covariances of excess returns in the data, arbitrageurs' portfolio is already pinned down up to leverage, as argued in section 4.1 — their risk aversion Γ is the only remaining degree of freedom we have to pin down their portfolio, as it governs their leverage. Consistent with Propositions 1 and 2, arbitrageurs' leverage, ξ, and $\rho_{\nu i1}$ shape the asset price responses to U.S. monetary surprises, and can thus be identified by them.

Formally, we use the generalized method of moments (GMM) to choose these parameters to minimize the squared deviation of exchange rate, two-year yield, and 10-year yield responses between data and model,[28] simulating the U.S. monetary surprise as a positive innovation to $\nu_{i1,t}$ which raises the two-year dollar yield by $1pp$ as in the data. We use identical weights on all moments, and 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 three years.[29] We discipline ω_a to match a $1.05pp$ decline in the 10-year dollar yield on impact of the shock.[30] We return to discuss the implications of the calibrated ω_a when we study FX interventions in response to U.S. monetary policy.

Figure 2 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 29 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

[28] 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.

[29] 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 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%.

[30] 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.

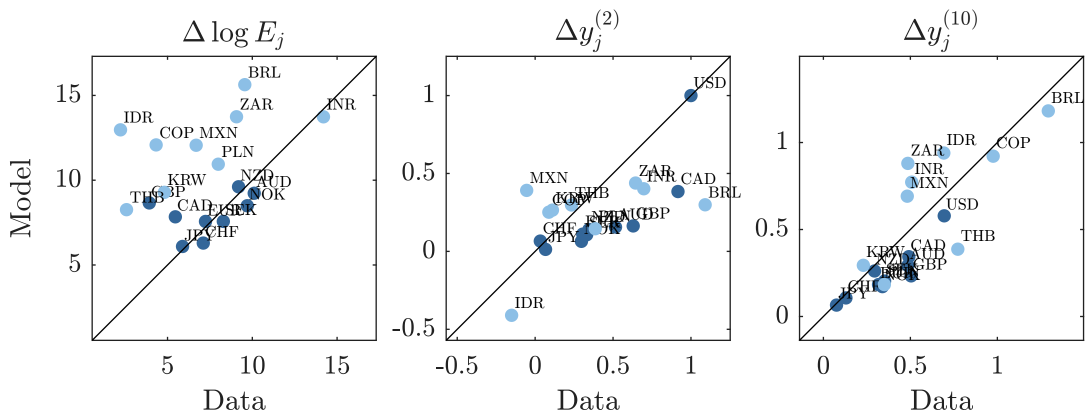

Figure 2: asset price responses to U.S. monetary tightening in model vs. data

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

emerging markets may reflect two forces. First, many of these economies likely engage in exchange rate intervention to prevent their exchange rate from depreciating without relying on an increase in policy rates, consistent with the arguments in Albagli et al. (2019) and Albagli, Ceballos, Claro, and Romero (2024). We study such interventions later in this paper. Second, it may be that we are especially mismeasuring the true currency and bond term premia for these economies, which control their exposure to a change in risk pricing.

Figure 3 depicts how Γ, governing arbitrageurs' leverage and thus revaluation of wealth upon a dollar appreciation, is identified. The estimated Γ implies that arbitrageurs' leverage in non-dollar bonds

$$\frac{\sum_{k=2}^{n} q_k^{-1}(b_{ak} + b_{ak}^{\delta})}{w_a} \approx 25.$$

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. Appendix D demonstrates how ξ, governing arbitrageurs' persistence of wealth, is similarly identified by these cross-sectional patterns. We estimate $\xi = 0.026$, so that the expected lifetime of an individual arbitrageur is approximately three quarters.

The model requires a very high degree of arbitrageur leverage to rationalize the

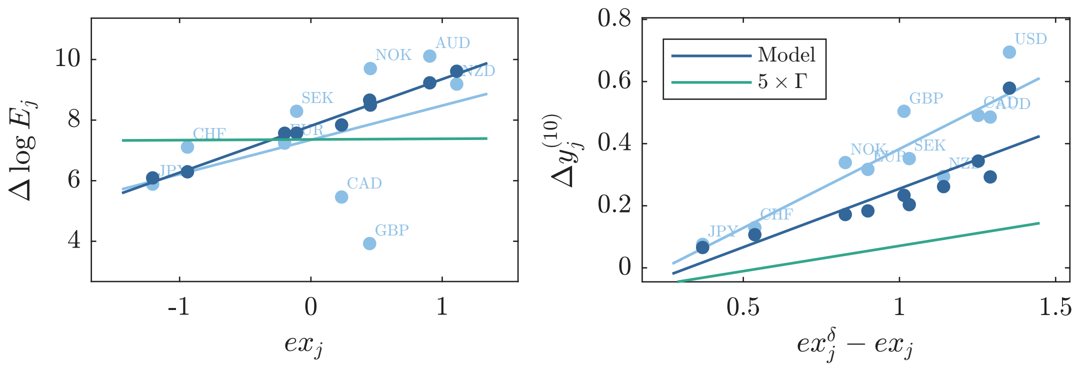

Figure 3: identification of arbitrageur leverage

Notes: both model and data responses reported for surprise that raises two-year dollar yield by $1pp$. Expected excess returns along x-axis are annualized.

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. The first panel of Figure 4 depicts the impulse response of arbitrageur wealth to the same shock depicted in Figure 2. While arbitrageur log-wealth falls by a dramatic 800 pp, it rapidly rebounds to be only 15 pp below its steady-state value by one year after the shock.[31] The response of expected excess currency and bond excess returns in the second and third panels inherit these dynamics.

We can consider an alternative specification of arbitrageurs' wealth process (1) in which they retain only the unexpected component of excess returns:

$$W_{a,t+1} = \xi P_{1,t+1}\omega_a + (1-\xi)\left((1+i_{1,t})W_{a,t} + \Pi_{a,t+1} - E_t\Pi_{a,t+1}\right).$$

This could be microfounded by arbitrageurs having to pay the expected component of their excess returns to households each period. Given this process, we re-calibrate $\{\Gamma, \xi, \rho_{\nu i1}, \omega_a\}$ to match the same moments as in the baseline model. Appendix D demonstrates that the asset price responses, as well as effects on the real allocation (studied later in this section), are comparable to the baseline model. However, the calibrated Γ implies arbitrageurs' leverage in non-dollar bonds which is now only

$$\frac{\sum_{k=2}^{n} q_k^{-1}(b_{ak} + b_{ak}^\delta)}{w_a} \approx 4.5,$$

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

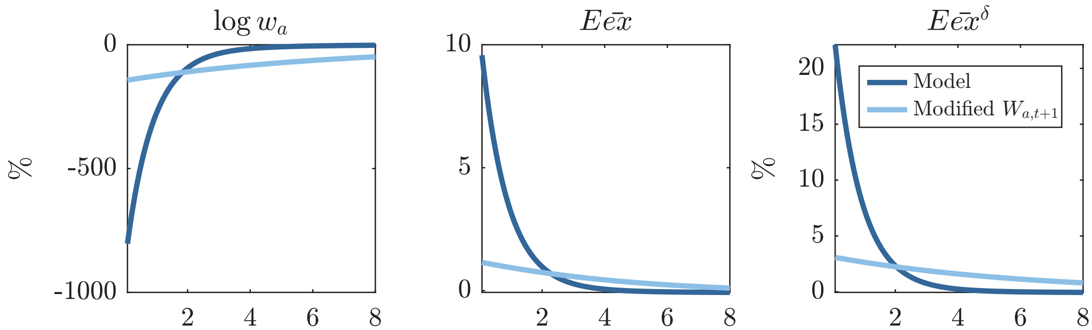

Figure 4: response of arbitrageur wealth and risk premia

Notes: responses reported for surprise that raises two-year dollar yield by $1pp$. Environment with modified wealth process assumes that arbitrageurs only accumulate wealth according to unexpected component of their profits. $E_t\bar{ex}_{t+1} \equiv \frac{1}{n}\sum_{j=2}^{n} E_t ex_{j,t+1}$ denotes average (annualized) expected excess return on non-dollar currencies and $E_t\bar{ex}^{\delta}_{t+1} \equiv \frac{1}{n}\sum_{j=1}^{n} E_t ex^{\delta}_{j,t+1}$ denotes average (annualized) expected excess return on all consols. x-axis denotes quarters since the shock.

far smaller than in the baseline model. The reason is that arbitrageurs' wealth does not rebound so aggressively following the shock, as depicted by the light blue line in Figure 4. Hence, this calibration still matches the observed exchange rate and yield responses on impact because risk premia adjust more persistently.

5.2 The specialness of U.S. monetary policy

We can use our model to understand what makes U.S. monetary policy special.

We first demonstrate the specialness of U.S. monetary policy in the data. Figure 5 compares the high-frequency responses of exchange rates to Federal Reserve surprises versus European Central Bank (ECB) surprises. We focus on Euro area surprises because it is an economy of comparable size to the U.S. We measure ECB monetary surprises using the "forward guidance" surprise measure from Altavilla, Brugnolini, Gurkaynak, Motto, and Ragusa (2019).[32] In both panels, we consider a surprise that

[32]That paper provides several measures of Euro area monetary surprises extracted from the intraday response of asset prices around ECB monetary announcements. One, called the "timing factor", concerns the near-term path of monetary policy (explaining most variation in OIS rates around the monetary announcement with maturity less than or equal to one year). The second, a "forward guidance factor", concerns the longer-term path of monetary policy (explaining most variation in OIS rates around the monetary announcement with maturity between two and five years). We focus for concreteness on the second measure, constructed from the event window that encapsulates both the press release and press conference, but similar results hold for the first measure. The surprises begin in January 1999 and we study all such surprises through December 2019, for

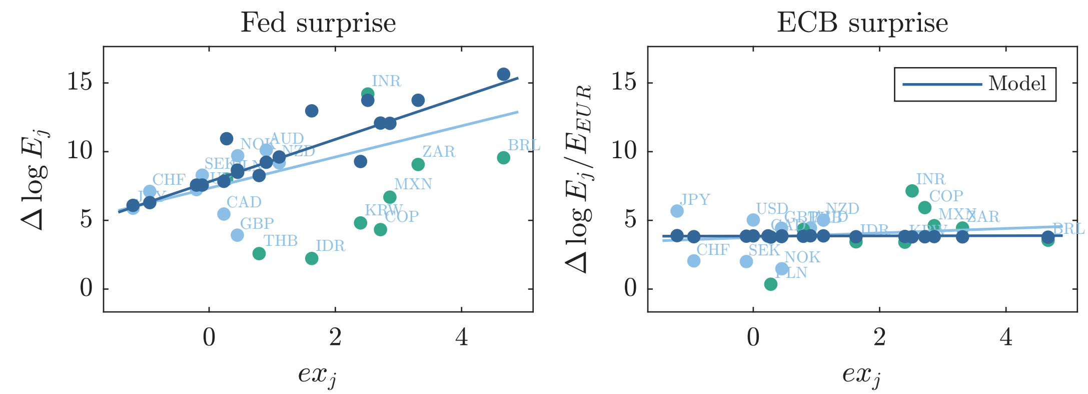

Figure 5: asset price responses to U.S. vs. Euro area monetary tightening

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. Expected excess returns along x-axis are annualized.

raises the two-year domestic yield by $1pp$. 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. 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 5 then demonstrates that our model generates precisely such an asymmetry between 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. We see that 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, unlike in the case of a U.S. monetary surprise.[33] This follows from

comparability with the sample of Federal Reserve surprises studied in the last section.

[33]The model still does generate heterogeneity in exchange rate responses to a Euro area surprise, largely explained by trade relationships with the Euro area. For instance, Switzerland and the Scandinavian currencies depreciate by less, and the U.S dollar depreciates by more, in response to

the fact that arbitrageurs fund only a small fraction of their portfolio in euros, as described in the discussion of 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. The dollar's dominant role as a funding currency is thus essential for the disproportionate effects of U.S. monetary policy on financial markets.

5.3 The importance of risk premium spillovers

We now quantify the importance of risk premium spillovers of U.S. monetary policy, both relative to other channels of international transmission, and for the real economy.

The model first clarifies that the change in risk pricing dominates trade in generating cross-country variation in response to a U.S. tightening. Figure 6 compares exchange rate and yield curve responses in the calibrated model versus a counterfactual environment in which arbitrageurs are risk neutral, 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 risk pricing are thus shut down in the latter environment, and the only remaining source of heterogeneous spillovers across countries are heterogeneous trade linkages (as disciplined by bilateral expenditure shares) and heterogeneous country sizes. These simply cannot rationalize the cross-section of exchange rate and yield curve responses to a U.S. tightening that we see in the data.[34]

The model also clarifies that these heterogeneous spillovers through risk pricing in turn matter for real allocations. Figure 7 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. A U.S. monetary tightening is accompanied by an endogenous policy tightening in the rest of the world. Relative to the G10 economies, emerging markets see a bigger increase in policy rate and thus short-dated yields, a larger decline in consumption, and a larger increase in inflation. In other words, emerging markets experience a relative "sudden stop" in response to a U.S. tightening.[35] These simulations thus

a Euro area tightening, as in the data.

[34]In particular, since the U.S. trades more with Canada and Mexico than any other economy, a U.S. tightening induces a larger increase in interest rates in Canada and Mexico, and a smaller movement in the Canadian dollar and Mexican peso, relative to all other economies. There is little difference in the exchange rate and yield curve responses among the remaining economies.

[35]Hence, the model at least qualitatively generates the heterogeneous dynamic responses to U.S.

