Global risk and the dollar[*]

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Abstract

The dollar is a safe-haven currency and dominates the global financial system. We investigate its role for the transmission of global risk within a Bayesian Proxy-SVAR. We identify global risk shocks using high-frequency surprises in the price of gold—the ultimate safe asset—around narratively selected events. Global risk shocks appreciate the dollar, induce a synchronized contraction of global economic activity and tighter global financial conditions. We benchmark these effects against a counterfactual in which the dollar does not appreciate. In this case, the contractionary impact of a global risk shock is much weaker, notably outside of US. We then put forward a two-country DCP^2 model of the world economy which features dollar dominance in trade and finance. We show that both aspects are necessary to account for the evidence on the effects of global risk shocks.

Keywords: Dollar dominance, global risk shocks, international transmission, Bayesian proxy structural VAR, minimum relative entropy counterfactual, DCP^2 model

JEL-Classification: F31, F42, F44

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1 Introduction

According to the received wisdom the dollar appreciates when global risk goes up. The Global Financial Crisis (GFC) and the COVID-19 pandemic are striking examples. We illustrate this in Figure 1, which shows the expected volatility index (VIX) as a proxy for global risk and the broad dollar index: both rise strongly at the height of the GFC (left panel) and the early stage of the pandemic (right panel). This co-movement is a general pattern of the data and testifies to a fundamental asymmetry in a global financial system centered around the dollar (Bruno & Shin 2015; Farhi & Gabaix 2016; Bocola & Lorenzoni 2020).[1] But the dollar's dominance is not limited to international finance, it also extends to trade (Gopinath et al. 2020). This dominance, in turn, is key for the transmission of global risk shocks. We establish this insight based on new time series evidence and a model of the world economy which features dollar dominance in both, finance and trade. For lack of a better term, we refer to it as the 'DCP^2 model'.

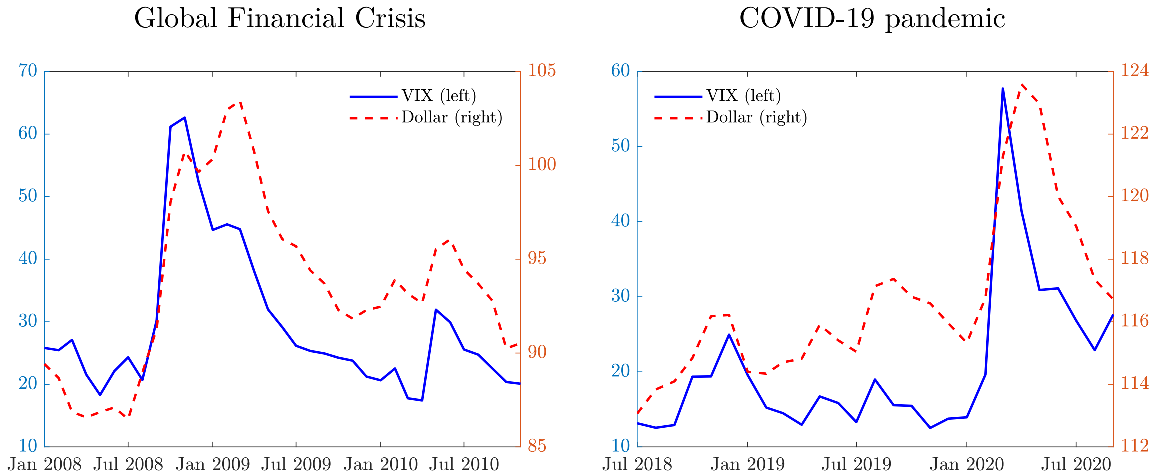
Does the dollar's dominance help the world economy in coping with global risk shocks or does it amplify their adverse impact? In this paper, we shed light on this question by exploring empirically the dollar's role in the transmission of global risk. First, we support the received wisdom with rigorous evidence: We show that structurally identified global risk shocks appreciate the dollar. Second, we find that the appreciation of the dollar amplifies the adverse effects of global risk shocks through tighter financial conditions; expenditure switching does little to stabilize economic activity outside the US.

In order to estimate the causal effects of global risk shocks, we use intra-daily surprises in the price of gold—the ultimate safe asset—as an external instrument in a Bayesian proxy vector-autoregressive model. As predicted by theory, we find that global risk shocks induce an appreciation of the dollar and other safe-haven currencies, 'flight-to-safety' as foreign holdings of US Treasury securities increase, a rise in the US Treasury premium, the dollar liquidity buffers of banks and the share of dollar-denominated debt in total international debt issuance. We also establish that global risk shocks induce a contraction in global economic activity, consistent with findings for the US (Baker et al. 2016; Basu & Bundick 2017; Bloom et al. 2018). Reflecting a 'trade channel', US net exports contract, suggesting that the dollar appreciation induces expenditure switching (Gopinath et al. 2020). And reflecting a 'financial channel', global equity prices drop, spreads increase, and cross-border bank credit contracts (Bruno & Shin 2015). These patterns conform well with the notions of a global financial cycle and an 'exorbitant duty' of the US (Miranda-Agrippino & Rey 2020, 2021; Gourinchas et al. 2012, 2017). As a distinct contribution, we show that they are *caused* by and not just correlated with variations in global risk.

To quantify the relative importance of the trade and financial channel, we then construct a counterfactual that simulates the effects of a global risk shock in the absence of dollar appreciation. We find that the contraction in real activity in the rest of the world is substantially weaker when

[1] In a regression of changes in the VIX on changes in the dollar exchange rate over the period 01/1990-12/2020 the t-value is 5.8, and 2.2 when excluding the period 7/2008-12/2009 and after 03/2020. Consistent with the findings in Lilley et al. (forthcoming), the t-value is essentially zero for the time period prior to the GFC, it is 4.3 for the post-GFC period 1/2010-12/2020, and 3.6 for the inter-crises period 1/2010-3/2020.

Figure 1: The US dollar and the VIX



Note: VIX is an index of expected stock market volatility compiled by Chicago Board of Options Exchange; dollar is the price of dollar expressed in foreign currency (in effective terms) such that an increase represents an appreciation.

dollar appreciation is absent. The contractionary effects of dollar appreciation that materialize through tighter financial conditions thus dominate expansionary effects due to expenditure switching. Indeed, without dollar appreciation the response of US net exports hardly changes, while global financial conditions tighten much less.

Finally, we present a structural two-country model for the US and the rest of the world that can match our empirical estimates for the effects of a global risk shock. The model incorporates a special role for the dollar by bringing together a dominant-currency paradigm in *both* trade and cross-border banking (Gopinath et al. 2020; Akinci & Queralto 2019). For lack of a better term, we refer to this framework as the 'DCP^2 model'. In the model, dollar dominance in trade means that US imports and a share of domestic transactions in the rest of the world are priced in dollar; the latter reflects that in the data a large share of third-country, non-US trade is invoiced in dollar (Boz et al. 2022). In turn, dollar dominance in cross-border banking means that US banks intermediate dollar liquidity to banks in the rest of the world that are subject to currency mismatches and lend to domestic borrowers. We show that the impulse responses to a global risk shock in the DCP^2 model with a standard parameter calibration match the impulse responses in the data. Moreover, we show that dollar dominance in both trade and cross-border banking are necessary for doing so.

In more detail, we estimate a Bayesian proxy structural vector-autoregressive (BPSVAR) model as proposed by Arias et al. (2018, 2021). We use monthly observations for the period 1990–2019 and in the baseline specification include the VXO, industrial production in the US and the rest of the world (RoW), the consumer price index and the excess bond premium in the US, the 1-year Treasury Bill rate as an indicator of US monetary policy, RoW policy rates, and the US dollar nominal effective exchange rate. In order to speak to the theoretical literature on the foundations of the dominant role of the dollar, we consider extended specifications which feature other economies' exchange rates, the US Treasury premium, foreign holdings of US Treasury securities, banks' dollar asset liquidity ratio, the share of dollar-denominated in total international debt securities of non-US issuers, US exports and imports, cross-border bank credit flows to non-US borrowers, the Emerging Markets Bond Index (EMBI) spread, equity prices, and the global factors in risky asset prices and capital flows of Miranda-Agrippino & Rey (2020) and Miranda-Agrippino et al. (2020).

Consistent with recent theoretical work we conceive of a global risk shock as an incident that is associated with an exogenous increase in the demand for safe and liquid assets (Maggiori 2017; Jiang et al. 2021a; Kekre & Lenel 2021). In order to identify a global risk shock we rely on an external instrument (Stock & Watson 2012; Mertens & Ravn 2013). In particular, as in Piffer & Podstawski (2018) and Ludvigson et al. (2021) we use as external instrument the change in the gold price around narrow intra-daily windows bracketing the time stamps of global risk events selected narratively originally by Bloom (2009); we document that results are similar when we instead use surprises in long-term Treasury yields and the dollar-euro exchange rate for which the intra-daily windows are narrower than for gold, as well as for monthly changes in the Geopolitical Risk Index of Caldara & Iacoviello (2022) which by construction reflect exogenous variation in risk so that we do not have to rely on narratively selected events. In order to explore a policy experiment in which the Federal Reserve (Fed) stabilizes the dollar, in an extension to our baseline analysis we use intra-daily high-frequency surprises in short-term Treasury futures around Federal Open Market Committee announcements as an external instrument to additionally identify US monetary policy shocks (Gertler & Karadi 2015; Jarociński & Karadi 2020).