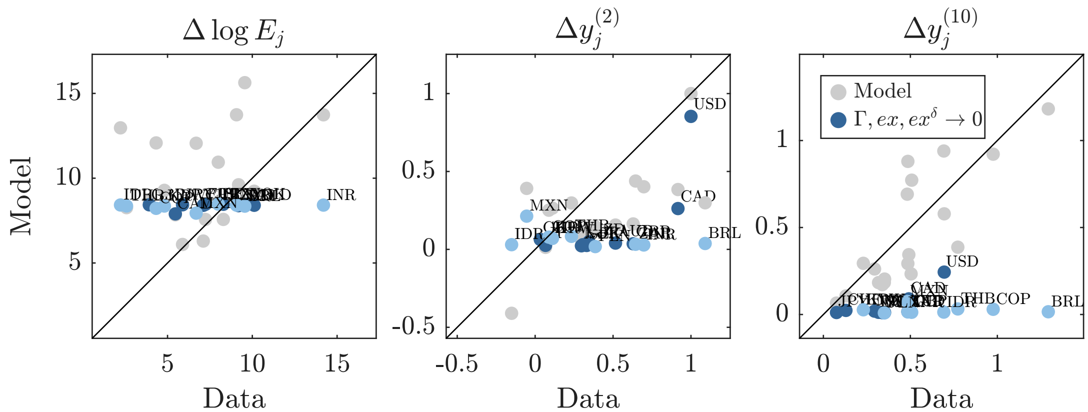

Figure 6: comparing transmission mechanisms of U.S. tightening

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

demonstrate that the effects of U.S. monetary policy on global risk pricing have first-order implications for real transmission in the rest of the world.

5.4 Policy responses to a U.S. tightening

We can also use the model to ask how the response to a U.S. tightening can shape domestic outcomes. Focusing on Brazil as a representative emerging market, we consider two alternative policy responses to the U.S. tightening than is assumed in our baseline model: an alternative response of conventional monetary policy, and exchange rate intervention.[36]

Figure 8 adds a term targeting the (real) exchange rate to Brazil's monetary policy rule,

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

monetary surprises between emerging market and advanced economies documented in Mendoza-Fernandez and Pelin (2025).

[36] In both cases, we are keeping the covariance matrix of excess returns $\boldsymbol{C}$ unchanged. Hence, this should be thought of as a "one-off" change in the reaction to a U.S. tightening that does not change anticipated future comovements. Jeanne and Rose (2002), Itskhoki and Mukhin (2023), Caballero, Caravello, and Simsek (2025), and Caballero and Simsek (2026) emphasize how the policy rule affects these second moments and thereby risk pricing.

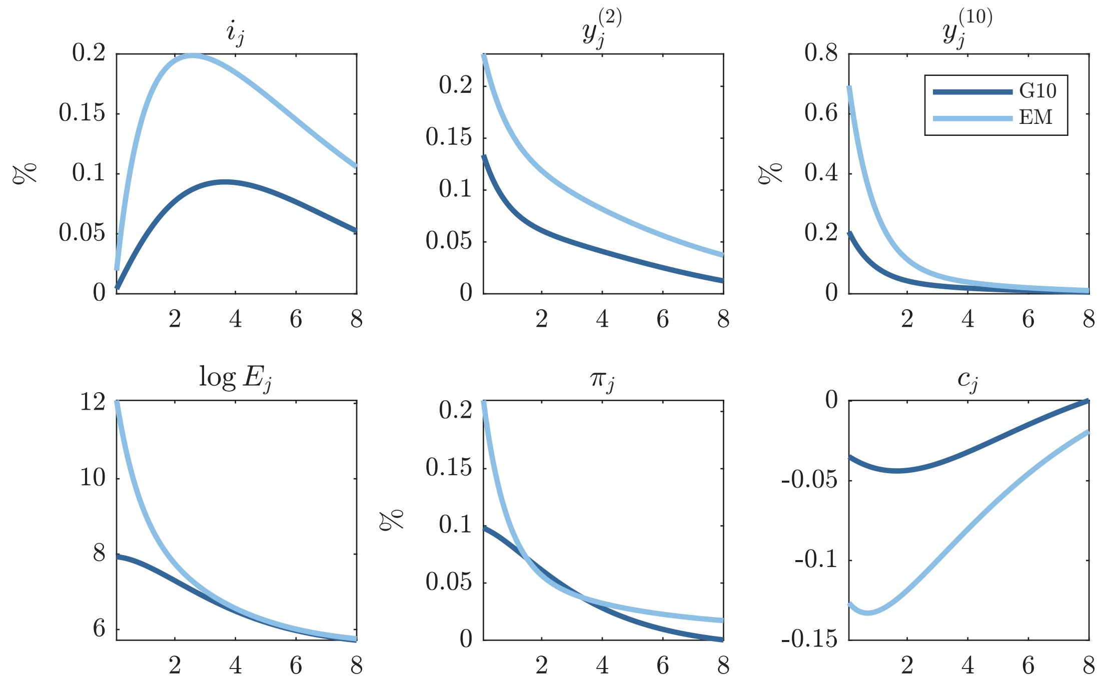

Figure 7: impulse responses to U.S. monetary tightening

Notes: responses reported for same $\nu_{i1,t}$ innovation depicted in Figure 2. x-axis denotes quarters since the shock.

With a positive value $\phi_{q,BRL}$, Brazil raises its policy rate by more in response to a U.S. tightening. As a result, it experiences a smaller depreciation and smaller inflation (in fact, deflation), but at the cost of a more severe decline in consumption. Quantitatively, the effect on the nominal exchange rate is minimal, while that on inflation is more meaningful, owing to the decline in domestic demand.

Perhaps more interesting is the possibility that Brazil can engage in exchange rate intervention which, as Figure 9 demonstrates, more directly addresses the risk premium shock at its source. We model such intervention as the Brazilian government selling one-period dollar bonds and buying one-period Brazilian bonds. Since Brazil is Ricardian with respect to the 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. We thus consider a rule

$$\bar{b}_{BRL1,t} = -(\phi_{fxi,BRL}/52)\left(\log q_{BRL,t} - \log q_{BRL}\right).$$

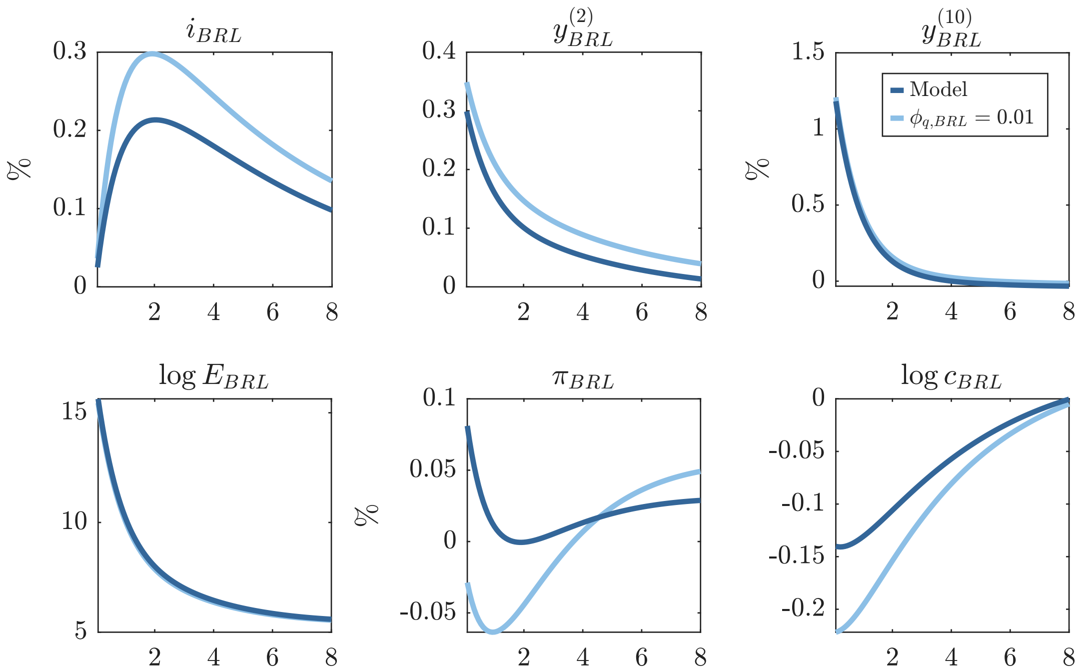

Figure 8: effects of alternative monetary policy response to U.S. monetary tightening

Notes: responses reported for same $\nu_{i1,t}$ innovation depicted in Figure 2. x-axis denotes quarters since the shock.

With a positive value $\phi_{fxi,BRL}$, the government sells dollar bonds on behalf of local households, who then 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, resulting in both dampened inflation and a dampened decline in consumption. Moreover, because the conditional covariance of the Brazilian currency excess return and Brazilian local bond carry return are positive — that is, the Brazilian real tends to depreciate when Brazilian long-term yields rise[37] — the lower amount of short-term Brazilian bonds on arbitrageurs' balance sheet also reduces the term premium on long-term Brazilian bonds, reflected in a decline in long-term yields.

We choose the parameter $\phi_{fxi,BRL}$ in this example so that Brazil's position in dollar bonds falls by 20% of its annual GDP.[38] A key parameter governing the price (and

[37] This finding for emerging markets like Brazil echoes points made in Kekre and Lenel (2024) and De Leo, Keller, Simoncelli, Villamizar-Villegas, and Williams (2025).

[38] We choose this amount of intervention because, according to Milesi-Ferretti (2026), Brazil's reserve balances excluding gold have averaged 15-25% of its annual GDP in recent years. Thus, it

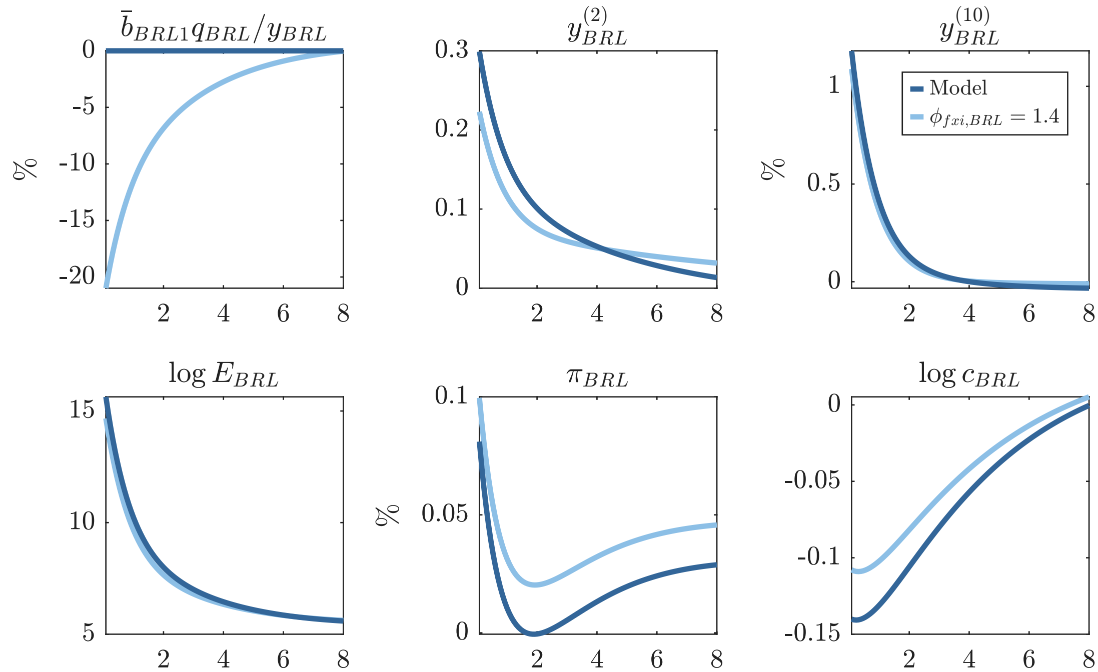

Figure 9: effects of FX intervention in response to U.S. monetary tightening

Notes: responses reported for same $\nu_{i1,t}$ innovation depicted in Figure 2. x-axis denotes quarters since the shock.

thus macroeconomic) impact of this intervention is arbitrageurs' initial endowment, ω_a. At a higher value of ω_a and thus steady-state arbitrageur wealth, the same dollar flow absorbed by arbitrageurs will change required excess returns by less. We disciplined ω_a using the price impact of QE1. Hence, Figure 9 implies that, disciplined by the price impact of QE, FX intervention by 20% of Brazil's GDP can dampen the Brazilian depreciation by roughly $1pp$ in response to a U.S. tightening.

Relative to previous models studying small open economies exposed to exogenous risk premium shocks,[39] a key novelty in the present environment is that the sudden stop affecting Brazil is endogenized in a multi-country setting. In particular, the depreciation of emerging market currencies like Brazil plays a key role in driving the increase in risk pricing to begin with, owing to the leverage of arbitrageurs in these currencies. While a formal normative analysis is outside the scope of the present

roughly corresponds to Brazil selling all of its reserves, assuming they are in short-term dollar bonds.

[39]See, for instance, Schmitt-Grohe and Uribe (2012), Farhi and Werning (2012, 2014), Bianchi and Lorenzoni (2022), and Corsetti, Dedola, and Leduc (2023).

paper, we conjecture that the gains from policy coordination will be different in the present environment than previous models of small open economies facing exogenous risk premium shocks. This is because countries will not internalize the cumulative effects of their policy actions on the broad dollar exchange rate and, in turn, the risk-bearing capacity of global arbitrageurs.[40]

5.5 A future with higher dollar interest rates

We finally use the model to study a potential future with 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.[41] 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 25 in the baseline model, to 5 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 2). 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

[40] This is distinct from the role for policy coordination recently emphasized in Fontanier (2025). In that paper, arbitrageurs are not subject to currency risk, but emerging market firms who borrow in dollar-denominated debt are. From a global perspective, that paper argues that countries excessively tighten their policy rate in response to an external risk premium shock, because they do not internalize that the associated capital inflows and increase in dollar borrowing costs affects other countries. By contrast, we anticipate that in our model, countries may *insufficiently* tighten their policy rate in response to an external risk premium shock, because they do not internalize that the resulting dollar depreciation would mitigate the risk premium shock to begin with.