The BPSVAR framework of Arias et al. (2018, 2021) is suited particularly well for our purposes relative to traditional frequentist approaches (Mertens & Ravn 2013; Lakdawala 2019). First, it makes more efficient use of the information contained in the external instruments by avoiding estimation in multiple steps; this also facilitates coherent and exact finite-sample inference, especially when the external instruments are weak (Caldara & Herbst 2019; Montiel Olea et al. 2021). Second, in a setting in which multiple structural shocks are jointly identified by multiple external instruments it allows us to avoid restrictions on the contemporaneous relationships between endogenous variables, which might seem controversial in the context of risk shocks (Angelini et al. 2019; Alessandri et al. 2020; Redl 2020; Carriero et al. 2021).

We find that a one-standard-deviation global risk shock appreciates the dollar by about 0.5%. Other safe-haven currencies such as the Japanese yen and the Swiss franc also appreciate; non-safe-haven currencies such as the euro and the British pound depreciate. The US Treasury premium rises by about 5 basis points, reflecting an increase in the relative, currency-hedged convenience yield between US and other G10 government bonds. Foreign holdings of US Treasury securities increase by up to 1%, indicating 'flight-to-safety' capital flows. US and RoW industrial production decline in a hump-shaped and highly synchronized pattern; the recessionary impact is strongest after about six months, when US and RoW industrial production fall by up to 0.4%. Monetary policy loosens, with rates declining by up to 10 basis points in the US and the RoW. US exports and imports contract by about 0.6% and 0.2% on impact, respectively; consistent with dominant-currency pricing the maximum contraction in US exports occurs on impact, while it is delayed for US imports (Gopinath et al. 2020). Global financial conditions in terms of risky asset prices and capital flows tighten. Specifically, cross-border bank credit to non-US borrowers contracts by up to 1%, RoW equity prices fall by 1.6%, and the EMBI spread rises by up to 25 basis points.

We then construct counterfactuals in which the dollar does not respond in order to assess its contribution to the transmission of a global risk shock to the RoW. The counterfactual is based on a 'minimum relative entropy' (MRE) approach previously used in the context of forecasting (Robertson et al. 2005; Cogley et al. 2005; Giacomini & Ragusa 2014). The original idea is to improve forecasts by incorporating restrictions implied by economic theory in the least 'disruptive' way. We apply the MRE approach to construct impulse responses for a counterfactual in which the dollar is unresponsive to a global risk shock while the impulse responses of the remaining variables are minimally different compared to the baseline in an information-theoretic sense. We find that in this counterfactual the contractionary effect of a global risk shock on RoW industrial production is roughly halved compared to the baseline. This implies that the contractionary effects via the financial channel dominate the expansionary effects due to expenditure switching that operate via the trade channel. Indeed, while US net exports only fall somewhat less in the counterfactual, global financial conditions tighten much less.

In extensions to this counterfactual we illustrate that the dollar's role in the transmission of global risk shocks is special and not shared by other safe-haven currencies: in the counterfactual the global risk shock is associated with a weaker drop in especially *dollar*-denominated compared to non-dollar cross-border credit, consistent with the findings of Ivashina et al. (2015). Also, suppressing appreciation of other safe-haven currencies instead of the dollar in the counterfactual is inconsequential for the effects of global risk shocks.

Finally, we consider an alternative approach to construct a counterfactual by carrying out a policy experiment. The MRE approach is purely data driven and thereby agnostic as to why dollar appreciation in response to global risk shocks is absent in the counterfactual. In order to construct an alternative counterfactual we assume that US monetary policy deviates from its past behaviour and stabilizes the dollar in the face of a global risk shock. The experiment is motivated by the unprecedented emergency liquidity the Fed provided to many economies through various facilities during the COVID-19 pandemic. It is widely believed that these measures were crucial for preventing a global financial crisis (see Cetorelli et al. 2020). Theoretically, Fed swap lines can be conceived as increasing the supply of safe assets by crediting RoW central banks with dollar reserves, which reduces the convenience yield and thereby depreciates—or dampens appreciation pressures on—the dollar (Jiang et al. 2021a). We implement this policy experiment by specifying a sequence of US monetary policy shocks which offsets the effect of a global risk shock on the dollar (e.g. Bachmann & Sims 2012; Epstein et al. 2019). We find that—similarly to the MRE counterfactual—by adopting a more accommodative stance that prevents dollar appreciation US monetary policy would mitigate substantially the contractionary effects of a global risk shock in the RoW.

We then show that that the impulse responses to a global risk shock in the two-country DCP[2] model for the world economy with a standard parameter calibration match the impulse responses in the data. Moreover, we show that dollar dominance in both trade and cross-border banking are necessary for doing so, especially for the contractionary effects in the rest of the world. In particular, on the one hand, as in Gopinath et al. (2020) dollar dominance in US trade mutes the response of US import prices following dollar appreciation, and therefore eliminates traditional Mundellian expenditure switching that would dampen the contractionary effects of a global risk shock in the rest of the world. And as in Mukhin (2022) and Zhang (2022) dollar dominance in a subset of domestic rest-of-the-world transactions entails that consumer prices increase strongly following dollar appreciation, which induces rest-of-the-world monetary policy to tighten, amplifying the contractionary effects of a global risk shock. On the other hand, as in Bruno & Shin (2015) and Akinci & Queralto (2019) dollar dominance in cross-border banking entails that dollar appreciation reduces the net worth of rest-of-the-world banks as they are subject to currency mismatches, which leads to a tightening financing conditions and thereby also amplifies the contractionary effects of a global risk shock.

Related literature. Our paper first speaks to recent theoretical work on the special role of the dollar and US assets in the international monetary system (Farhi & Gabaix 2016; Maggiori 2017; Jiang et al. 2021a; Kekre & Lenel 2021; Bianchi et al. 2021). Our contribution is to assess the empirical relevance of these mechanisms spelled out theoretically in a unified VAR framework. More generally, our analysis also informs empirically the theoretical literature on the role of exchange rates for the cross-border transmission of shocks through financial channels (Banerjee et al. 2016; Aoki et al. 2018; Akinci & Queralto 2019). Second, our paper is related to empirical work that studies the role of the dollar as a global risk factor (Lustig et al. 2014; Verdelhan 2018), the predictive power of convenience yields (Engel & Wu 2018; Jiang et al. 2021b) and global risk (Lilley et al. forthcoming; Hassan et al. 2021) for the dollar, as well as the relationship between global risk, deviations from covered interest parity, the dollar and cross-border credit (Avdjiev, Du, et al. 2019; Erik et al. 2020). We complement these lines of research by moving from forecasting and reduced-form regressions to isolating the effect of exogenous innovations to global risk on the dollar, the Treasury premium, and cross-border bank credit. Third, our paper contributes to empirical work on the role financial channels play in the global transmission of risk and uncertainty shocks (Carriere-Swallow & Cespedes 2013; Liu et al. 2017; Cesa-Bianchi et al. 2018; Epstein et al. 2019; Shousha 2019; Bhattarai et al. 2020). Relative to the existing work, we zoom in on and quantify the role of the dollar within the broader class of financial channels for the transmission of global risk shocks. Moreover, we consider exogenous variation in global risk, and in doing so we avoid recursiveness assumptions by using external instruments in a flexible BPSVAR model. Our findings on the role of the dollar for financial spillovers complement existing evidence based on micro data (Shim et al. 2021; Banerjee et al. 2020; Avdjiev, Bruno, et al. 2019; Bruno & Shin 2021; Meisenzahl et al. 2019; Niepmann & Schmidt-Eisenlohr forthcoming). Relative to this work, our analysis allows us to assess the net contribution of dollar appreciation—contrasting trade and financial channels—on the effects of global risk shocks on the aggregate economy. Finally, in the DCP^2 model we bring together dominant-currency paradigms in trade (Gopinath et al. 2020) and cross-border banking (Akinci & Queralto 2019). We additionally introduce dollar dominance in domestic, intra-rest-of-the-world transactions to reflect that in the data a large share of third-country, non-US trade is invoiced in dollar (Boz et al. 2022).

The rest of the paper is organized as follows. Section 2 lays out the BPSVAR framework and describes our empirical specification. Section 3 presents our results for the effects of global risk shocks on the dollar and the world economy. Section 4 zooms in on the role of the dollar based on a counterfactual, and Section 4.3 carries out the policy experiment. Finally, Section 6 concludes.

2 Empirical strategy

We first briefly outline the general BPSVAR model framework put forth by Arias et al. (2021) and then discuss our specification and identification assumptions.

2.1 General framework

Using the notation of Rubio-Ramirez et al. (2010), consider the structural VAR model

$$\boldsymbol{y}_t'\boldsymbol{A}_0 = \boldsymbol{y}_{t-1}'\boldsymbol{A}_1 + \boldsymbol{\epsilon}_t', \tag{1}$$

where $\boldsymbol{y}_t$ is an $n \times 1$ vector of endogenous variables and $\boldsymbol{\epsilon}_t$ an $n \times 1$ vector of structural shocks. In our empirical specification below we include additional lags and deterministic terms, but omit them here for simplicity.