[41] 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 the mechanism by which FX interventions are non-neutral in the prior subsection.

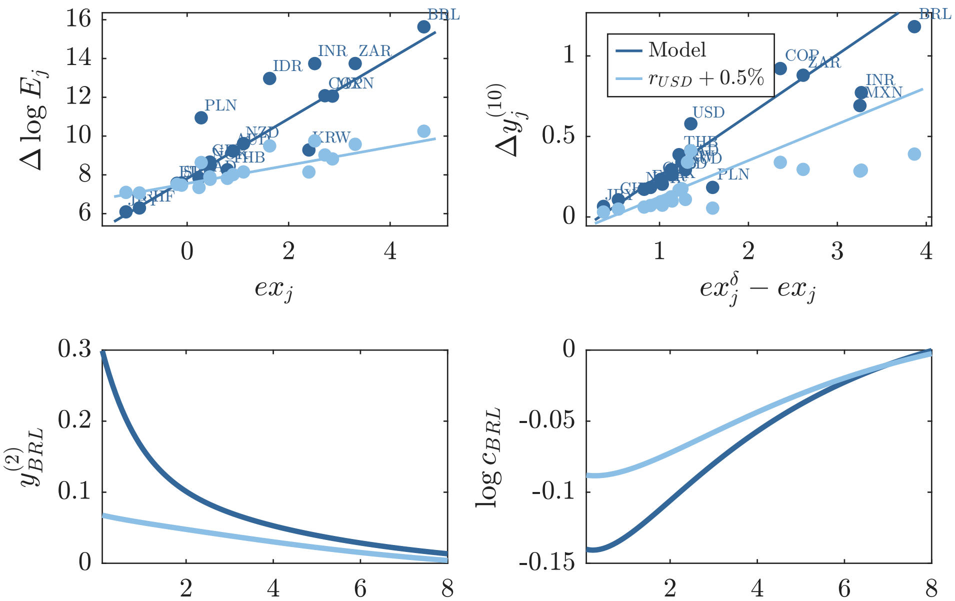

Figure 10: responses to U.S. monetary tightening with 0.5*pp* higher U.S. interest rate

Notes: responses reported for same $\nu_{i1,t}$ innovation in Figure 2. Expected excess returns along x-axis in top panels are annualized. x-axis in bottom panels denote quarters since the shock.

bond yields of currencies with high excess bond carry returns.

These modified effects on U.S. monetary policy transmission to asset prices are in turn reflected in the real spillovers from U.S. tightening. The second row of Figure 10 compares the impulse responses to a U.S. tightening on outcomes in Brazil in the baseline model, versus this counterfactual environment. Brazil experiences a smaller increase in near-dated yields and a smaller decline in consumption, consistent with a dampened effective risk premium shock. In this way, the model implies that a future with higher dollar interest rates will feature dampened U.S. monetary spillovers through risk pricing, particularly among emerging markets.

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 quantify arbitrageurs' leverage to rationalize this evidence, and we assess the implications for global economic activity. In a potential future with higher dollar interest rates, the model implies that arbitrageurs will be less currency-mismatched and the spillovers of U.S. monetary policy through risk pricing will be 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.[42] 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.[43] Second, our model can be used to study normative questions. We have conjectured that 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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[42] 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 (2024).

[43] 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

$$\begin{aligned} w_{j,t} &\equiv W_{j,t}/P_{j,t}, \\ \pi_{kj,t} &\equiv \log P_{kj,t}/P_{kj,t-1}, \\ b_{jk,t} &\equiv B_{jk,t}/P_{k,t}, \\ b^{\delta}_{jk,t} &\equiv B^{\delta}_{jk,t}/P_{k,t}. \end{aligned}$$

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

$$1 = E_t \beta_j \left(\frac{c_{j,t+1}}{c_{j,t}} \right)^{-1/\psi} \exp(-\pi_{j,t+1})(1 + i_{j,t}), \ \ \forall j, \tag{6}$$

$$b_{jk,t} = \bar{b}_{jk}, \ \ \forall j, k \neq j, \tag{7}$$

$$b^{\delta}_{jk,t} = \bar{b}^{\delta}_{jk}, \ \ \forall j, k, \tag{8}$$

$$\bar{\varphi} \ell^{1/\varphi}_{j,t} c^{1/\psi}_{j,t} = w_{j,t}, \ \ \forall j, \tag{9}$$

$$c_{jk,t} = \eta_{jk} \left(p_{jk,t} \right)^{-\sigma} c_{j,t}, \ \ \forall j, k. \tag{10}$$

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:

$$c_{j,t} + \sum_{k=1}^{n} q_{j,t} q^{-1}_{k,t} \left(b_{jk,t} + b^{\delta}_{jk,t} \right) = \zeta^{-1}_j q_{j,t} \sum_{k=1}^{n} \zeta_k q^{-1}_{k,t} p_{kj,t} c_{kj,t} + a_{j,t}, \ \ \forall j > 1, \tag{11}$$

$$a_{j,t+1} = \sum_{k=1}^{n} \exp(\Delta \log q_{j,t+1} - \Delta \log q_{k,t+1} - \pi_{k,t+1}) \times$$

$$\left((1+i_{k,t})q_{j,t}q_{k,t}^{-1}b_{jk,t}+\left(\frac{1+\delta P_{k,t+1}^{\delta}}{P_{k,t}^{\delta}}\right)q_{j,t}q_{k,t}^{-1}b_{jk,t}^{\delta}\right),\quad \forall j>1. \tag{12}$$

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

$$\begin{aligned}(\exp(\pi_{kj,t})-1)\exp(\pi_{kj,t}) &= \frac{\varepsilon-1}{\chi_p}\left[\frac{\varepsilon}{\varepsilon-1}\frac{w_{j,t}}{q_{j,t}q_{k,t}^{-1}p_{kj,t}}-1\right] \\ &+ E_t\beta\left(\frac{c_{j,t+1}}{c_{j,t}}\right)^{-1/\psi}\exp(\Delta\log q_{j,t+1}-\Delta\log q_{k,t+1})\frac{p_{kj,t+1}c_{kj,t+1}}{p_{kj,t}c_{kj,t}}\times \\ &(\exp(\pi_{kj,t+1})-1)\exp(\pi_{kj,t+1}),\quad \forall j,k. \end{aligned}\tag{13}$$

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

$$ex_{j,t+1}=\frac{E_{j,t}}{E_{j,t+1}}(1+i_{j,t})-(1+i_{1,t}),\quad \forall j>1, \tag{14}$$

$$ex_{j,t+1}^{\delta}\equiv\frac{E_{j,t}}{E_{j,t+1}}\frac{1+\delta P_{j,t+1}^{\delta}}{P_{j,t}^{\delta}}-(1+i_{1,t}),\quad \forall j, \tag{15}$$

$$\begin{aligned}E_t ex_{j,t+1} &= \frac{\gamma}{w_{a,t}}\Bigg[\sum_{k=2}^{n}Cov_t\left(ex_{j,t+1},ex_{k,t+1}\right)q_{k,t}^{-1}b_{ak,t} \\ &+\sum_{k=1}^{n}Cov_t\left(ex_{j,t+1},ex_{k,t+1}^{\delta}\right)q_{k,t}^{-1}b_{ak,t}^{\delta}\Bigg],\quad \forall j>1,\end{aligned}\tag{16}$$

$$\begin{aligned}E_t ex_{j,t+1}^{\delta} &= \frac{\gamma}{w_{a,t}}\Bigg[\sum_{k=2}^{n}Cov_t\left(ex_{j,t+1}^{\delta},ex_{k,t+1}\right)q_{k,t}^{-1}b_{ak,t} \\ &+\sum_{k=1}^{n}Cov_t\left(ex_{j,t+1}^{\delta},ex_{k,t+1}^{\delta}\right)q_{k,t}^{-1}b_{ak,t}^{\delta}\Bigg],\quad \forall j,\end{aligned}\tag{17}$$

$$\sum_{j=1}^{n}q_{j,t}^{-1}\left[b_{aj,t}+b_{aj,t}^{\delta}\right]=w_{a,t}, \tag{18}$$

$$\begin{aligned}w_{a,t+1} &= \xi\omega_a+(1-\xi)\exp(-\pi_{1,t+1})\Bigg[(1+i_{1,t})w_{a,t} \\ &+\sum_{j=2}^{n}ex_{j,t+1}q_{j,t}^{-1}b_{aj,t}+\sum_{j=1}^{n}ex_{j,t+1}^{\delta}q_{j,t}^{-1}b_{aj,t}^{\delta}\Bigg].\end{aligned}\tag{19}$$

The definitions of CPIs and mapping between prices are:

$$1 = \left(\sum_{k=1}^{n} \eta_{jk} \left(p_{jk,t} \right)^{1-\sigma} \right)^{\frac{1}{1-\sigma}}, \ \ \forall j, \tag{20}$$

$$\pi_{kj,t} = \Delta \log p_{kj,t} + \pi_{k,t}, \ \ \forall j, k, \tag{21}$$

$$\Delta \log q_{j,t} = \Delta \log E_{j,t} + \pi_{1,t} - \pi_{j,t}, \ \ \forall j > 1. \tag{22}$$

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

$$\sum_{k=1}^{n} \zeta_k c_{kj,t} = \zeta_j \ell_{j,t}, \ \ \forall j, \tag{23}$$

$$\sum_{k=1}^{n} \zeta_k b_{kj,t} + b_{aj,t} = 0, \ \ \forall j, \tag{24}$$

$$\sum_{k=1}^{n} \zeta_k b^{\delta}_{kj,t} + b^{\delta}_{aj,t} = 0, \ \ \forall j. \tag{25}$$

The specification of policy is finally:

$$\log(1 + i_{j,t}) = \log(1 + i_j) + \rho_i \left(\log(1 + i_{j,t-1}) - \log(1 + i_j) \right) + (1 - \rho_i) \phi_\pi \pi_{j,t} + \nu_{ij,t}, \ \ \forall j. \tag{26}$$

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

$$c_{1,t} + \sum_{k=1}^{n} q_{k,t}^{-1} \left(b_{1k,t} + b_{ak,t} + b^{\delta}_{ak,t} + b^{\delta}_{1k,t} \right) = \sum_{k=1}^{n} \zeta_k q_{k,t}^{-1} p_{k1,t} c_{k1,t}$$
$$+ \sum_{k=1}^{n} q_{k,t}^{-1} \exp(-\pi_{k,t}) \left((1 + i_{k,t-1})(b_{1k,t-1} + b_{ak,t-1}) + \left(\frac{1 + \delta P^{\delta}_{k,t}}{P^{\delta}_{k,t-1}} \right) \left(b^{\delta}_{1k,t-1} + b^{\delta}_{ak,t-1} \right) \right)$$

is satisfied as well.

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

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

given the endogenous state variables $\{\{p_{jk,t-1}\}, \{q_{j,t-1}, E_{j,t-1}, a_{j,t}\}_{j>1}, w_{a,t}\}$, exogenous state variables $\{\nu_{ij,t}\}$ 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 (16)-(17) is standard, we focus on the latter. The approach and exposition closely follows that in Kekre and Lenel (2024), 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

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

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 $\log q_{k,t}$, $b_{ak,t}$, $b^\delta_{ak,t}$, $\log w_{a,t}$, $ex_{j,t+1}$, and $ex^\delta_{j,t+1}$ around the point of approximation is given by

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

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

By the property of lognormally distributed variables, we have that

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

where Σ is the covariance matrix of the innovations, scaled appropriately by σ^2. An analogous argument holds for $Cov_t\left(ex_{j,t+1}, ex^\delta_{k,t+1}\right)$ and $Cov_t\left(ex^\delta_{j,t+1}, ex^\delta_{k,t+1}\right)$. Thus, substituting the first-order approximations for all variables into (16) yields

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

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

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

Differentiating the above condition with respect to each state variable and evaluating at the point of approximation, the coefficients will solve

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

It is straightforward to show that a similar argument applies for consol pricing (17). 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,

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

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 (6)-(26) around an arbitrary steady-state, but already using that $\chi_p \to \infty$, which implies

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

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:

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

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:

$$c_j\hat{c}_{j,t} + \hat{b}_{jj,t} + \sum_{k=1}^{n} q_j q_k^{-1}(b_{jk} + b_{jk}^{\delta})\left(\hat{q}_{j,t} - \hat{q}_{k,t}\right)$$

$$= \zeta_j^{-1} q_j \sum_{k=1}^{n} \zeta_k q_k^{-1} p_{kj} c_{kj}\left(\hat{q}_{j,t} - \hat{q}_{k,t} + \hat{c}_{k,t}\right) + \hat{a}_{j,t}, \quad \forall j > 1,$$

$$\hat{a}_{j,t+1} = (1+i_j)\hat{b}_{jj,t} + \sum_{k=1}^{n}\Bigg[\left((1+i_k)q_j q_k^{-1} b_{jk} + \left(\frac{1+\delta P_k^{\delta}}{P_k^{\delta}}\right) q_j q_k^{-1} b_{jk}^{\delta}\right)\left(\hat{q}_{j,t+1} - \hat{q}_{k,t+1}\right)$$

$$+ (1+i_k) q_j q_k^{-1} b_{jk} \hat{i}_{k,t} + \left(\frac{1+\delta P_k^{\delta}}{P_k^{\delta}}\right) q_j q_k^{-1} b_{jk}^{\delta}\left(\frac{\delta P_k^{\delta}}{1+\delta P_k^{\delta}}\hat{P}_{k,t+1}^{\delta} - \hat{P}_{k,t}^{\delta}\right)\Bigg], \quad \forall j > 1.$$