To achieve identification the BPSVAR framework exploits a $k \times 1$ vector of observed proxy variables—or, in alternative jargon, external instruments—$\boldsymbol{m}_t$. The proxy variables are assumed to be correlated with the k unobserved structural shocks of interest $\boldsymbol{\epsilon}_t^*$ (relevance condition) and orthogonal to the remaining unobserved structural shocks $\boldsymbol{\epsilon}_t^o$ (exogeneity condition). Formally, the identifying assumptions are

$$E[\boldsymbol{m}_t\boldsymbol{\epsilon}_t^{*\prime}] = \underset{(k\times k)}{\boldsymbol{V}}, \tag{2a}$$

$$E[\boldsymbol{m}_t\boldsymbol{\epsilon}_t^{o\prime}] = \underset{(n-k\times k)}{\boldsymbol{0}}. \tag{2b}$$

These identifying assumptions are operationalized by augmenting the model in Equation (1) with equations for the k proxy variables $\boldsymbol{m}_t$ so that

$$\tilde{\boldsymbol{y}}_t'\tilde{\boldsymbol{A}}_0 = \tilde{\boldsymbol{y}}_{t-1}'\tilde{\boldsymbol{A}}_1 + \tilde{\boldsymbol{\epsilon}}_t', \tag{3}$$

where $\tilde{\boldsymbol{y}}_t' \equiv (\boldsymbol{y}_t', \boldsymbol{m}_t')$, $\tilde{\boldsymbol{\epsilon}} \equiv (\boldsymbol{\epsilon}_t', \boldsymbol{v}_t')' \sim N(\boldsymbol{0}, \boldsymbol{I}_{n+k})$, $\boldsymbol{v}_t$ denotes measurement error in the proxy variables, and $\tilde{\boldsymbol{A}}_\ell$ coefficient matrices of dimension $\tilde{n} \times \tilde{n}$, $\tilde{n} = n + k$. The latter satisfy

$$\tilde{\boldsymbol{A}}_\ell = \begin{pmatrix} \underset{(n \times n)}{\boldsymbol{A}_\ell} & \underset{(n \times k)}{\boldsymbol{\Gamma}_{\ell,1}} \\ \underset{(k \times n)}{\boldsymbol{0}} & \underset{(k \times k)}{\boldsymbol{\Gamma}_{\ell,2}} \end{pmatrix}, \qquad \ell = 0, 1, \tag{4}$$

to ensure that augmenting the model in Equation (1) with equations for the proxy variables does not alter the dynamics of the endogenous variables.

As

$$\tilde{\boldsymbol{A}}_0^{-1} = \begin{pmatrix} \boldsymbol{A}_0^{-1} & -\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1} \\ 0 & \boldsymbol{\Gamma}_{0,2}^{-1} \end{pmatrix}, \tag{5}$$

in the reduced form of the model given by

$$\tilde{\boldsymbol{y}}_t' = \tilde{\boldsymbol{y}}_{t-1}'\tilde{\boldsymbol{A}}_1\tilde{\boldsymbol{A}}_0^{-1} + \tilde{\boldsymbol{\epsilon}}_t{}'\tilde{\boldsymbol{A}}_0^{-1}, \tag{6}$$

the last k equations read as

$$\boldsymbol{m}_t' = \tilde{\boldsymbol{y}}_{t-1}'\tilde{\boldsymbol{A}}_1 \begin{pmatrix} -\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1} \\ \boldsymbol{\Gamma}_{0,2}^{-1} \end{pmatrix} - \boldsymbol{\epsilon}_t'\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1} + \boldsymbol{v}_t'\boldsymbol{\Gamma}_{0,2}^{-1}. \tag{7}$$

Ordering the structural shocks as $\boldsymbol{\epsilon}_t = (\boldsymbol{\epsilon}_t^{o\prime}, \boldsymbol{\epsilon}_t^{*\prime})'$ we have

$$E\left[\boldsymbol{\epsilon}_t\boldsymbol{m}_t'\right] = -\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1} = \begin{pmatrix} \underset{((n-k)\times k)}{\boldsymbol{0}} \\ \underset{(k \times k)}{\boldsymbol{V}} \end{pmatrix}. \tag{8}$$

The first equality is implied by Equation (7) and the assumption that the structural shocks $\boldsymbol{\epsilon}_t$ are orthogonal to $\boldsymbol{y}_{t-1}$ and $\boldsymbol{v}_t$. The second equality is implied by the exogeneity and relevance conditions in Equations (2a) and (2b).

Equation (8) shows that the identifying assumptions of the BPSVAR model imply restrictions on the last k columns of the contemporaneous structural impact coefficients in $\tilde{\boldsymbol{A}}_0^{-1}$ in Equation (5). In particular, if the exogeneity condition in Equation (2b) holds, the first $n-k$ rows of the upper right-hand side sub-matrix $\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1}$ of $\tilde{\boldsymbol{A}}_0^{-1}$ are zero. From the reduced form in Equation (6) it can be seen that this implies that the first $n-k$ structural shocks do not impact the proxy variables contemporaneously. In turn, if the relevance condition in Equation (2a) holds, the last k rows of $\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1}$ are different from zero. From the reduced form in Equation (6) it can be seen that this implies that the last k structural shocks impact the proxy variables contemporaneously. Arias et al. (2021) develop an algorithm that estimates $\boldsymbol{A}_\ell$ and $\boldsymbol{\Gamma}_{0,\ell}$ while the restrictions on $\tilde{\boldsymbol{A}}_0^{-1}$ implied by Equations (2a) and (2b) are satisfied, and hence the estimation identifies the structural shocks of interest in $\boldsymbol{\epsilon}_t^*$.

In Appendix A we discuss in detail the numerous advantages of the BPSVAR framework of Arias et al. (2021) over the traditional frequentist external instruments SVAR framework. In short: first, the BPSVAR framework allows us to refrain from imposing potentially contentious recursiveness assumptions between the endogenous variables when we point-identify multiple structural shocks with multiple proxy variables.[2] Instead, we can impose the additional necessary identification assumptions on the relationship between the proxy variables and the structural shocks of interest, for example that the proxy variable for the global risk shock is not affected by, say, the monetary policy shock.[3] Second, by avoiding a two-step approach the single-step estimation of the BPSVAR model is more efficient and facilitates coherent inference; in fact, the Bayesian set-up allows exact finite sample inference, and does not require an explicit theory to accommodate weak instruments. Third, the BPSVAR framework is relatively flexible in that Equation (7) allows the proxy variables to be serially correlated and to be affected by lags of the endogenous variables as well as by measurement error.

2.2 Empirical specification

Our point of departure is the US VAR model of Gertler & Karadi (2015) which includes as endogenous variables in $\boldsymbol{y}_t$ the logarithms of US industrial production and consumer prices, the excess bond premium of Gilchrist & Zakrajsek (2012), and the 1-year Treasury Bill rate as monetary policy indicator. We augment $\boldsymbol{y}_t$ with the VXO as a measure of global risk (Londono & Wilson 2018), the logarithm of an index of non-US, RoW industrial production, a weighted average of advanced economies' (AEs) policy rates, and the logarithm of the US dollar nominal effective exchange rate (NEER).[4] We use monthly data for the time period from February 1990 to December 2019. We use flat priors for the VAR parameters. Below we consider a robustness check for a larger VAR model that includes many additional variables estimated with informative Minnesota-type priors and optimal hyperpriors/prior tightness as suggested by Giannone et al. (2015). Data descriptions are provided in Table D.1.

2.3 Identification

We think of global risk shocks as incidents that are associated with an increase in the demand for safe and/or liquid assets. This notion is well supported by the data and rationalized by theory. In the theory, it has been shown that demand for safe and liquid assets may rise during times of elevated risk due to differences in economies' resilience to rare disasters (Farhi & Gabaix 2016), risk-bearing capacity (Maggiori 2017), or frictions in interbank markets (Bianchi et al. 2021). In the

[2]For example, it has been documented that US monetary policy shocks have large contemporaneous effects on global risk (Bekaert et al. 2013; Rey 2016; Miranda-Agrippino & Rey 2020). In fact, the results we present below based on identification assumptions that do not impose restrictions on the contemporaneous relationship between endogenous variables confirm this finding.

[3]Other approaches to overcome the limitations of recursive identification in the context of global risk shocks exploit heteroskedasticity across regimes or over time (Angelini et al. 2019; Carriero et al. 2021), use a bridge-proxy SVAR model that imposes recursiveness only at a higher frequency (Alessandri et al. 2020), or rely on narrative restrictions (Redl 2020).

[4]We use AE instead of RoW policy rates as the latter exhibit spikes reflecting periods of hyperinflation in some EMEs. We discuss an extension below in which we include AE and EME industrial production, consumer prices and policy rates separately. We consider the VXO instead of the VIX because the latter is not available before 1994.

data, it has been documented that investors impute a non-pecuniary convenience yield to especially US Treasury securities (Krishnamurthy & Vissing-Jorgensen 2012; Jiang et al. 2021b), which rises during episodes of elevated global risk amidst a 'flight-to-safety'.