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

$$\hat{ex}_{j,t+1} = (1+i_j)\left(-\Delta\hat{q}_{j,t+1} + \hat{i}_{j,t}\right) - (1+i_1)\hat{i}_{1,t}, \quad \forall j > 1,$$

$$\hat{ex}_{j,t+1}^{\delta} = \left(\frac{1+\delta P_j^{\delta}}{P_j^{\delta}}\right)\left(-\Delta\hat{q}_{j,t+1} + \frac{\delta P_j^{\delta}}{1+\delta P_j^{\delta}}\hat{P}_{j,t+1}^{\delta} - \hat{P}_{j,t}^{\delta}\right) - (1+i_1)\hat{i}_{1,t}, \quad \forall j,$$

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

$$+ \sum_{k=1}^{n} Cov\left(ex_{j,+1}, ex_{k,+1}^{\delta}\right)\left(q_k^{-1}\hat{b}_{ak,t}^{\delta} - q_k^{-1} b_{ak}^{\delta}\hat{q}_{k,t}\right)\Bigg], \quad \forall j > 1,$$

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

$$+ \sum_{k=1}^{n} Cov\left(ex_{j,+1}^{\delta}, ex_{k,+1}^{\delta}\right)\left(q_k^{-1}\hat{b}_{ak,t}^{\delta} - q_k^{-1} b_{ak}^{\delta}\hat{q}_{k,t}\right)\Bigg], \quad \forall j,$$

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

$$\hat{w}_{a,t+1} = (1-\xi)\Bigg[(1+i_1) w_a \hat{i}_{1,t} + (1+i_1)\hat{w}_{a,t}$$

$$+ \sum_{j=2}^{n} q_j^{-1}\left(b_{aj}\hat{ex}_{j,t+1} - ex_j b_{aj}\hat{q}_{j,t} + ex_j\hat{b}_{aj,t}\right)$$

$$+\sum_{j=1}^{n} q_j^{-1}\left(b_{aj}^{\delta}\hat{ex}_{j,t+1}^{\delta} - ex_j^{\delta} b_{aj}^{\delta}\hat{q}_{j,t} + ex_j^{\delta}\hat{b}_{aj,t}^{\delta}\right)\Bigg].$$

The bond market clearing conditions are:

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

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:

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

As stated in the claim, we will now characterize the effects of a U.S. monetary shock $\nu_{i1,t}$ 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:

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

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

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

$$\tilde{b}_{jk} \equiv b_{jk}/\kappa, \;\; \forall j, k \neq j.$$

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

$$1 = \beta_j(1 + i_j), \;\; \forall j,$$

$$i_j - i_1 = \frac{\Gamma}{w_a} \sum_{k=2}^{n} Cov\left(ex_{j,+1}, ex_{k,+1}\right) q_k^{-1} b_{ak}, \;\; \forall j > 1,$$

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

$$\sum_{j=1}^{n} q_j^{-1} b_{aj} = w_a,$$

$$w_a = \xi \omega_a + (1 - \xi)\left[(1 + i_1) w_a + \sum_{j=2}^{n} (i_j - i_1) q_j^{-1} b_{aj}\right].$$

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

$$0 = \beta^{-1}\left(\tilde{\beta}_j - \beta\right) + \beta \tilde{i}_1, \;\; \forall j,$$

$$\tilde{ex}_j = \frac{\tilde{\Gamma}}{\tilde{w}_a} \sum_{k=2}^{n} Cov\left(ex_{j,+1}, ex_{k,+1}\right) q_k^{-1} \tilde{b}_{ak}, \;\; \forall j > 1,$$

$$\tilde{ex}_j^\delta = \frac{\tilde{\Gamma}}{\tilde{w}_a} \sum_{k=2}^{n} Cov\left(ex_{j,+1}^\delta, ex_{k,+1}\right) q_k^{-1} \tilde{b}_{ak}, \;\; \forall j,$$

$$\sum_{j=1}^{n} q_j^{-1} \tilde{b}_{aj} = 0,$$

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

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

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

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:

$$\begin{aligned}
c_j &= 1, \ \ \forall j, \\
\ell_j &= 1, \ \ \forall j, \\
c_{jk} &= \eta_{jk}, \ \ \forall j, k, \\
p_{jk} &= 1, \ \ \forall j, k, \\
q_j &= 1, \ \ \forall j.
\end{aligned}$$

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

$$\mathbf{0}_n = -\frac{1}{\psi} E_t \Delta \hat{\boldsymbol{c}}_{t+1} + \left(\mathbf{1}_n - \boldsymbol{J}'_{n-1} \mathbf{1}_{n-1}\right) \nu_{i1,t}, \tag{27}$$

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

$$\boldsymbol{J}_{n-1} \hat{\boldsymbol{c}}_t + \beta \hat{\boldsymbol{a}}_t = \hat{\boldsymbol{q}}_t - \boldsymbol{J}_{n-1} \boldsymbol{\zeta}^{-1} (\boldsymbol{\zeta\eta})' \boldsymbol{J}'_{n-1} \hat{\boldsymbol{q}}_t + \boldsymbol{J}_{n-1} \boldsymbol{\zeta}^{-1} (\boldsymbol{\zeta\eta})' \hat{\boldsymbol{c}}_t + \hat{\boldsymbol{a}}_{t-1}, \tag{28}$$

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

$$\hat{\boldsymbol{ex}}_{t+1} = \beta^{-1} \left(-\Delta \hat{\boldsymbol{q}}_{t+1} - \mathbf{1}_{n-1} \nu_{i1,t} \right). \tag{29}$$

The arbitrageur optimality conditions for one-period bonds are

$$E_t \hat{\boldsymbol{ex}}_{t+1} = -\boldsymbol{ex} \frac{1}{w_a} \hat{w}_{a,t} - \beta \frac{\Gamma}{w_a} \boldsymbol{C} \boldsymbol{J}_{n-1} \boldsymbol{\zeta} \boldsymbol{J}'_{n-1} \hat{\boldsymbol{a}}_t, \tag{30}$$

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

$$\hat{w}_{a,t+1} = (1 - \xi) \left[\beta^{-1} \hat{w}_{a,t} + \boldsymbol{b}'_a \hat{\boldsymbol{ex}}_{t+1} - \boldsymbol{ex}' \boldsymbol{J}_{n-1} \boldsymbol{\zeta} \boldsymbol{J}'_{n-1} \hat{\boldsymbol{a}}_t \right], \tag{31}$$

where $\boldsymbol{ex}$ is a $(n-1) \times 1$ vector of ex_j; and $\boldsymbol{b}_a$ is a $(n-1) \times 1$ vector of b_{aj} for $j > 1$. Note here that we have eliminated terms of order κ^2 from the evolution of $\hat{w}_{a,t+1}$, but retained terms of order κ, as these will be relevant for risk pricing.[44]

Then, (27)-(31) 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 $\{\hat{\boldsymbol{a}}_{t-1}, \hat{w}_{a,t}, \nu_{i1,t}\}$ and the evolution of the exogenous state variable $\nu_{i1,t}$. We conjecture a solution of the form

$$\begin{aligned}
\hat{\boldsymbol{c}}_t &= \boldsymbol{\Delta}_a^c \hat{\boldsymbol{a}}_{t-1} + \boldsymbol{\delta}_w^c \hat{w}_{a,t} + \boldsymbol{\delta}_\nu^c \nu_{i1,t}, \\
\hat{\boldsymbol{q}}_t &= \boldsymbol{\Delta}_a^q \hat{\boldsymbol{a}}_{t-1} + \boldsymbol{\delta}_w^q \hat{w}_{a,t} + \boldsymbol{\delta}_\nu^q \nu_{i1,t}, \\
\hat{\boldsymbol{a}}_t &= \boldsymbol{\Delta}_a^a \hat{\boldsymbol{a}}_{t-1} + \boldsymbol{\delta}_w^a \hat{w}_{a,t} + \boldsymbol{\delta}_\nu^a \nu_{i1,t}, \\
\hat{w}_{a,t+1} &= \boldsymbol{\delta}_a^{Ew} \hat{\boldsymbol{a}}_{t-1} + \delta_w^{Ew} \hat{w}_{a,t} + \delta_\nu^{Ew} \nu_{i1,t} + \delta_\epsilon^w \epsilon_{\nu 1,t+1}, \\
\hat{\boldsymbol{ex}}_{t+1} &= \boldsymbol{\Delta}_a^{Eex} \hat{\boldsymbol{ex}}_{t-1} + \boldsymbol{\delta}_w^{Eex} \hat{w}_{a,t} + \boldsymbol{\delta}_\nu^{Eex} \nu_{i1,t} + \boldsymbol{\delta}_\epsilon^{ex} \epsilon_{\nu 1,t+1},
\end{aligned}$$

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 $\epsilon_{\nu 1,t}$ on

[44] This in turn is because they are multiplied by $\boldsymbol{ex} \frac{1}{w_a}$, which is of order κ^{-1}, in (30).

exchange rates. This is given by

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

By (28), we have that

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

where

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

assuming the expression in parenthesis is invertible. It follows that

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

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 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

$$\Bigg[\left(1-\beta(1-\xi)\left[\left(\beta^{-1} - \boldsymbol{b}_a'\boldsymbol{ex}\frac{1}{w_a}\right) - \left(\boldsymbol{ex}' + \beta\frac{\Gamma}{w_a}\boldsymbol{b}_a'\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}'\boldsymbol{\delta}_w^a\right]\right)\boldsymbol{I}_{n-1}$$
$$+\beta\left(\boldsymbol{I}_{n-1} - \boldsymbol{\Delta}_a^a\right) + \beta^2\frac{\Gamma}{w_a}\boldsymbol{Q}_{n-1}^{-1}\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}'\Bigg]\boldsymbol{\delta}_w^a = -\beta\boldsymbol{Q}_{n-1}^{-1}\boldsymbol{ex}\frac{1}{w_a}. \quad (32)$$

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

$$
\begin{aligned}
&(\beta\boldsymbol{\Delta}_a^a - \boldsymbol{I}_{n-1})\,(\boldsymbol{\Delta}_a^a - \boldsymbol{I}_{n-1}) \\
&\quad = \left[\beta(1-\xi)\boldsymbol{\delta}_w^a\left(\boldsymbol{e}\boldsymbol{x}' + \beta\frac{\Gamma}{w_a}\boldsymbol{b}_a'\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}' + \beta^2\frac{\Gamma}{w_a}\boldsymbol{Q}_{n-1}^{-1}\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}'\right]\boldsymbol{\Delta}_a^a. \quad (33)
\end{aligned}
$$

As is evident, in general we cannot solve (32) and (33) 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

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

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

$$
\boldsymbol{\delta}_\nu^c = -(1-\rho_{i1})^{-1}\psi\left(\mathbf{1}_n - \boldsymbol{J}_{n-1}'\mathbf{1}_{n-1}\right).
$$

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 and central banks which target inflation.

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

$$
\begin{aligned}
\boldsymbol{\delta}_\nu^a = &\left[(1-\beta\rho_{i1})\boldsymbol{I}_{n-1} + \beta\left(\boldsymbol{I}_{n-1} - \boldsymbol{\Delta}_a^a\right) + \beta(1-\xi)\boldsymbol{\delta}_w^a\left(\boldsymbol{e}\boldsymbol{x}' + \beta\frac{\Gamma}{w_a}\boldsymbol{b}_a'\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}'\right. \\
&\left. + \beta^2\frac{\Gamma}{w_a}\boldsymbol{Q}_{n-1}^{-1}\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}'\right]^{-1}(1-\psi)\boldsymbol{Q}_{n-1}^{-1}\mathbf{1}_{n-1}. \quad (34)
\end{aligned}
$$

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

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

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

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 (32)-(34) is

$$\lim_{\boldsymbol{ex},\Gamma\to 0} \boldsymbol{\Delta}_{a}^{a} = \boldsymbol{I}_{n-1},$$
$$\lim_{\boldsymbol{ex},\Gamma\to 0} \boldsymbol{\delta}_{w}^{a} = \boldsymbol{0}_{n-1},$$
$$\lim_{\boldsymbol{ex},\Gamma\to 0} \boldsymbol{\delta}_{\nu}^{a} = (1-\beta\rho_{i1})^{-1}(1-\psi)\boldsymbol{Q}_{n-1}^{-1}\boldsymbol{1}_{n-1}.$$

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

$$\lim_{\boldsymbol{ex},\Gamma\to 0}\left[\boldsymbol{\delta}_{w}^{q}\delta_{\epsilon}^{w} + \boldsymbol{\delta}_{\nu}^{q}\right] = \left[\beta^{-1}-\rho_{i1}\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_{i1})^{-1}\psi\right]\boldsymbol{1}_{n-1}.$$

We can then differentiate (32)-(34) with respect to $\boldsymbol{ex}$ and evaluate at the point of approximation, and we obtain

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

Using as well that

$$\lim_{\boldsymbol{ex},\Gamma\to 0} \delta_{\epsilon}^{w} = -(1-\xi)\left(\sum_{k=2}^{n} b_{ak}\right)\beta^{-1}\left[\beta^{-1}-\rho_{i1}\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_{i1})^{-1}\psi\right],$$

it follows that

$$\lim_{\boldsymbol{ex},\Gamma\to 0} \frac{d\left[\boldsymbol{\delta}_{w}^{q}\delta_{\epsilon}^{w} + \boldsymbol{\delta}_{\nu}^{q}\right]}{d\boldsymbol{ex}} = \frac{1-\xi}{\xi}\beta\frac{\sum_{k=2}^{n} b_{ak}}{w_{a}}\left[\beta^{-1}-\rho_{i1}\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_{i1})^{-1}\psi\right]\boldsymbol{I}_{n-1}.$$

Finally, we can differentiate (32)-(34) with respect to Γ/w_a and evaluate at the point of approximation, and we obtain

$$\lim_{\boldsymbol{ex},\Gamma\to 0} \frac{d\boldsymbol{\Delta}_a^a}{d[\Gamma/w_a]} = -(1-\beta)^{-1}\beta^2 \boldsymbol{Q}_{n-1}^{-1}\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}',$$
$$\lim_{\boldsymbol{ex},\Gamma\to 0} \frac{d\boldsymbol{\delta}_w^a}{d[\Gamma/w_a]} = \boldsymbol{0}_{n-1},$$
$$\lim_{\boldsymbol{ex},\Gamma\to 0} \frac{d\boldsymbol{\delta}_\nu^a}{d[\Gamma/w_a]} = -(\beta^{-1}-\rho_{i1})^{-2}(1-\beta)^{-1}\boldsymbol{Q}_{n-1}^{-1}\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\eta}_1'(1-\psi),$$

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

$$\lim_{\boldsymbol{ex},\Gamma\to 0} \frac{d\left[\boldsymbol{\delta}_w^q\delta_\epsilon^w + \boldsymbol{\delta}_\nu^q\right]}{d[\Gamma/w_a]} = -(\beta^{-1}-\rho_{i1})^{-2}(1-\beta)^{-1}\beta\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\eta}_1'(1-\psi).$$

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

$$\begin{aligned}\boldsymbol{\delta}_w^q\delta_\epsilon^w + \boldsymbol{\delta}_\nu^q = {} & \left[\beta^{-1}-\rho_{i1}\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_{i1})^{-1}\psi\right]\boldsymbol{1}_{n-1} \\ & + \frac{1-\xi}{\xi}\beta\frac{\sum_{k=2}^n b_{ak}}{w_a}\left[\beta^{-1}-\rho_{i1}\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_{i1})^{-1}\psi\right]\boldsymbol{ex} \\ & - (\beta^{-1}-\rho_{i1})^{-2}(1-\beta)^{-1}\beta\frac{\Gamma}{w_a}\boldsymbol{C}\boldsymbol{J}_{n-1}\boldsymbol{\eta}_1'(1-\psi).\end{aligned}$$

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

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

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

$$\hat{\boldsymbol{ex}}^{\delta}_{t+1} = -\beta^{-1}\boldsymbol{J}'_{n-1}\Delta\hat{\boldsymbol{q}}_{t+1} + \delta\hat{\boldsymbol{P}}^{\delta}_{t+1} - \beta^{-1}\hat{\boldsymbol{P}}^{\delta}_{t} - \beta^{-1}\mathbf{1}_n\nu_{i1,t}.$$

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

$$\hat{\boldsymbol{ex}}^{\delta}_{t+1} = \boldsymbol{J}'_{n-1}\hat{\boldsymbol{ex}}_{t+1} + \delta\hat{\boldsymbol{P}}^{\delta}_{t+1} - \beta^{-1}\hat{\boldsymbol{P}}^{\delta}_{t} - \beta^{-1}\left(\mathbf{1}_n - \boldsymbol{J}'_{n-1}\mathbf{1}_{n-1}\right)\nu_{i1,t}.$$

Combining these results and using (30) yields

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

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

$$\hat{\boldsymbol{P}}^{\delta}_{t} = \boldsymbol{\Delta}^{P^{\delta}}_{a}\hat{\boldsymbol{a}}_{t-1} + \boldsymbol{\delta}^{P^{\delta}}_{w}\hat{w}_{a,t} + \boldsymbol{\delta}^{P^{\delta}}_{\nu}\nu_{i1,t},$$

the method of undetermined coefficients implies

$$\begin{aligned}\boldsymbol{\Delta}^{P^{\delta}}_{a} &= \left[\delta\boldsymbol{\delta}^{P^{\delta}}_{w}\boldsymbol{\delta}^{Ew}_{a} + \beta\frac{\Gamma}{w_a}\left(\boldsymbol{C}^{\delta} - \boldsymbol{J}'_{n-1}\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}'_{n-1}\boldsymbol{\Delta}^{a}_{a}\right]\left(\beta^{-1}\boldsymbol{I}_{n-1} - \delta\boldsymbol{\Delta}^{a}_{a}\right)^{-1}, \\ \boldsymbol{\delta}^{P^{\delta}}_{w} &= \left[\delta\boldsymbol{\Delta}^{P^{\delta}}_{a}\boldsymbol{\delta}^{a}_{w} + \left(\boldsymbol{ex}^{\delta} - \boldsymbol{J}'_{n-1}\boldsymbol{ex}\right)\frac{1}{w_a} + \beta\frac{\Gamma}{w_a}\left(\boldsymbol{C}^{\delta} - \boldsymbol{J}'_{n-1}\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}'_{n-1}\boldsymbol{\delta}^{a}_{w}\right]\left(\beta^{-1} - \delta\delta^{Ew}_{w}\right)^{-1}, \\ \boldsymbol{\delta}^{P^{\delta}}_{\nu} &= \left[\delta\left(\boldsymbol{\Delta}^{P^{\delta}}_{a}\boldsymbol{\delta}^{a}_{\nu} + \boldsymbol{\delta}^{P^{\delta}}_{w}\delta^{Ew}_{\nu}\right) - \beta^{-1}\left(\mathbf{1}_n - \boldsymbol{J}'_{n-1}\mathbf{1}_{n-1}\right) + \beta\frac{\Gamma}{w_a}\left(\boldsymbol{C}^{\delta} - \boldsymbol{J}'_{n-1}\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}'_{n-1}\boldsymbol{\delta}^{a}_{\nu}\right]\times \\ &\quad \left(\beta^{-1} - \delta\rho_{i1}\right)^{-1}.\end{aligned}$$

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

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

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

$$\begin{aligned}
\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0} \boldsymbol{\Delta}_a^{P^{\delta}} &= \mathbf{0}_{n\times(n-1)}, \\
\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0} \boldsymbol{\delta}_w^{P^{\delta}} &= \mathbf{0}_{n}, \\
\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0} \boldsymbol{\delta}_\nu^{P^{\delta}} &= -\left(1-\beta\delta\rho_{i1}\right)^{-1}\left(\mathbf{1}_n - \boldsymbol{J}'_{n-1}\mathbf{1}_{n-1}\right).
\end{aligned}$$

It follows that

$$\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0}\left[\boldsymbol{\delta}_w^{P^{\delta}}\delta_\epsilon^w + \boldsymbol{\delta}_\nu^{P^{\delta}}\right] = -\left(1-\beta\delta\rho_{i1}\right)^{-1}\left(\mathbf{1}_n - \boldsymbol{J}'_{n-1}\mathbf{1}_{n-1}\right).$$

Moreover,

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

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

$$\begin{aligned}
&\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0}\frac{d\left[\boldsymbol{\delta}_w^{P^{\delta}}\delta_\epsilon^w + \boldsymbol{\delta}_\nu^{P^{\delta}}\right]}{d\boldsymbol{ex}} \\
&= \left(1-\delta(1-\xi)\right)^{-1}(1-\xi)\left(\frac{\sum_{k=2}^{n} b_{ak}}{w_a}\right)\left[\beta^{-1}-\rho_{i1}\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_{i1})^{-1}\psi\right]\boldsymbol{J}'_{n-1}.
\end{aligned}$$

Similarly,

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

It follows that

$$\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0} \frac{d\left[\boldsymbol{\delta}_w^{P^\delta}\delta_\epsilon^w + \boldsymbol{\delta}_\nu^{P^\delta}\right]}{d\boldsymbol{ex}^\delta}$$
$$= -\left(1-\delta(1-\xi)\right)^{-1}(1-\xi)\left(\frac{\sum_{k=2}^n b_{ak}}{w_a}\right)\left[\beta^{-1}-\rho_{i1}\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_{i1})^{-1}\psi\right]\boldsymbol{I}_n.$$

Finally,

$$\begin{aligned}
\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0} \frac{d\boldsymbol{\Delta}_a^{P^\delta}}{d[\Gamma/w_a]} &= (1-\beta\delta)^{-1}\left(\boldsymbol{C}^\delta - \boldsymbol{J}_{n-1}'\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\zeta}\boldsymbol{J}_{n-1}',\\
\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0} \frac{d\boldsymbol{\delta}_w^{P^\delta}}{d[\Gamma/w_a]} &= \boldsymbol{0}_n,\\
\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0} \frac{d\boldsymbol{\delta}_\nu^{P^\delta}}{d[\Gamma/w_a]} &= (\beta^{-1}-\rho_{i1})^{-1}\left(\beta^{-1}-\delta\rho_{i1}\right)^{-1}\left(\boldsymbol{C}^\delta - \boldsymbol{J}_{n-1}'\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\eta}_1'(1-\psi).
\end{aligned}$$

It follows that

$$\lim_{\boldsymbol{ex},\boldsymbol{ex}^{\delta},\Gamma\to 0} \frac{d\left[\boldsymbol{\delta}_w^{P^\delta}\delta_\epsilon^w + \boldsymbol{\delta}_\nu^{P^\delta}\right]}{d[\Gamma/w_a]} = (\beta^{-1}-\rho_{i1})^{-1}\left(\beta^{-1}-\delta\rho_{i1}\right)^{-1}\left(\boldsymbol{C}^\delta - \boldsymbol{J}_{n-1}'\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\eta}_1'(1-\psi).$$

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

$$\begin{aligned}
\boldsymbol{\delta}_w^{P^\delta}\delta_\epsilon^w + \boldsymbol{\delta}_\nu^{P^\delta} &= -\left(1-\beta\delta\rho_{i1}\right)^{-1}\left(\boldsymbol{1}_n - \boldsymbol{J}_{n-1}'\boldsymbol{1}_{n-1}\right)\\
&-\left(1-\delta(1-\xi)\right)^{-1}(1-\xi)\left(\frac{\sum_{k=2}^n b_{ak}}{w_a}\right)\left[\beta^{-1}-\rho_{i1}\right]^{-1}\left[1+(\beta^{-1}-1)(1-\rho_{i1})^{-1}\psi\right]\left(\boldsymbol{ex}^\delta - \boldsymbol{J}_{n-1}'\boldsymbol{ex}\right)\\
&+(\beta^{-1}-\rho_{i1})^{-1}\left(\beta^{-1}-\delta\rho_{i1}\right)^{-1}\frac{\Gamma}{w_a}\left(\boldsymbol{C}^\delta - \boldsymbol{J}_{n-1}'\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\eta}_1'(1-\psi).
\end{aligned}$$

Noting that

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

and again noting that

$$\left(\boldsymbol{C}^{\delta}-\boldsymbol{J}_{n-1}^{\prime}\boldsymbol{C}\right)\boldsymbol{J}_{n-1}\boldsymbol{\eta}_{1}^{\prime}=\left[\begin{array}{c} Cov\left(ex_{1,+1}^{\delta},\sum_{k=2}^{n}\eta_{1k}ex_{k,+1}\right) \\ Cov\left(ex_{2,+1}^{\delta}-ex_{2,+1},\sum_{k=2}^{n}\eta_{1k}ex_{k,+1}\right) \\ \vdots \\ Cov\left(ex_{n,+1}^{\delta}-ex_{n,+1},\sum_{k=2}^{n}\eta_{1k}ex_{k,+1}\right) \end{array}\right],$$

we obtain the claimed result in the proposition. $\square$

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

	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
CLP	Oct-05	Oct-05			Apr-91
COP	Aug-03	Oct-04			Aug-92
HUF*	Feb-99	Feb-99			Feb-99
IDR*	Feb-03	Jul-03			Jun-99
ILS	Mar-05	Mar-05			Apr-91
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).

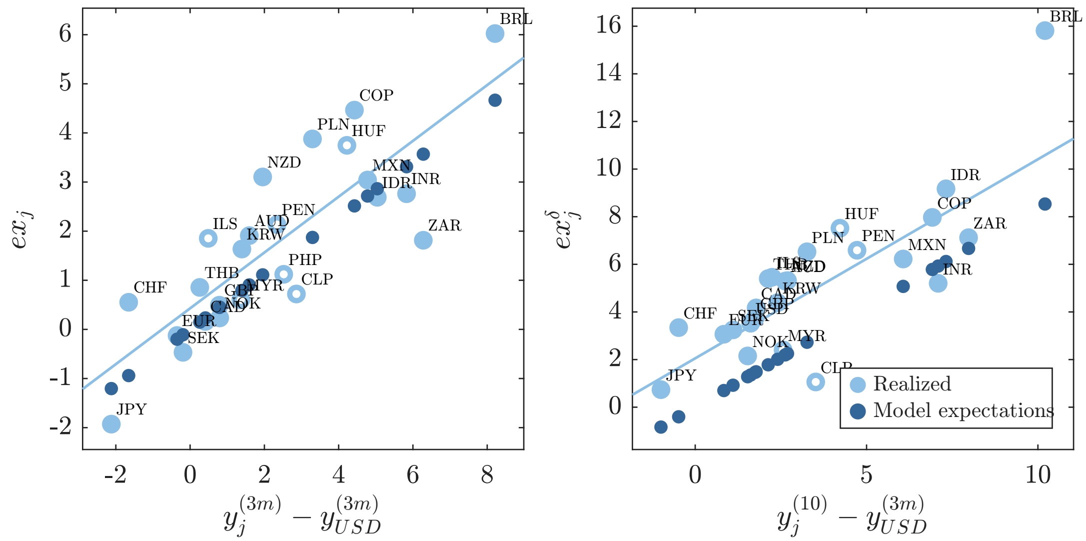

Figure 11: average realized excess returns vs. estimates of expected excess returns

Notes: all excess returns are annualized. 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.

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. We do not include the constant from these cross-sectional regressions in our fitted values. This is motivated by our view that the secular dollar appreciation and decline in global interest rates over our sample period may have been unexpected ex-ante.

Figure 11 compares average realized excess returns on currencies and long-term bonds against our estimates of expected excess returns. There are two messages. First, average interest rate differentials and yield curve slopes indeed do a very good job explaining the cross-section of average excess returns.[45] Second, eliminating the constant from the cross-sectional regressions when constructing expected returns is a conservative assumption, because it means that we generally estimate lower expected excess returns. It has a small quantitative effect for currencies, but is more meaningful

[45]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).