2.3.1 Proxy variables

Our proxy variable for global risk shocks is constructed on the basis of intra-daily data in the spirit of work on the high-frequency identification of monetary policy shocks (see Gertler & Karadi 2015, and references therein). Specifically, building on the work of Bloom (2009) as well as Piffer & Podstawski (2018) we consider the intra-daily changes in the price of gold—the ultimate safe asset—on narratively selected events as proxy variable for global risk shocks. Piffer & Podstawski (2018) first extend the list of exogenous risk events compiled by Bloom (2009). Second, as a quantitative measure of the size of the unobserved shock, they calculate the change in the price of gold between the last auction before and the first auction after the news about the event was released to markets.[5] In the baseline, we consider the events labelled as 'global' and 'US' by Piffer & Podstawski (2018); we include those labelled as 'European' and 'other' risk events in a robustness check. We aggregate the daily gold price surprises to monthly frequency as in Gertler & Karadi (2015). In particular, we first cumulative the daily surprises, then, second, take average within months, and, third, take the monthly first difference.[6]

It is important to emphasize that what is critical for the exogeneity condition $E[p_t^{\epsilon,r}\boldsymbol{\epsilon}_t^o] = \mathbf{0}$ in Equation (2b) to be satisfied in our context is that the gold price surprises around the intra-daily windows were driven *systematically* across the narratively selected events only by the global risk shock. For this, the selection of events and the narrowness of the intra-daily windows around the corresponding time stamps rather than the specific asset price for which the surprises are calculated are crucial. Below we explore robustness checks in which we consider alternative assets for which we can calculate surprises around narrower intra-daily windows than for gold, namely long-term Treasury securities and the US dollar-euro exchange rate. Moreover, we consider a robustness check which does not rely on narratively selected events at all, using as proxy variable monthly changes in the Geopolitical Risk Index of Caldara & Iacoviello (2022) that by construction reflects exogenous variation in risk. And finally, we consider a robustness check in which we relax the exogeneity condition allowing also non-global risk structural shocks to have affected the gold price surprises systematically across events, and instead only require that their effect was smaller than for global risk shocks so that $|E[p_t^{\epsilon,r}\epsilon_t^r]| > |E[p_t^{\epsilon,r}\epsilon_t^\ell]|$ for $\ell \neq r$.

We additionally identify a US monetary policy shock in order to carry out a policy experiment after our main analysis. We follow Gertler & Karadi (2015) and use the change in the 3-month

[5] The analysis of Piffer & Podstawski (2018) covers the time period until 2015; we use the update of Bobasu et al. (2021) that spans until 2019. In their analysis, Piffer & Podstawski (2018) refer to 'uncertainty' shocks rather than 'risk' shocks. Our use of the term 'risk' is meant to be broad so as to encompass both uncertainty and risk aversion. In robustness checks below we explore the effect of a global risk shock on distinct measures of the risk and uncertainty components in the VXO (Bekaert et al. forthcoming).

[6] Note that while this may induce serial correlation in the interest rate surprises, this is explicitly allowed for in the BPSVAR framework (see Equation (7)).

Federal Funds rate futures in a narrow time window around FOMC announcements as a proxy variable. We purge these surprises from central bank information shocks using the 'poor-man's' approach of Jarociński & Karadi (2020): When the interest rate surprise has the same sign as the equity price surprise, it is classified as central bank information shock; when the interest rate and the equity price surprises have the opposite sign, it is classified as a 'pure' monetary policy shock.

2.3.2 Identifying assumptions

Our identifying assumptions are given by

$$E[\boldsymbol{\epsilon}_t^* \boldsymbol{m}_t'] = \begin{pmatrix} E[p_t^{\epsilon,r} \epsilon_t^r] & E[p_t^{\epsilon,mp} \epsilon_t^r] \\ E[p_t^{\epsilon,r} \epsilon_t^{mp}] & E[p_t^{\epsilon,mp} \epsilon_t^{mp}] \end{pmatrix} = \boldsymbol{V}, \tag{9a}$$

$$E[\boldsymbol{\epsilon}_t^o \boldsymbol{m}_t'] = \begin{pmatrix} E[p_t^{\epsilon,r} \boldsymbol{\epsilon}_t^o] & E[p_t^{\epsilon,mp} \boldsymbol{\epsilon}_t^o] \end{pmatrix} = \boldsymbol{0}, \tag{9b}$$

where $\boldsymbol{\epsilon}_t^* \equiv (\epsilon_t^r, \epsilon_t^{mp})'$, ϵ_t^r denotes the global risk shock and ϵ_t^{mp} the US monetary policy shock, and $\boldsymbol{m}_t \equiv (p_t^{\epsilon,r}, p_t^{\epsilon,mp})'$ contains the corresponding proxy variables.

First, in the relevance condition in Equation (9a) we assume that global risk shocks drive gold price surprises on the narratively selected dates, $E[p_t^{\epsilon,r} \epsilon_t^r] \neq 0$. Intuitively, increases in precautionary savings push up the price of gold in response to risk shocks (Baur & McDermott 2010). Piffer & Podstawski (2018) provide evidence that gold price surprises are relevant instruments for risk shocks based on F-tests and Granger-causality tests with the VXO and the macroeconomic uncertainty measure constructed in Jurado et al. (2015). Ludvigson et al. (2021) also use gold price changes as a proxy variable for global risk shocks; Engel & Wu (2018) use the gold price as a proxy for risk. Regarding the exogeneity condition $E[p_t^{\epsilon,r} \boldsymbol{\epsilon}_t^o] = 0$ in Equation (9b), Piffer & Podstawski (2018) document that the intra-daily gold price surprises on the narratively selected dates are not systematically correlated with a range of measures of non-risk shocks. This is consistent with the notion that the only shock that systematically occurred in the intra-daily windows across the narratively selected dates is the global risk shock.[7]

Second, we assume that US monetary policy shocks drive the Federal Funds futures surprises in narrow windows around FOMC announcements in the relevance condition in Equation (9a), $E[p_t^{\epsilon,mp} \epsilon_t^{mp}] \neq 0$ (Gertler & Karadi 2015; Caldara & Herbst 2019; Jarociński & Karadi 2020). Regarding the exogeneity condition $E[p_t^{\epsilon,mp} \boldsymbol{\epsilon}_t^o] = 0$ in Equation (9b), it seems plausible that around these narrow windows—especially after cleansing from central bank information shocks—Federal Funds futures surprises are only driven by monetary policy shocks.

Note that when multiple proxy variables are used to identify multiple structural shocks the relevance and exogeneity conditions are not sufficient for point identification (see Appendix A and Mertens & Ravn 2013). In this case, additional restrictions need to be imposed on $\boldsymbol{V}$ in Equation

[7] Note that the events considered by Bloom (2009), Piffer & Podstawski (2018) as well as Bobasu et al. (2021) are very diverse. Therefore, even if on each and every event not only a global risk shock materialized, the non-risk shock is likely to have been of a different nature across events. For example, while the collapse of AIG may have in part been a financial, the 9/11 attack was arguably not.

(9a). A natural idea is to impose that $\boldsymbol{V}$ is a diagonal matrix, implying that Federal Funds futures surprises around FOMC announcements are not driven by global risk shocks and that the narratively selected gold price surprises are not driven by US monetary policy shocks. Unfortunately, this implies over-identifying restrictions, which cannot be implemented by the estimation algorithm of Arias et al. (2021). We therefore impose a *weaker* set of additional restrictions, namely only that Federal Funds futures surprises on FOMC announcement days are not driven by global risk shocks, $E[p_t^{\epsilon,mp}\epsilon_t^r] = 0$.[8] This assumption is implicitly maintained in the literature on the effects of monetary policy shocks (Gertler & Karadi 2015; Caldara & Herbst 2019; Jarociński & Karadi 2020). The assumption is also mild because we purge the Federal Funds futures surprises of central bank information shocks.

Finally, for consistency we follow Caldara & Herbst (2019) as well as Arias et al. (2021) and impose a 'relevance threshold' to express a prior belief that the proxy variables are relevant instruments. In particular, we require that at least a share $\gamma = 0.1$ of the variance of the proxy variables is accounted for by the global risk and US monetary policy shocks, respectively; this is weaker than the relevance threshold of $\gamma = 0.2$ used by Arias et al. (2021), and—although not straightforward to compare conceptually—lies below the 'high-relevance' prior of Caldara & Herbst (2019). Below we consider a robustness check in which we omit the relevance threshold.

3 Results

We first present results for the effects of global risk shocks in the baseline specification. Then, we present results for additional variables to speak to the theoretical literature on the dominant role of the dollar in the international monetary system and flesh out the transmission to the RoW through the trade and financial channel. Finally, we discuss several robustness checks.

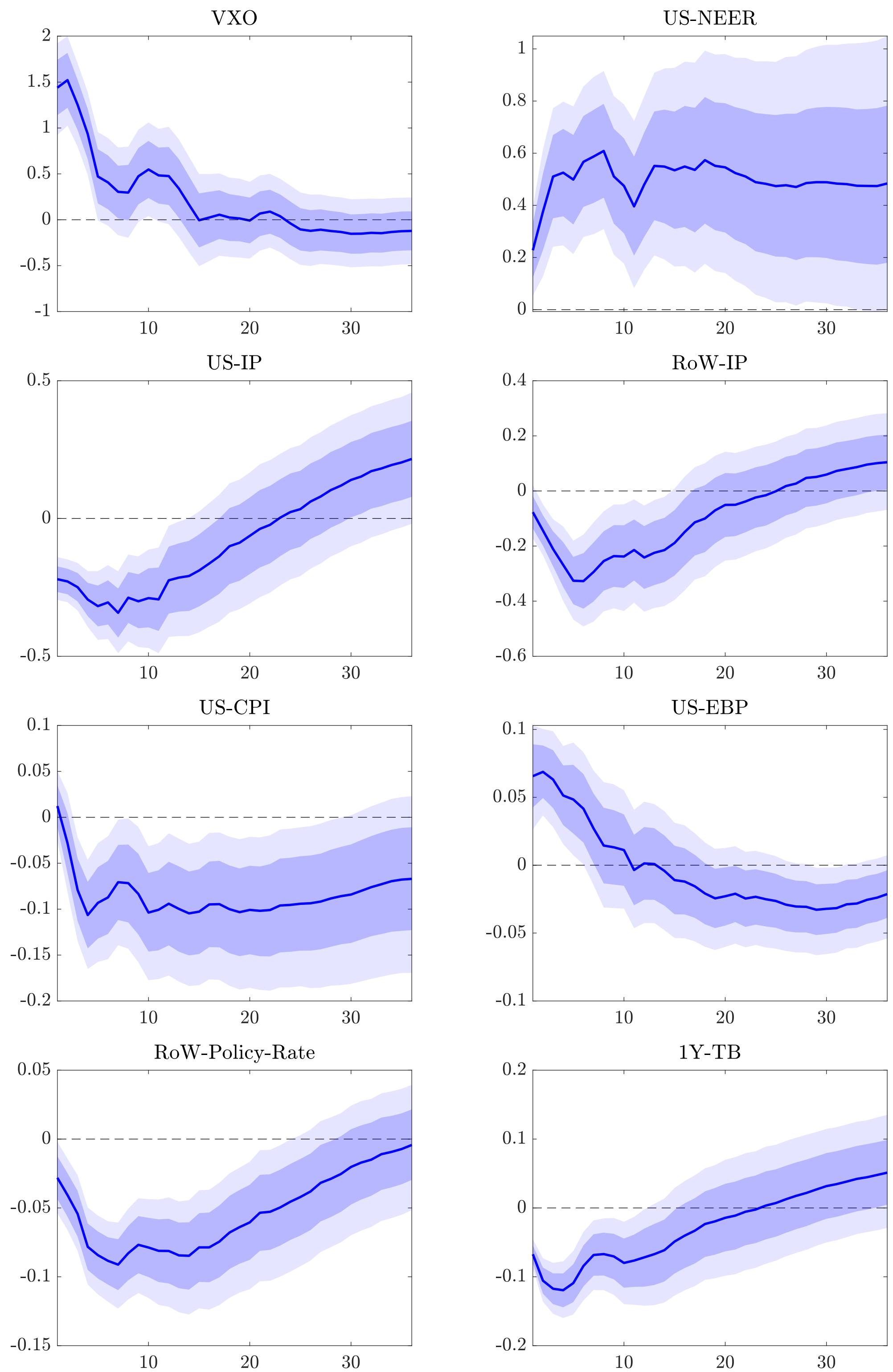
3.1 The effect of global risk shocks

Figure 2 presents the effects of a one-standard deviation global risk shock on the variables in the baseline specification. The global risk shock causes an increase in the VXO and a persistent appreciation of the dollar. This result implies that the positive co-movement between global risk and the dollar in the data can at least to some extent be accounted for by global risk shocks. That the dollar appreciates in response to a global risk shock is consistent with the predictions of the models of Farhi & Gabaix (2016), Jiang et al. (2021a), Bianchi et al. (2021), and Kekre & Lenel (2021); at the same time, it underscores the 'reserve currency *paradox*' in the model in Maggiori (2017).

US and RoW industrial production contract in tandem. The contraction is more immediate and somewhat larger in the US than in the RoW, consistent with the notion that the reserve asset issuer bears the brunt of 'safety traps' (Kekre & Lenel 2021; Caballero et al. 2020). The finding of a dollar appreciation and a global contraction is consistent with predictions from Jiang et al. (2021a) and

[8] Note that when two proxy variables are used to identify two structural shocks, a single additional zero restriction on $\boldsymbol{V}$ in Equation (9a) is sufficient for point-identification (Giacomini et al. forthcoming). This is appealing also because under set-identification credible sets are wider and results may depend on the choice of the prior distribution for the construction of the rotation matrices in the estimation (Baumeister & Hamilton 2015).

Figure 2: Impulse responses to a global risk shock



Note: Horizontal axis measures time in months, vertical axis measures deviation from pre-shock level; size of shock is one standard deviation; blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets. VXO is measured in levels, the dollar NEER, US and RoW industrial production, US consumer prices in logs, and the excess bond premium, the RoW policy as well as the US 1-year Treasury Bill rates in percent.

Kekre & Lenel (2021). For example, in the model of Jiang et al. (2021a) a demand for safe dollar assets gives rise to a non-pecuniary convenience yield and global financial accelerator effects. Finally, US consumer prices fall and the US excess bond premium rises persistently. US and RoW monetary policy are loosened.

We briefly mention results for some extensions for which we provide details in the Appendix. First, we distinguish effects for AEs and EMEs (Figure B.1). Here the evidence points to 'fear-of-floating' (Calvo & Reinhart 2002) in EMEs, consistent with the notion that local monetary policy attempts to limit exchange rate depreciation (Ahmed et al. 2021; Corsetti et al. 2021) and that this turns out to be self defeating as it induces currency risk premia to rise (Kalemli-Özcan 2019). Second, we use sign restrictions to additionally identify global demand shocks because they have been found to exhibit similar patterns as global risk shocks (Leduc & Liu 2016). While the effects on several variables are indeed qualitatively similar, the dollar does not appreciate and the gold price—when included as additional endogenous variable—does not increase in response to demand shocks (Figure C.2).[9,10] And finally, both a more accurate measure of risk aversion ('price of risk') and uncertainty ('quantity of risk') constructed by Bekaert et al. (forthcoming) increase in response to a global risk shock (Figure C.3).

3.2 Responses of additional variables

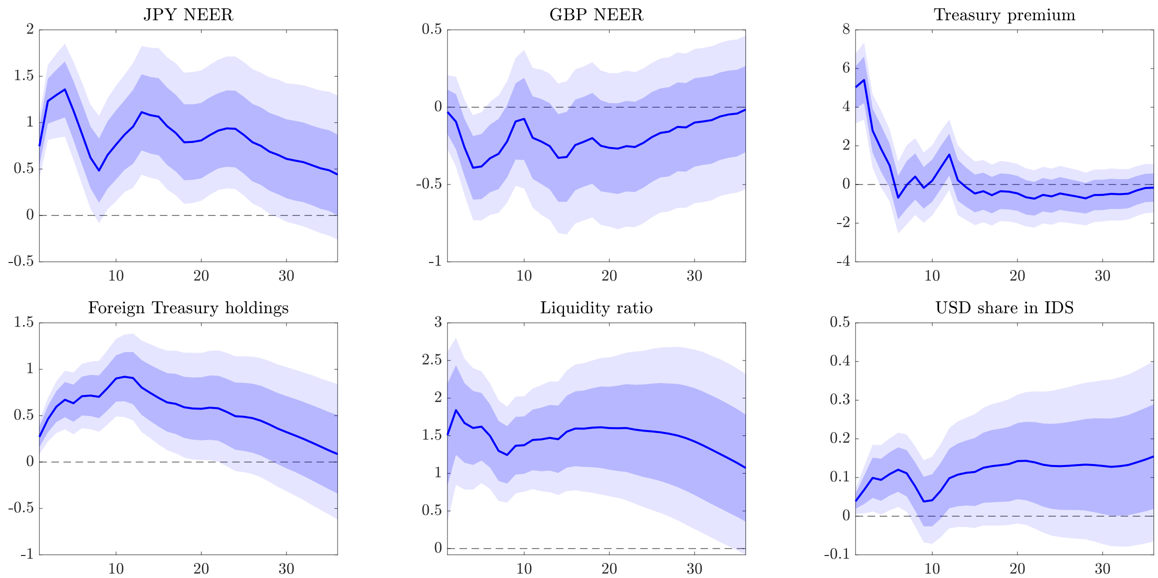
We next pull together several lines of theoretical work that are concerned with the implications of global risk for the world economy and the role of the dollar in its international transmission and assess the empirical relevance of the predicted mechanisms. Specifically, we present the effects of global risk shocks on several additional variables to explore the empirical relevance of the mechanisms articulated in the theoretical models of Farhi & Gabaix (2016), Maggiori (2017), Jiang et al. (2021a), Bianchi et al. (2021), and Kekre & Lenel (2021). To do so, we modify the baseline specification by including one additional variable at a time. In this way we keep the dimensionality of $\boldsymbol{y}_t$ limited; we consider a large VAR model in which we include all additional variables simultaneously in $\boldsymbol{y}_t$ as a robustness check below.

First, if the appreciation of the dollar in Figure 2 is indeed driven by 'flight-to-safety' against the backdrop of a global risk shock, we expect other safe-haven currencies to appreciate as well (Farhi & Gabaix 2016). Indeed, Figure 3 documents that the Japanese yen—typically considered a safe-haven currency (Ranaldo & Söderlind 2010; De Bock & de Carvalho Filho 2015)—also appreciates in response to a global risk shock. In contrast, the British pound—typically not considered a safe-haven currency—depreciates; results for the Swiss franc and the euro as alternative safe-haven and non-safe-haven currencies are very similar (Figure C.4). These patterns are consistent with the

[9] Note that following Enders et al. (2011) we leave the sign of the dollar response unrestricted, and impose that US and RoW industrial production, US consumer prices as well as US and RoW interest rates fall, and that the excess bond premium rises. Note that in case of the global risk shock our identifying assumptions only impose that the gold price increases on the *day* of a risk event and not during the entire month.