	CHF, EUR, JPY	Other non-U.S. G10	EM	Sum
$\Gamma = 1$	-1.9	-2.4	8.8	4.5
$\Gamma = 1$, fitted with intercept	-2.3	-1.7	10.1	6.1
$\Gamma = 1$, realized mean	-5.2	-2.9	12.2	4.1

Table 5: mean-variance efficient portfolio given alternative estimates of expected excess returns

Notes: first row reports baseline results with expected returns estimated as in prior subsection.

for long-term bond returns given the large decline in global interest rates in sample.

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 Table 5 reports the sensitivity of the mean-variance efficient portfolio using alternative estimates of expected excess returns. In our baseline specification reported in the first row (and in Table 1), we use expected excess returns as constructed in the prior subsection: using the fitted value from cross-sectional regressions of realized average excess returns on interest rate differentials and yield curve slopes, and removing the intercept. The second row of Table 5 retains the intercept, and third row simply uses the realized average excess returns. The basic structure of the mean-variance efficient portfolio is unchanged across these cases.

Figure 12 shows that the cross-sectional effects of U.S. monetary policy surprises are similarly robust when using realized rather than predicted returns along the x-axies. As in Figure 1 in the main text, currencies with higher average realized excess returns depreciate more, and long-dated yields for currencies with higher average

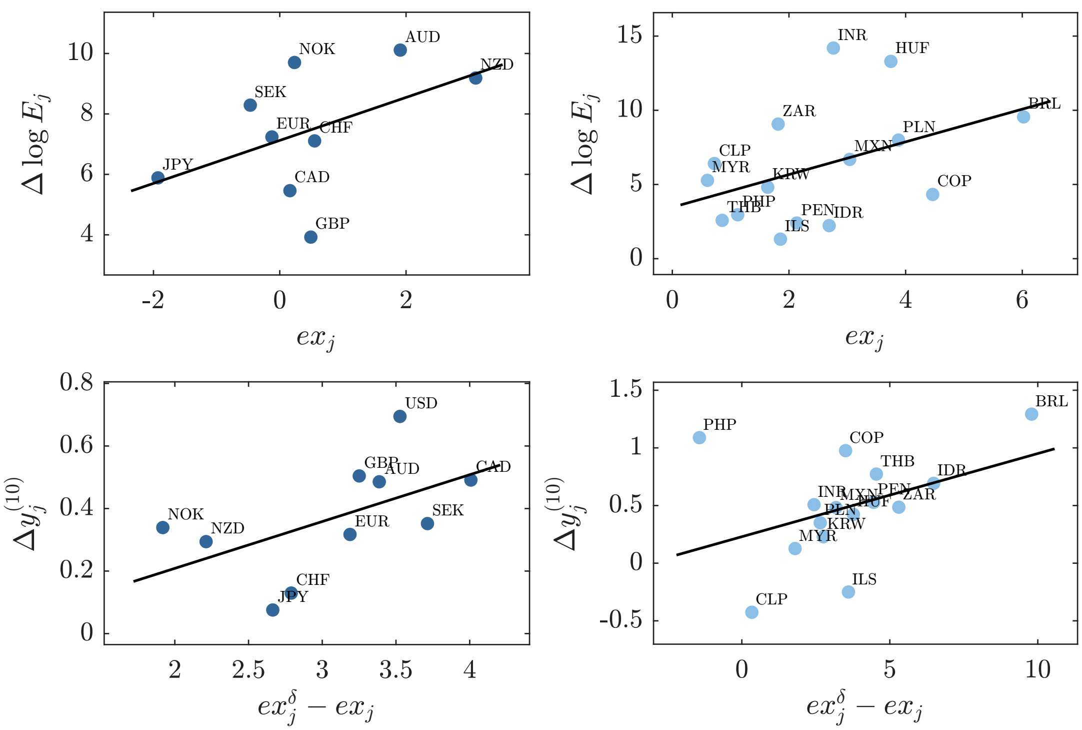

Figure 12: asset price responses to U.S. monetary tightening in data against average realized excess returns

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 $1pp$. Average realized excess returns along x-axis are annualized.

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 1.

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 markets (e.g., Lustig et al. (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

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

where $\boldsymbol{ex}_t$, $\boldsymbol{\alpha}$, and $\boldsymbol{\epsilon}_t$ are $n \times 1$ vectors, $\boldsymbol{\beta}$ is a $n \times k$ matrix, and $\boldsymbol{f}_t$ is a $k \times 1$ vector. We can further take expectations to yield

$$\boldsymbol{ex} = \boldsymbol{\alpha} + \boldsymbol{\beta\lambda},$$

where $\lambda^k \equiv E[f_t^k]$ can be interpreted as the k-th factor's price of risk, and $\boldsymbol{\lambda}$ is a $k \times 1$ vector of these factor prices stacked up. Furthermore, the conditional covariance matrix of excess returns is

$$\boldsymbol{C} = \boldsymbol{\beta F \beta}' + \text{diag}\left(\sigma_1^2, \ldots, \sigma_n^2\right),$$

where $\boldsymbol{F}$ is conditional covariance matrix of the factors, and where we assume constant conditional variances of innovations. Denoting $\boldsymbol{D} \equiv \text{diag}\left(\sigma_1^2, \ldots, \sigma_n^2\right)$, we can write this as

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

This factor-based estimate of the covariance matrix also allows us to understand the drivers of the efficient portfolio. By the Woodbury identity, we have that

$$\boldsymbol{C}^{-1} = \boldsymbol{D}^{-1} - \boldsymbol{D}^{-1}\boldsymbol{\beta}\left(\boldsymbol{F}^{-1} + \boldsymbol{\beta}'\boldsymbol{D}^{-1}\boldsymbol{\beta}\right)^{-1}\boldsymbol{\beta}'\boldsymbol{D}^{-1}.$$

It follows that the mean-variance efficient portfolio shares are given by

$$\begin{aligned} \frac{1}{\Gamma}\boldsymbol{C}^{-1}\boldsymbol{ex} &= \frac{1}{\Gamma}\left(\boldsymbol{D}^{-1} - \boldsymbol{D}^{-1}\boldsymbol{\beta}\left(\boldsymbol{F}^{-1} + \boldsymbol{\beta}'\boldsymbol{D}^{-1}\boldsymbol{\beta}\right)^{-1}\boldsymbol{\beta}'\boldsymbol{D}^{-1}\right)(\boldsymbol{\alpha} + \boldsymbol{\beta\lambda}), \\ &= \frac{1}{\Gamma}\left(\boldsymbol{D}^{-1}\boldsymbol{\alpha} + \boldsymbol{D}^{-1}\boldsymbol{\beta}\left(\boldsymbol{F}^{-1} + \boldsymbol{\beta}'\boldsymbol{D}^{-1}\boldsymbol{\beta}\right)^{-1}\left(\boldsymbol{F}^{-1}\boldsymbol{\lambda} - \boldsymbol{\beta}'\boldsymbol{D}^{-1}\boldsymbol{\alpha}\right)\right), \end{aligned} \tag{35}$$

where the second equality follows from straightforward matrix algebra. This motivates focusing on two components of the mean-variance efficient portfolio. The first, $\boldsymbol{D}^{-1}\boldsymbol{\beta}\left(\boldsymbol{F}^{-1}+\boldsymbol{\beta}'\boldsymbol{D}^{-1}\boldsymbol{\beta}\right)^{-1}\boldsymbol{F}^{-1}\boldsymbol{\lambda}$, summarizes the efficient portfolio due only to exposures to common factors ($\boldsymbol{\alpha}=\mathbf{0}$). The second, $\boldsymbol{D}^{-1}\boldsymbol{\alpha}$, summarizes the efficient portfolio in the absence of any loading of excess returns on common factors ($\boldsymbol{\beta}=\mathbf{0}$). The investor would go especially long in currencies offering high expected returns, α_j, relative to their (idiosyncratic) risk σ_j^2.

We first implement this approach using as factors the first k principal components (PCs) of excess currency returns. That is, we use standard PC analysis to define $f_t^1,\ldots,f_t^k$. Let $\boldsymbol{C}$ denote the conditional covariance matrix of $\boldsymbol{ex}_{t+1}$ estimated from daily data, and let $\boldsymbol{V}_k$ collect the first k eigenvectors of $\boldsymbol{C}$ (ordered by descending eigenvalue). We define PC factor returns as the returns on the corresponding PC portfolios,

$$\boldsymbol{f}_t\equiv\boldsymbol{V}_k'\boldsymbol{ex}_t.$$

Under this definition, the factor loadings are $\boldsymbol{\beta}=\boldsymbol{V}_k$ and the factor prices are $\boldsymbol{\lambda}\equiv E[\boldsymbol{f}_t]=\boldsymbol{V}_k'\boldsymbol{ex}$. We then define $\boldsymbol{\alpha}$ as the component of average realized excess returns orthogonal to the span of the first k PCs,

$$\boldsymbol{\alpha}\equiv\boldsymbol{ex}-\boldsymbol{\beta}\boldsymbol{\lambda}=\boldsymbol{ex}-\boldsymbol{V}_k\boldsymbol{V}_k'\boldsymbol{ex}.$$

Finally, we estimate the residual covariance matrix and set all non-diagonal elements to zero to obtain our estimate of $\boldsymbol{D}$.

Table 6 describes the share of variance in currency returns explained by the principal components as well as the factor loadings β_j^k. As argued in Lustig et al. (2011), the loadings on the principal components suggest that the first factor essentially captures the average excess return to non-dollar currencies, while the second captures the return to the carry trade, with quite negative loadings on low interest rate currencies like Japan and Switzerland, and less negative or positive loadings on high interest rate currencies like Australia, New Zealand, and emerging markets.

Table 7 decomposes the mean-variance efficient portfolio shares using $\Gamma=1$ and the first two, three, or four factors. The first line reports the shares given average realized excess currency returns and the covariance matrix estimated as described in the main text.[46] The second line reports the mean-variance efficient portfolio shares

[46] The only difference from the 4th row of Table 1 in the main text 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.

Cumulative share of $Var(ex_{j,t+1})$ explained by k PCs					
k	Share				
1	0.411				
2	0.532				
3	0.599				
4	0.662				

Estimated β_j^k					
k	JPY	CHF	NZD	AUD	EM (avg)
1	0.055	0.188	0.305	0.318	0.188
2	-0.253	-0.347	-0.088	-0.050	0.142
3	-0.121	-0.197	0.139	0.144	0.112
4	0.077	0.051	0.067	0.063	0.096

Table 6: principal component analysis of excess currency returns

given average excess returns and the covariance matrix estimated using $k = 2$ factors, as described in the prior paragraphs. Imposing this structure on the covariance matrix leaves the qualitative features of the efficient portfolio unchanged. The third and fourth lines report the mean-variance efficient portfolio shares implied when $\boldsymbol{\alpha} = \mathbf{0}$ or $\boldsymbol{\beta} = \mathbf{0}$ in (35). The message is that the dominant role of dollar funding is driven by both factor loadings and by average excess returns not explained by the loadings.[47] The remaining rows of the table demonstrate that similar results are obtained using three or four factors as well.

We next repeat the factor-based approach using dollar and carry factors as in Lustig et al. (2011). At each quarter end, we sort currencies into six portfolios based on their interest rate differential versus the U.S. We define the next quarter's return on each portfolio as the equal-weighted average of excess currency returns within the portfolio. Following Lustig et al. (2011), we define two traded factors: the average excess currency return across all currencies (the dollar factor), and the excess currency return on the highest interest rate-sorted portfolio less the lowest interest rate-sorted

[47] As described in the footnote of the table, using the Gibbons, Ross, and Shanken (1989) test, we cannot reject the null that these alphas are zero. However, it is still notable that at the point estimates, $\sum_{j=2}^{n} \frac{1}{\sigma_j^2} \alpha_j > 0$ — that is, the efficient portfolio based on alphas alone features a positive position in non-dollar currencies.

	CHF, EUR, JPY	Other non-U.S. G10	EM	Sum
$\Gamma = 1$, no 10-year bonds	-2.2	-4.9	12.5	5.4
$\Gamma = 1$, $k = 2$ factor model	-0.7	-5.5	16.4	10.2
$\boldsymbol{\alpha} = \mathbf{0}$	-1.5	0.5	3.2	2.2
$\boldsymbol{\beta} = \mathbf{0}$	-0.3	-7.5	13.1	5.3
$\Gamma = 1$, $k = 3$ factor model	0.8	-6.4	17.1	11.5
$\boldsymbol{\alpha} = \mathbf{0}$	-1.8	0.6	4.0	2.7
$\boldsymbol{\beta} = \mathbf{0}$	1.5	-7.7	14.9	8.7
$\Gamma = 1$, $k = 4$ factor model	-0.0	-7.0	13.1	6.0
$\boldsymbol{\alpha} = \mathbf{0}$	-1.4	0.7	5.9	5.3
$\boldsymbol{\beta} = \mathbf{0}$	0.2	-9.9	3.3	-6.5

Table 7: mean-variance efficient portfolio given PC factors

Notes: Gibbons-Ross-Shanken test statistic for jointly zero alphas has p-value 0.991 for $k = 2$, 0.996 for $k = 3$, and 0.999 for $k = 4$.

portfolio (the carry factor). We also consider an implementation with fixed portfolios over the sample based on the average interest rate differential over the entire period. In either case, denoting the factors $\bar{ex}_t$ and hml_t, we estimate factor loadings and alphas using standard time series regressions

$$ex_{j,t} = \alpha_j + \beta_j^1 \bar{ex}_t + \beta_j^2 hml_t + \epsilon_{j,t},$$

and estimate σ_j^2 using the variance of these residuals. We estimate the currency covariance matrix as

$$\boldsymbol{C} = \boldsymbol{\beta}\boldsymbol{F}\boldsymbol{\beta}' + \operatorname{diag}\left(\sigma_1^2, \ldots, \sigma_n^2\right),$$

where $\boldsymbol{\beta}$ and $\boldsymbol{F}$ are matrices of factor loadings and the covariance matrix of factors as above. We finally estimate the risk prices of each factor using the time series averages of $\bar{ex}_t$ and hml_t.

Table 8 decomposes the mean-variance efficient portfolio shares using $\Gamma = 1$ and the dollar and carry factors. The efficient portfolio is again largely short the dollar and long emerging markets. While both factor loadings and alphas contribute to

	CHF, EUR, JPY	Other non-U.S. G10	EM	Sum
$\Gamma = 1$, no 10-year bonds	-2.2	-4.9	12.5	5.4
$\Gamma = 1$, re-sorted portfolios	-2.4	-4.6	15.0	8.0
$\boldsymbol{\alpha} = \mathbf{0}$	-4.0	0.1	4.5	0.6
$\boldsymbol{\beta} = \mathbf{0}$	11.6	4.8	11.8	28.2
$\Gamma = 1$, fixed portfolios	-4.6	-3.6	16.3	8.1
$\boldsymbol{\alpha} = \mathbf{0}$	-0.6	0.5	1.9	1.7
$\boldsymbol{\beta} = \mathbf{0}$	-4.6	-3.2	17.0	9.1

Table 8: mean-variance efficient portfolio given dollar and carry factors

Notes: Gibbons-Ross-Shanken test statistic for jointly zero alphas has p-value 0.36 for re-sorted portfolios, and 0.59 for fixed portfolios.

	1	2	3	4	5	6	Sum
$\Gamma = 1$, re-sorted portfolios	-4.0	-2.5	2.0	-0.1	-3.0	7.7	0.2
$\Gamma = 1$, fixed portfolios	-0.9	-4.2	-11.2	9.3	2.5	2.3	-2.2

Table 9: mean-variance efficient portfolio given dollar and carry factors and sorted portfolios as test assets

Notes: portfolio 1 refers to currencies with lowest interest rate differential versus U.S., and portfolio 6 refers to currencies with highest interest rate differential versus U.S. Gibbons-Ross-Shanken test statistic for jointly zero alphas has p-value 0.72 for re-sorted portfolios, and 0.48 for fixed portfolios.

this portfolio structure, it is largely the alphas which drive this portfolio structure. Consistent with this result, Table 9 reports the efficient portfolio shares when we focus on the six portfolios of currencies sorted by interest rates, rather than the individual currencies themselves. In this case, the efficient portfolio is short the lowest interest rate portfolio and long the highest interest rate portfolio — i.e., it loads on the carry factor — with a dollar position that is small or even positive. We conclude that the alphas on individual currencies are an important driver of the dollar-funded portfolio we obtain in Table 8 and earlier results.

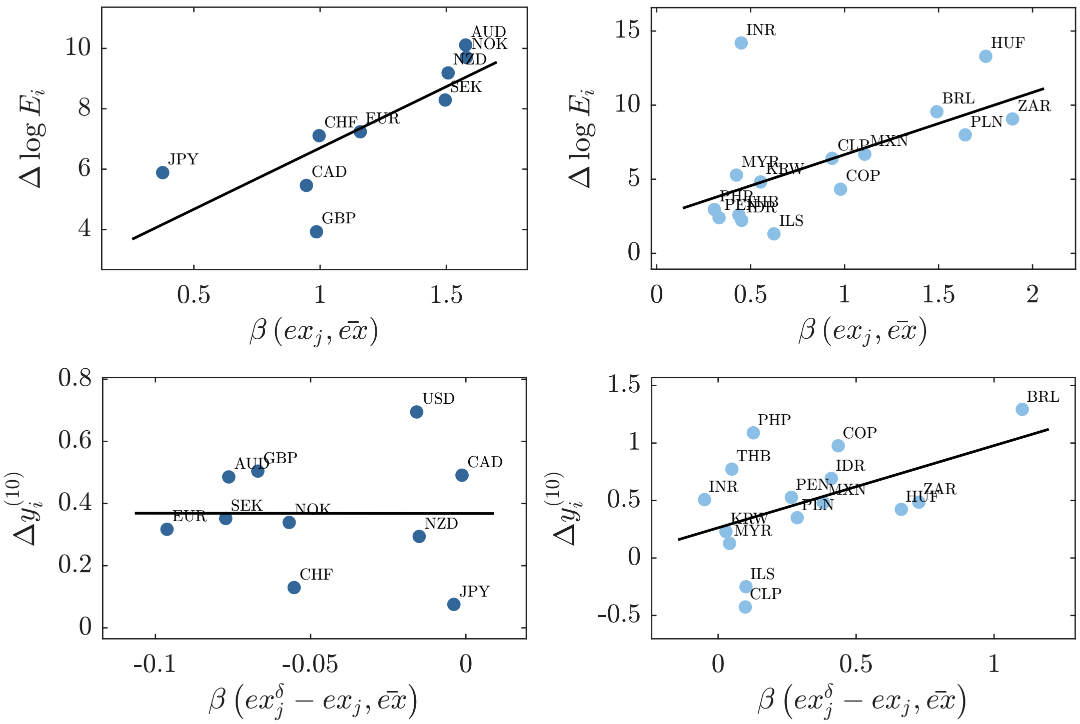

Figure 13: asset price responses to U.S. monetary tightening against covariances with average currency carry return

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 $1pp$. $\beta(y, x) \equiv Cov(y, x)/Var(x)$, and $\bar{ex}$ denotes the simple average of currency carry returns.

C.4 Evidence against portfolio balance

Figure 13 depicts the same responses of international asset prices to U.S. monetary shocks as in Figure 1, 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 emerging market bonds,[48]

[48] 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.

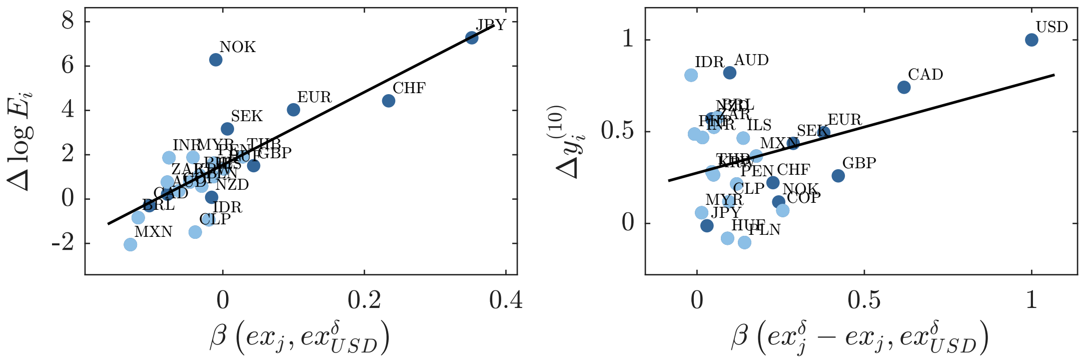

Figure 14: asset price responses to U.S. quantitative tightening against covariances with average U.S. bond carry return

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 $1pp$. $\beta(y,x) \equiv Cov(y,x)/Var(x)$.

nor for either set of currencies. 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 currency carry trade.

As noted in the main text, 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.

C.5 Global effects of quantitative tightening

The results of the last subsection suggest that the portfolio balance mechanism does not explain well the asset price responses to U.S. interest rate surprises. Here we study surprises to large-scale asset purchases (LSAP) and demonstrate that the spillovers *are* well explained by the portfolio balance mechanism. This 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.

We measure LSAP surprises using Swanson (2021)'s LSAP factor using the high-frequency response of long yields in tight intraday windows around Federal Reserve announcements. We project two-day changes in global yields and exchange rates on two-day changes in the U.S. 10-year yield, instrumented with the intraday Swanson

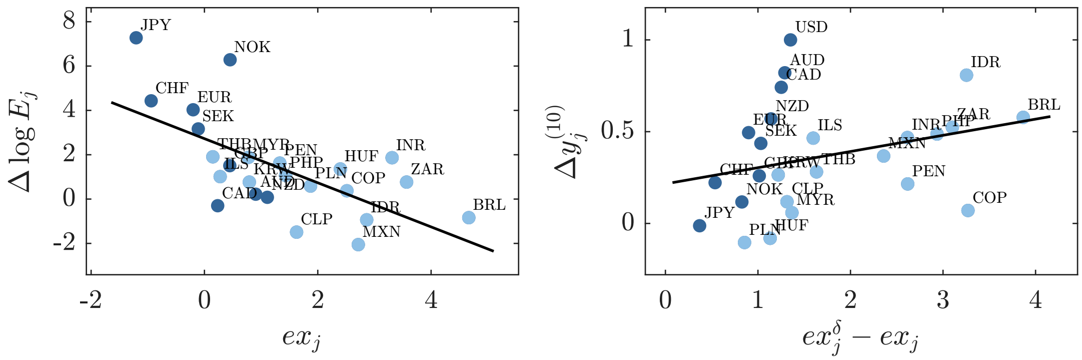

Figure 15: asset price responses to U.S. quantitative tightening against average excess returns

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*. Expected excess returns along x-axis are annualized.

(2021) LSAP surprise. Figure 14 depicts the estimated effects corresponding to a 1*pp* increase in the U.S. 10-year yield, and plots them against covariances with the U.S. bond carry return. Figure 15 depicts the same effects against average excess currency and local currency bond carry returns.

The responses to LSAP surprises are consistent with the portfolio balance mechanism. In response to more long-term dollar bonds on arbitrageur balance sheets, the portfolio balance mechanism predicts that 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). Figure 14 demonstrates 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. 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 15.

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), the main effects on global risk premia following U.S. quantitative tightening operate through the portfolio balance mechanism. In response to interest rate policy, the latter effects are instead much weaker.

D Additional quantitative results

In this appendix we provide additional results accompanying those in sections 5-5.5.

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 all of these calibrated parameter values and, where relevant, the associated model moments.

The model features the U.S., nine other G10 economies, nine emerging markets, and one "rest of world" (ROW) which is a residual ensuring market clearing given the targets for all other economies.

Table 10 provides the population parameters in each economy, which are targeted to match relative consumption from the IMF International Financial Statistics. The ROW has population such that world consumption is 3.5 times U.S. consumption.

Table 11 provides the discount factors in each economy and consol positions of U.S. households, which are targeted to match the one-month dollar interest rate and excess returns estimated using Bloomberg data. We set β_{ROW} so that the average β_j globally is the same as that in the U.S., and $\bar{b}^{\delta}_{1ROW} = 0$ for simplicity. 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.

Finally, Table 12 provides the taste parameters governing trade shares, which are targeted to match expenditure shares by origin from the IMF International Financial Statistics and Direction of Trade Statistics. 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.

		$\zeta_j q_j^{-1} c_j / c_1$	
Country	ζ_j	Target	Achieved
USD	1	normalization	
AUD	0.05	0.05	0.05
CAD	0.08	0.08	0.08
CHF	0.03	0.03	0.03
EUR	0.66	0.66	0.66
GBP	0.17	0.17	0.17
JPY	0.34	0.34	0.34
NOK	0.02	0.02	0.02
NZD	0.01	0.01	0.01
SEK	0.02	0.02	0.02
BRL	0.09	0.09	0.09
IDR	0.04	0.04	0.04
INR	0.09	0.09	0.09
KRW	0.06	0.06	0.06
MXN	0.07	0.07	0.07
PLN	0.02	0.02	0.02
ZAR	0.02	0.02	0.02
THB	0.02	0.02	0.02
COP	0.02	0.02	0.02
ROW	0.70	0.70	0.70

Table 10: baseline ζ_j parameters

Notes: data targets estimated over 1991 Q2 – 2024 Q1.

		i_j			$52\log((1+\delta P_j^\delta)/P_j^\delta)$	
Country	β_j^{52}	Target	Achieved	$\bar{b}_{1j}^\delta/(52\cdot y_1)$	Target	Achieved
USD	0.9901	1.00%	1.00%	9.27%	2.35%	2.35%
AUD	0.9811	1.90%	1.90%	1.40%	3.20%	3.19%
CAD	0.9877	1.23%	1.23%	4.53%	2.49%	2.49%
CHF	0.9994	0.06%	0.06%	0.65%	0.60%	0.60%
EUR	0.9920	0.80%	0.80%	0.08%	1.70%	1.70%
GBP	0.9856	1.45%	1.45%	-0.18%	2.46%	2.46%
JPY	1.0021	-0.21%	-0.21%	2.11%	0.16%	0.16%
NOK	0.9856	1.45%	1.45%	0.32%	2.28%	2.28%
NZD	0.9791	2.11%	2.11%	0.07%	3.25%	3.25%
SEK	0.9911	0.89%	0.89%	3.90%	1.92%	1.92%
BRL	0.9450	5.66%	5.66%	-2.32%	9.53%	9.52%
IDR	0.9621	3.86%	3.86%	-2.45%	7.12%	7.11%
INR	0.9578	4.31%	4.31%	-8.36%	6.93%	6.92%
KRW	0.9822	1.79%	1.79%	4.24%	3.01%	3.01%
MXN	0.9635	3.71%	3.71%	-0.30%	6.07%	6.07%
PLN	0.9717	2.87%	2.87%	-2.10%	3.73%	3.73%
ZAR	0.9554	4.57%	4.57%	-0.80%	7.67%	7.66%
THB	0.9886	1.15%	1.15%	6.57%	2.78%	2.78%
COP	0.9655	3.51%	3.51%	0.37%	6.78%	6.78%
ROW	1.0018	-0.18%	-0.18%	0.00%	1.00%	1.00%

Table 11: baseline β_j and $\bar{b}_{1j}^\delta$ parameters

Notes: data targets estimated 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 $1 + i_j + (ex_j^\delta - ex_j)$, where $ex_j^\delta - ex_j$ is estimated as described in section 5.