[10] The effects of global risk shocks we estimate are qualitatively different from what has been obtained for news shocks. In particular, Piffer & Podstawski (2018) find that while contractionary risk shocks are followed by a loosening of US monetary policy and a decline in US inflation, the opposite happens in response to adverse news shocks.

Figure 3: Impulse responses of additional variables



Note: See notes to Figure 2. Responses are obtained from estimating the baseline BPSVAR model with the vector $\boldsymbol{y}_t$ augmented with one additional variable at a time. Because data on the liquidity ratio is only available from 2001 we use informative priors and optimal hyperpriors/prior tightness as suggested by Giannone et al. (2015).

reduced-form results in Lilley et al. (forthcoming), who document a co-movement between bilateral dollar exchange rates and global risk measures, except for the Swiss franc and the Japanese yen. It is also consistent with the reduced-form results in Hassan et al. (2021), who find that besides the dollar the yen also correlates positively with their novel measure of global risk constructed on the basis of a textual analysis of quarterly earnings calls of thousands of publicly listed firms worldwide.

Second, the model of Jiang et al. (2021a) predicts that dollar appreciation is induced by an increase in US Treasury securities' currency-hedged relative convenience yield, which is in turn triggered by a drop in the supply or increase in the demand for safe and liquid dollar assets during a global crisis. Figure 3 shows that the Treasury premium of Du et al. (2018)—or, inversely defined, the Treasury basis of Jiang et al. (2021b)—over other G10 countries' sovereign bonds indeed increases sharply in response to a global risk shock.

Third, in Jiang et al. (2021a) a global risk shock appreciates the dollar through an increase in the convenience yield of US Treasury securities driven by a rise in the demand for safe and liquid US assets. The model of Maggiori (2017) also highlights 'flight-to-safety' capital flows by non-US financial intermediaries in global crises. Consistent with this prediction, Figure 3 shows that foreign holdings of US Treasury securities increase in response to the global risk shock. This is also in line with empirical work studying capital flows during risk-off periods (Habib & Stracca 2015) and specific prominent events such as the GFC (Vissing-Jorgensen 2021).[11] Interestingly, that

[11]Note that although we present it as a motivation in Figure 1, the sample period for our empirical analysis does not include the COVID-19 pandemic. This is worthwhile to point out, because as discussed by Vissing-Jorgensen (2021)

the Treasury premium rises well ahead of the foreign holdings of Treasury securities is consistent with the pattern documented by Krishnamurthy & Lustig (2019, pp. 458): "purchases of Treasuries on average tend to follow a widening of the Treasury basis, as Treasuries become more expensive relative to foreign bonds. Foreign investors buy Treasuries when they are expensive."

Fourth, in the model of Bianchi et al. (2021) banks hold dollar assets to insure against liquidity risk. When dollar funding becomes more volatile in times of elevated risk, banks raise the ratio of safe and liquid dollar assets to liabilities. This increases global demand for dollar assets as well as the convenience yield and thereby appreciates the dollar. Bianchi et al. (2021) provide evidence for this predicted correlation in regressions of the dollar on measures of banks' liquidity ratio. Figure 3 documents that global risk shocks indeed induce a positive correlation between dollar appreciation and the liquidity ratio.

Finally, the models of Jiang et al. (2021a) and Liao (2020) predict that a rise in the dollar convenience yield or in the currency-hedged corporate basis analogous to the US Treasury premium incentivizes firms to tilt the denomination of their bond issuance towards dollar. Indeed, Caramichael et al. (2021) document that for global non-US firms that issue bonds in multiple currencies cheaper relative borrowing costs in dollar correlate with a higher dollar share in their total corporate bond issuance. Figure 3 documents that the data are consistent with this prediction also from a causal perspective: the share of dollar-denominated in total international debt securities rises.

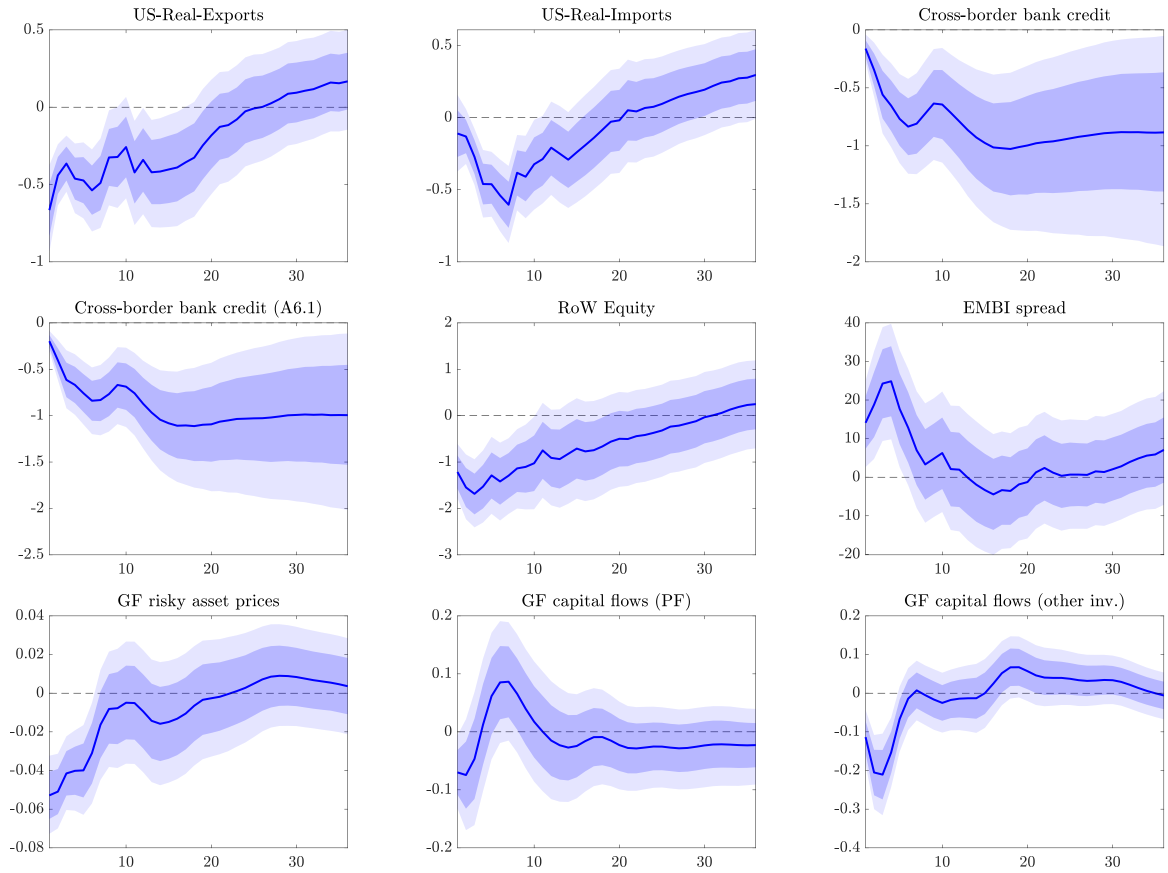
3.3 Trade and financial channels in the transmission of global risk shocks

We next explore the transmission of global risk shocks to the RoW through the trade channel and the financial channel (Bruno & Shin 2015; Jiang et al. 2021a; Obstfeld & Rogoff 1996; Gopinath et al. 2020). Figure 4 documents that global risk shocks cause a drop in both US real exports and imports. Consistent with the notion of dominant-currency pricing the decline of exports is more immediate and stronger than for imports (Gopinath et al. 2020): when both US import and export prices are sticky in dollar, dollar appreciation induces expenditure switching only away from US exports but not towards imports; the weaker and delayed decline of US imports tracks the hump-shaped contraction of economic activity in the US shown in Figure 2.

Figure 4 also presents the effects of global risk shocks on variables reflecting global financial conditions. Cross-border bank credit to non-US borrowers drops sharply and persistently in response to the global risk shock, whether measured in terms of cross-border liabilities (third panel in the

the early stage of the COVID-19 pandemic in March 2020 was marked by a 'dash for cash' in which domestic and foreign investors liquidated even US Treasury holdings.

Figure 4: Impulse responses of trade and financial variables to a global risk shock



Note: See notes to Figure 2 and Figure 3. The top-right panel depicts the response of cross-border bank credit reported in the BIS Locational Banking Statistics Table A7 based on the nationality principle (calculated as "External liabilities to all sectors of all reporting banks" less "External liabilities to all sectors of banks owned by US nationals"). The left panel in the second row shows the response cross-border bank credit taken reported in the BIS Locational Banking Statistics Table A6.1 based on the residency principle (calculated as "Banks' external claims on all sectors in all countries" less "Banks' external claims on all sectors in the US").

first row) or cross-border claims (first panel in the second row) of globally active banks.[12,13] The tightening in global financial conditions induced by global risk shocks also results in a drop in RoW

[12]As Bruno & Shin (2015) we rely on data reported in Table A7 of the BIS Locational Banking Statistics. We use linear interpolation to convert the data from quarterly to monthly frequency. We measure cross-border bank credit to non-US borrowers as 'External liabilities to all sectors of all reporting banks' less 'External liabilities to all sectors of banks owned by US nationals' (see Table D.1 for variable definitions/descriptions). The advantage of the data in Table A7 is that it is based on the nationality principle so that distortions introduced through financial centers are reduced. The disadvantage is that the data only reflect information on the *liabilities* of globally active banks in BIS reporting countries, which included between 24 in the 1990s and 48 countries at the end of our sample period (BIS 2020), and therefore potentially omits globally active banks in some important EMEs. However, the coverage of the BIS reporting banks even in the 1990s amounted to about 90%.

[13]Analogously to Lilley et al. (forthcoming), Avdjiev et al. (2020) find that variation in global risk measured by changes in the VIX has been much less correlated with cross-border bank credit after the GFC. Figure C.5 documents that while exogenous innovations to global risk caused contractions in global cross-border bank credit both in the pre-GFC period until 2006 and the post-GFC period starting from 2009, the effects were indeed stronger prior to the GFC.

equity prices and an increase in the EMBI spread. And also the global factor in risky asset prices covering equity, bonds and commodities originally introduced in Miranda-Agrippino & Rey (2020) and extended in Miranda-Agrippino et al. (2020) as well as their global factors in 'portfolio' and especially 'other investment' flows—which includes bank loans—drop markedly.

Note that our results also support the notion of an 'exorbitant duty' of the US (Gourinchas et al. 2012, 2017). In particular, the US is—roughly speaking—long in RoW-currency denominated portfolio equity and direct investment assets but short in dollar-denominated safe US portfolio debt liabilities, in particular US Treasuries. In this setting, when a global risk shock appreciates the dollar, raises the prices of US debt assets and pushes down the prices of risky assets in the RoW, it brings about a wealth transfer from the US to the RoW such that the US provides insurance to the RoW. Our results are qualitatively consistent with this notion of 'exorbitant duty': the dollar appreciates, RoW equity prices drop, and yields of US Treasury securities decline (Figure 2)—implying an increase in their prices—in response to global risk shocks. A more direct test for the 'exorbitant duty' would be based on impulse responses of the US net foreign asset position, for which (quarterly) data are unfortunately only available from 2006 onwards.

3.4 Robustness

Our results are robust across a number of alternative specifications. First, if the gold price surprises on the narratively selected dates were—despite being selected carefully by Bloom (2009), Piffer & Podstawski (2018) and Bobasu et al. (2021)—*systematically* driven not only by global risk shocks, then the exogeneity condition $E[p_t^{\epsilon,r}\boldsymbol{\epsilon}_t^o] = \mathbf{0}$ in Equation (9b) would not be satisfied. In order to address this concern, as in Ludvigson et al. (2021) we relax this identifying assumption by allowing for $E[p_t^{\epsilon,r}\boldsymbol{\epsilon}_t^o] \neq \mathbf{0}$; we only impose the weaker condition that the correlation between the gold price surprises and global risk shocks is stronger than for all other structural shocks: $|E[p_t^{\epsilon,r}\epsilon_t^r]| > |E[p_t^{\epsilon,r}\epsilon_t^\ell]|$ for $\ell \neq r$. Results are very similar to the baseline (Figure C.6).

Second, one might argue that the exogeneity condition $E[p_t^{\epsilon,r}\boldsymbol{\epsilon}_t^o] = \mathbf{0}$ in Equation (9b) might not be satisfied because there being only two auctions per day and only on weekdays implies that the windows around the time stamps of the narratively selected events for which the gold price surprises are calculated are not sufficiently narrow. In particular, the wider the windows the more likely it is that the gold price surprises also reflect the effects of other shocks that occurred shortly after or before the global risk shocks. Recall that for this to compromise our identification of global risk shocks, these other shocks must have occurred *systematically* across events, which is unlikely to be the case. Nevertheless, as a robustness check we consider surprises in other asset prices which we can calculate for narrower windows around the event time stamp. Specifically, we use as proxy variable the changes in long-term Treasury yields over a -30min/+90min window or the US dollar/euro exchange rate over a -10min/+20min window.[14] Results based on these alternative specifications

[14] We extract Treasury futures prices and exchange rates from Reuters Tick History. We convert Treasury futures prices to yields and reverse the sign so that a positive surprise reflects the effects of an increase in safe asset demand. Treasury futures are traded on the Chicago Board of Trade (CBT) from Sunday 5pm to Friday 4pm. Exchange rates are traded 24/7 and are rather liquid. We choose different window lengths for Treasury futures and US dollar/euro

are very similar to the baseline (Figures C.7 and C.8).

Third, one might be worried that even in the narrower intra-daily windows on the narratively selected events that we can use for the alternative asset prices shocks other than to global risk did occur systematically. To address this concern, we consider an alternative proxy variable for which we do not have to rely on narratively selected events because it by construction reflects exogenous variation in global risk. Specifically, we consider monthly changes in the Geopolitical Risk Index of Caldara & Iacoviello (2022). Also in this case results are quite similar to the baseline (Figure C.9).

Fourth, we consider only events with *positive* gold price surprises to focus on *increases* in the demand for safe assets. Results hardly change (Figure C.10).

Fifth, we estimate a large BPSVAR model that adds the variables simultaneously instead of one at a time to estimate the impulse responses shown in Figures 3 and 4; to do so we use informative Minnesota-type priors and optimal hyperpriors/prior tightness (Giannone et al. 2015). Results are very similar to the baseline (Figure C.11). Our results are also very similar if we do not impose a relevance threshold (Figure C.12).

Finally, Lilley et al. (forthcoming) demonstrate that a variety of common measures of global risk feature significant in-sample explanatory power for exchange rates after but not prior to the GFC. While these reduced-form correlations are in general not indicative of the strength of structural relationships, we verify that our results are similar for the time periods before and after 2007; we do find though that the dollar appreciation following a global risk shock was delayed before 2007 (Figure C.13).

4 The dollar in the transmission of global risk shocks to the RoW

The results in Figure 4 are consistent with the notion that the dollar appreciation impacts the RoW both through a trade and a financial channel. Given that these channels work in opposite directions, we now assess the net effects of dollar appreciation on the RoW. We do this by constructing a counterfactual in which a global risk shock does *not* cause dollar appreciation. We first discuss how we construct the counterfactual.

4.1 Minimum relative entropy counterfactuals

In the existing literature minimum relative entropy (MRE) is used to incorporate restrictions implied by economic theory in order to improve a forecast. For example, Robertson et al. (2005) improve forecasts of the Federal Funds rate, US inflation and the output gap by imposing the constraint that the inflation forecast over the next three years must average 2.5% (see also Cogley et al. 2005; Giacomini & Ragusa 2014). As in Breitenlechner et al. (2021), we apply MRE to impulse responses as these can be conceived as conditional forecasts.

exchange rate surprises due to differences in trading activity. One could also use intra-daily changes in the VIX reported by the Chicago Board Options Exchange (CBOE), but regular trading hours only cover 8.30am to 3pm on weekdays, and hence do not bracket intra-daily many of the global risk events we consider; extended trading hours do exist for CBOE, but reporting is much less comprehensive than for regular trading hours.

Assume for simplicity but without loss of generality that the VAR model in Equation (1) is stationary, that it does not include deterministic terms, and that it is in steady state in period T. Under these assumptions, the impulse response to a global risk shock over h periods is given by the conditional forecast $\widetilde{\boldsymbol{y}}_{T+1,T+h} \equiv [\widetilde{\boldsymbol{y}}'_{T+1}, \widetilde{\boldsymbol{y}}'_{T+2}, \ldots, \widetilde{\boldsymbol{y}}'_{T+h}]'$ with $\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \equiv [\widetilde{\boldsymbol{\epsilon}}'_{T+1}, \widetilde{\boldsymbol{\epsilon}}'_{T+2}, \ldots, \widetilde{\boldsymbol{\epsilon}}'_{T+h}]'$ featuring $\widetilde{\epsilon}^r_{T+1} = 1$, $\widetilde{\epsilon}^r_{T+s} = 0$ for $s = 2, 3, \ldots, h$ and $\widetilde{\epsilon}^\ell_{T+s} = 0$ for $s = 1, 2, \ldots, h$ and $\ell \neq r$. The impulse responses $\widetilde{\boldsymbol{y}}_{T+1,T+h}$ are a function of the structural VAR parameters $\boldsymbol{\psi} \equiv vec(\boldsymbol{A}_0, \boldsymbol{A}_1)$, which have to be estimated from the sample of data $\boldsymbol{y}_{1,T}$.

Bayesian estimation of the BPSVAR model delivers the posterior belief about the effects of a global risk shock

$$f(\widetilde{\boldsymbol{y}}_{T+h}|\boldsymbol{y}_{1,T}, \mathcal{I}_a, \widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}) \propto p(\boldsymbol{\psi}) \times \ell(\boldsymbol{y}_{1,T}|\boldsymbol{\psi}, \mathcal{I}_a) \times \nu, \tag{10}$$

where $\mathcal{I}_a$ represents the identification assumptions in Equations (9a) and (9b), $p(\cdot)$ is the prior about the structural VAR parameters, and ν the volume element of the latter's mapping into the impulse responses; the blue solid lines in Figure 2 are the point-wise means of f.