Country	η_{j1}	$P_{j1}c_{j1}/(P_jc_j)$ Target	Achieved	η_{1j}	$P_{1j}c_{1j}/(P_1c_1)$ Target	Achieved
USD	87.35%	87.35%	87.35%	87.35%	87.35%	87.35%
AUD	2.93%	2.93%	2.93%	0.06%	0.06%	0.06%
CAD	23.13%	23.13%	23.13%	2.19%	2.19%	2.19%
CHF	4.30%	4.30%	4.30%	0.15%	0.15%	0.15%
EUR	1.87%	1.87%	1.87%	1.86%	1.86%	1.86%
GBP	2.50%	2.50%	2.50%	0.41%	0.41%	0.41%
JPY	1.70%	1.70%	1.70%	1.03%	1.03%	1.03%
NOK	1.61%	1.61%	1.61%	0.04%	0.04%	0.04%
NZD	3.34%	3.34%	3.34%	0.02%	0.02%	0.02%
SEK	1.44%	1.44%	1.44%	0.08%	0.08%	0.08%
BRL	2.71%	2.71%	2.71%	0.16%	0.16%	0.16%
IDR	1.24%	1.24%	1.24%	0.09%	0.09%	0.09%
INR	1.25%	1.25%	1.25%	0.17%	0.17%	0.17%
KRW	4.57%	4.57%	4.57%	0.36%	0.36%	0.36%
MXN	18.62%	18.62%	18.62%	1.65%	1.65%	1.65%
PLN	0.63%	0.63%	0.63%	0.02%	0.02%	0.02%
ZAR	1.60%	1.60%	1.60%	0.04%	0.04%	0.04%
THB	4.14%	4.14%	4.14%	0.14%	0.14%	0.14%
COP	6.04%	6.04%	6.04%	0.08%	0.08%	0.08%
ROW	8.66%	8.66%	8.66%	4.09%	4.09%	4.09%

Table 12: baseline η_{j1} and η_{1j} parameters

Notes: data targets estimated over 1991 Q2 – 2024 Q1.

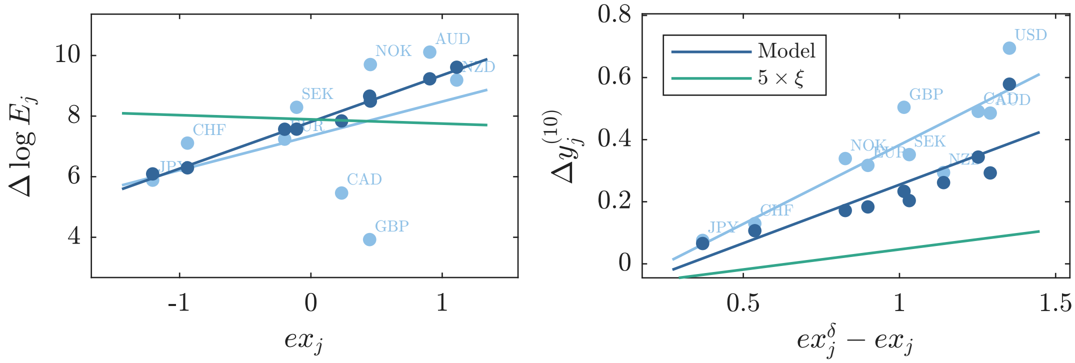

Figure 16: identification of arbitrageur exit rate

Notes: both model and data responses reported for surprise that raises two-year dollar yield by $1pp$. Expected excess returns along x-axis are annualized.

D.2 Identification of ξ

We estimate $\xi = 0.026$, implying a half-life of arbitrageurs that is roughly two quarters. Figure 16 demonstrates that, like arbitrageur risk aversion Γ, this parameter allows the model to match the cross-section of asset price responses for the G10 economies and how they line up with unconditional average excess returns. Consistent with Propositions 1 and 2, at a higher value of ξ and thus lower persistence of arbitrageur wealth, 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. 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 wealth process

As described in the main text, the baseline model implies that arbitrageurs endogenously earn high returns in the periods following a U.S. tightening, dampening the initial response of asset prices to the shock. Here we describe the calibration of an alternative version of the model in which arbitrageurs' wealth process is

$$W_{a,t+1} = \xi P_{1,t+1}\omega_a + (1-\xi)\left((1+i_{1,t})W_{a,t} + \Pi_{a,t+1} - E_t\Pi_{a,t+1}\right),$$

and thus they do not earn these expected excess returns ex-post (because they are paid out to households).

We re-calibrate $\{\Gamma, \xi, \rho_{\nu i1}, \omega_a\}$ to match the same targets as in the baseline model (it is largely only Γ and ω_a which change). Figure 17 depicts the asset price responses in this calibration, which are quite comparable to those in the baseline model (in Figure 2), and thus still comparable to the data. Figure 18 demonstrates the resulting effects of a U.S. tightening on real quantities are on impact comparable to those in the baseline model (in Figure 7) but inherit the higher persistence of arbitrageur wealth.

D.4 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, $\bar{b}_{11}^{\delta}$, but zero positions in other currencies' short or consol bonds, $\bar{b}_{1j} = \bar{b}_{1j}^{\delta} = 0$ 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, $\bar{b}_{jj}^{\delta}$ and $\bar{b}_{j1}$ for $j > 1$, while holding zero positions in other currencies' short or consol bonds. We use $\bar{b}_{jj}^{\delta}$ 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 19 demonstrates that the asset price responses to a U.S. tightening are comparable to those in Figure 2. Similarly, Figure 20 shows that the real effects on G10 and EM countries are similar to those in Figure 7. 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

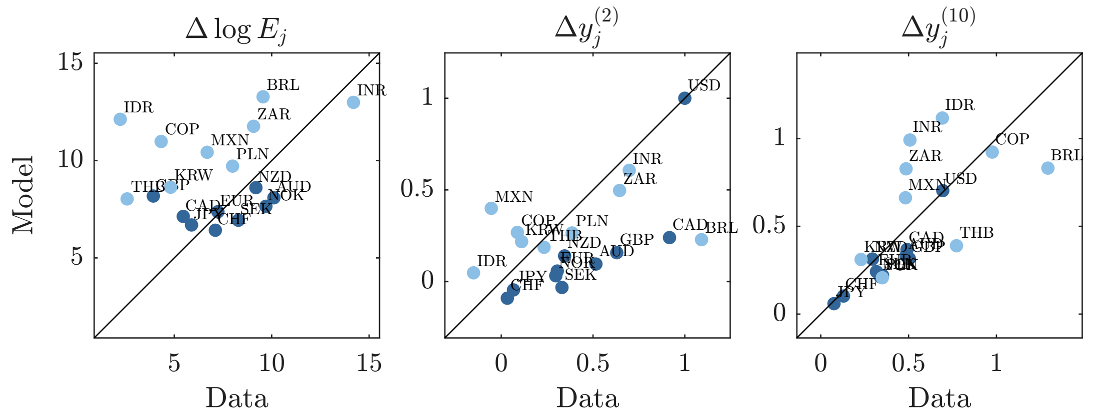

Figure 17: asset price responses to U.S. tightening given alternative wealth process

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

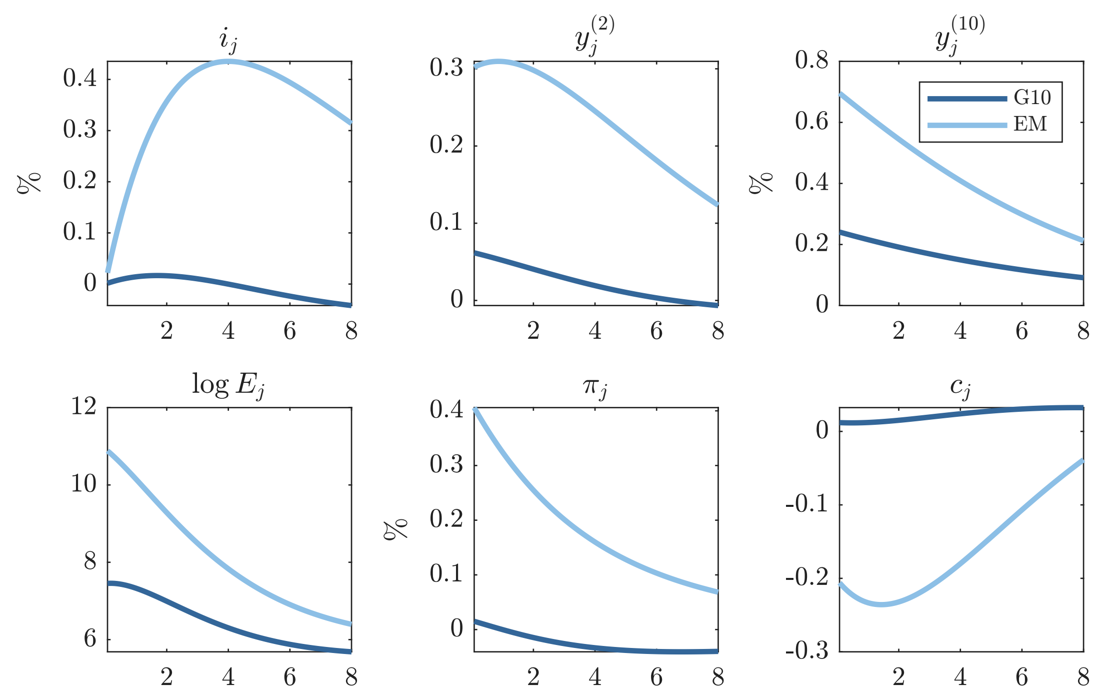

Figure 18: impulse responses to U.S. tightening given alternative wealth process

Notes: responses reported for same $\nu_{i1,t}$ innovation depicted in Figure 2. x-axis denotes quarters since the shock.

borrowing in local currencies (that is, $\sum_{j>1} \bar{b}_{j1} > 0$). This is consistent with estimates of the currency composition of NFAs (e.g., Allen and Juvenal (2024)). Notably, the magnitude of dollar-denominated NFAs are comparable between model and data: the rest of the world's dollar-denominated NFA is 28% of annual U.S. GDP in the model, as compared to 65% in the data from Allen and Juvenal (2024). 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.

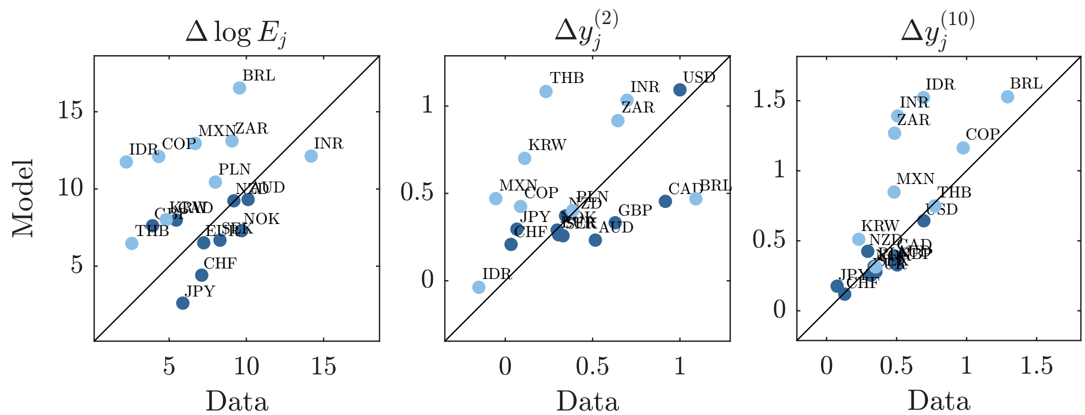

Figure 19: asset price responses to U.S. monetary tightening in calibration targeting NFAs

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

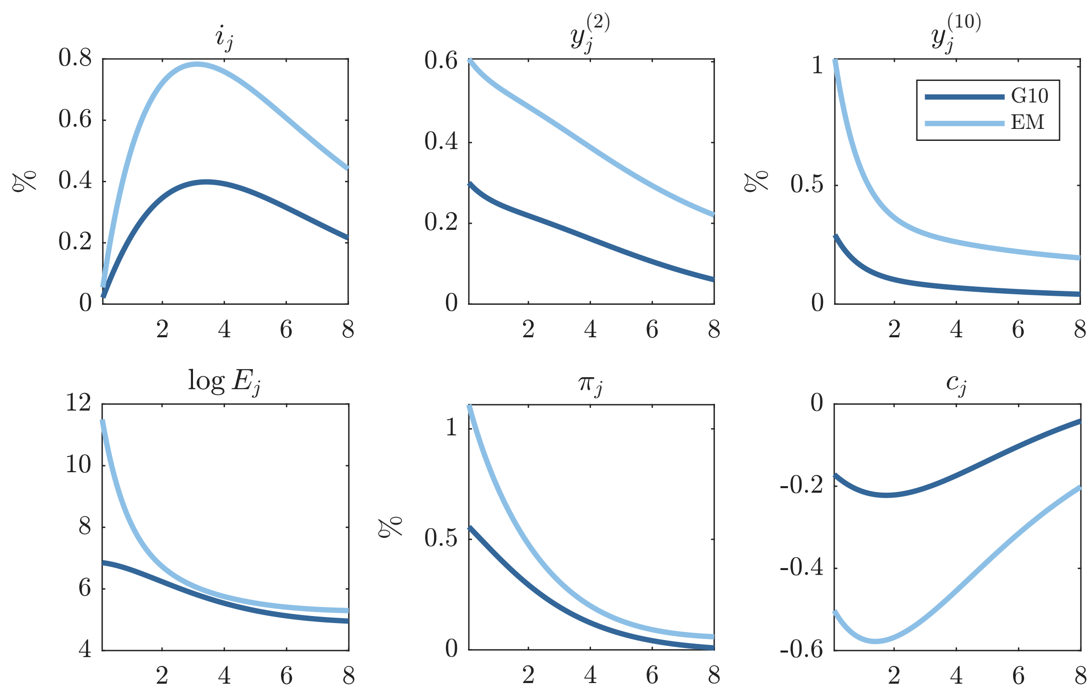

Figure 20: impulse responses to U.S. monetary tightening in calibration targeting NFAs

Notes: responses reported for same $\nu_{i1,t}$ innovation depicted in Figure 2. x-axis denotes quarters since the shock.