MRE determines the posterior beliefs about the effects of a global risk shock in a *counterfactual* VAR model with structural parameters $\boldsymbol{\psi}^*$ as follows:

$$Min_{\boldsymbol{\psi}} \ \mathcal{D}(f^*||f) \quad \text{s.t.} \tag{11}$$
$$\int f^*(\widetilde{\boldsymbol{y}})\widetilde{y}^{\$}d\widetilde{\boldsymbol{y}} = E(\widetilde{y}^{\$}) = 0, \ \int f^*(\widetilde{\boldsymbol{y}})d\widetilde{\boldsymbol{y}} = 1, \ \ f^*(\widetilde{\boldsymbol{y}}) \geq 0,$$

where $\mathcal{D}(\cdot)$ is the Kullback-Leibler divergence—the 'relative entropy'—between the counterfactual and baseline posterior (we drop the subscripts in $\widetilde{y}^{\$}_{T+h}/\widetilde{\boldsymbol{y}}_{T+h}$ in Equation (11) for simplicity). Intuitively, there is an infinite number of counterfactual beliefs f^* that satisfy the constraint that the dollar is unresponsive to a global risk shock. The MRE approach disciplines the choice of the counterfactual beliefs f^* by requiring that they are *minimally* different from the baseline posterior beliefs f in an information-theoretic sense. Roughly speaking, MRE determines that counterfactual VAR model in which the dollar is unresponsive to a global risk shock but whose dynamic properties in terms of impulse responses are otherwise minimally different from those of the actual VAR model.[15]

It turns out that the solution for the counterfactual distribution f^* in (11) can be computed by updating the baseline posterior f given the 'information' that the dollar is unresponsive to a global risk shock. Specifically, we have

$$\begin{aligned} &f^*\left(\widetilde{\boldsymbol{y}}_{T+h}|\boldsymbol{y}_{1,T}, \mathcal{I}_a, \widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}, \widetilde{y}^{\$}_{T+h} = 0\right) \propto \\ &\qquad f(\widetilde{\boldsymbol{y}}_{T+h}|\boldsymbol{y}_{1,T}, \mathcal{I}_a, \widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}) \times \tau\left(\widetilde{y}^{\$}_{T+h}(\boldsymbol{\psi})\right), \end{aligned} \tag{12}$$

[15] Earlier studies have instead constructed counterfactuals by constraining selected VAR coefficients to zero, either before or after estimation (see, for example, Carriere-Swallow & Cespedes 2013; Vicondoa 2019; Degasperi et al. 2020; Redl 2020). Imposing zeros before estimation implies a mis-specified model and induces biased estimates that are not informative about the strength of the channel that is being shut down (Georgiadis 2017). Simply setting VAR coefficients after estimation to zero is more similar to the MRE approach, but lacks the discipline MRE imposes on the choice of the counterfactual model.

where τ is a 'tilt' function (Robertson et al. 2005). The tilt τ down-weighs the actual posterior beliefs for those values of the VAR parameters that are associated with large deviations from the counterfactual constraint that the dollar is unresponsive to a global risk shock. In practice, Robertson et al. (2005) as well as Giacomini & Ragusa (2014) show that implementing the MRE approach boils down to tilting the weights of the draws of the approximated baseline posterior distribution (see Appendix E). Once the tilted weights are obtained, importance sampling techniques can be used to estimate mean and percentiles of f^*

4.2 A 'no-appreciation' counterfactual benchmark

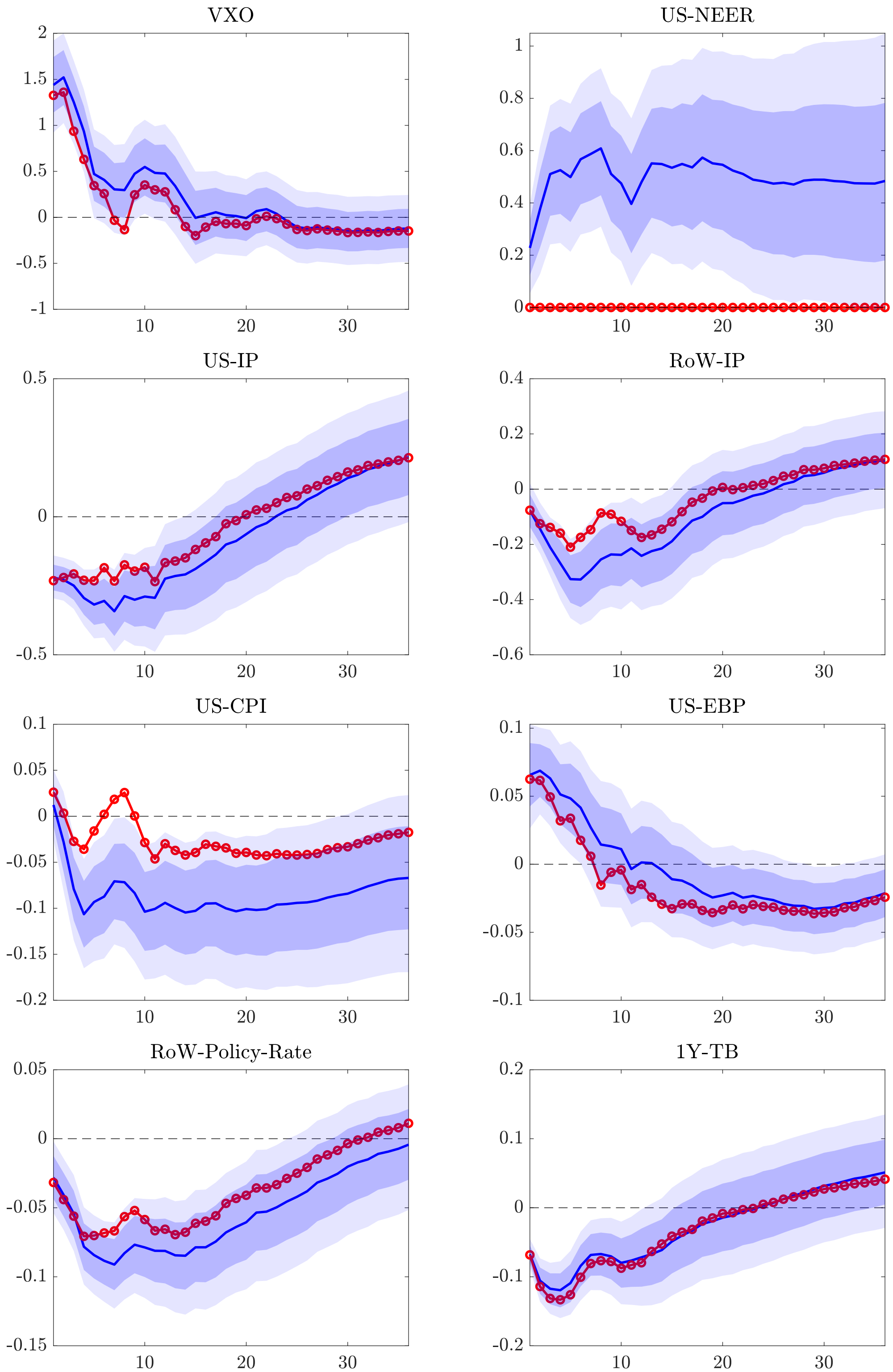
Figure 5 shows that in the counterfactual in which the dollar is unresponsive (red line with circles), a global risk shock is considerably less contractionary in the RoW: the decline in industrial production is reduced by roughly half. More generally, the effects in the counterfactual are less pronounced for all variables, except for US monetary policy. The finding that RoW industrial production falls by less in the counterfactual implies that the contractionary financial channel dominates the expansionary trade channel. This mirrors findings that financial channels dominate trade channels in the international transmission of US monetary policy shocks (Degasperi et al. 2020).

One may ask whether the differences between the baseline and the counterfactual are 'statistically significant'. This question is not as straightforward to answer as in a standard regression in which one tests the null of a coefficient estimate being equal to a non-random benchmark. Instead, in our context this involves assessing how 'different' the baseline and the counterfactual posterior distributions are. To do so, we use a test for first-order stochastic dominance. Intuitively, the baseline first-order stochastically dominates the MRE counterfactual at horizon h if for every possible value of the RoW industrial production response Y the probability of the estimated response $\widetilde{y}_h^{ip_{row}}$ being smaller than Y is higher in the baseline than in the counterfactual. Formally, the baseline first-order stochastically dominates the counterfactual at horizon h if $F^*_{\widetilde{y}_h^{ip_{row}}}(Y) < F_{\widetilde{y}_h^{ip_{row}}}(Y)$ for all Y. It turns out that the baseline first-order stochastically dominates the counterfactual for RoW industrial production essentially over all horizons in our estimation. We conclude that the MRE counterfactual is meaningfully different from the baseline.

That the financial channel dominates the trade channel also emerges from the counterfactual results for global financial conditions and trade shown in Figure 6: when dollar appreciation is absent there is only a slightly weaker drop in US exports and only a slightly stronger drop in US imports. This suggests the trade channel is not very powerful in the first place, consistent with the observed weakening of exchange rate pass-through to import prices over time that has resulted from the deepening of cross-border value chains (Ahmed et al. 2017; Georgiadis et al. 2019). We observe a much bigger difference between the baseline and the counterfactual for global financial conditions: cross-border bank credit and RoW equity prices contract by much less when dollar appreciation is absent[16]; also the EMBI spread rises by less, and the global factors in risky asset

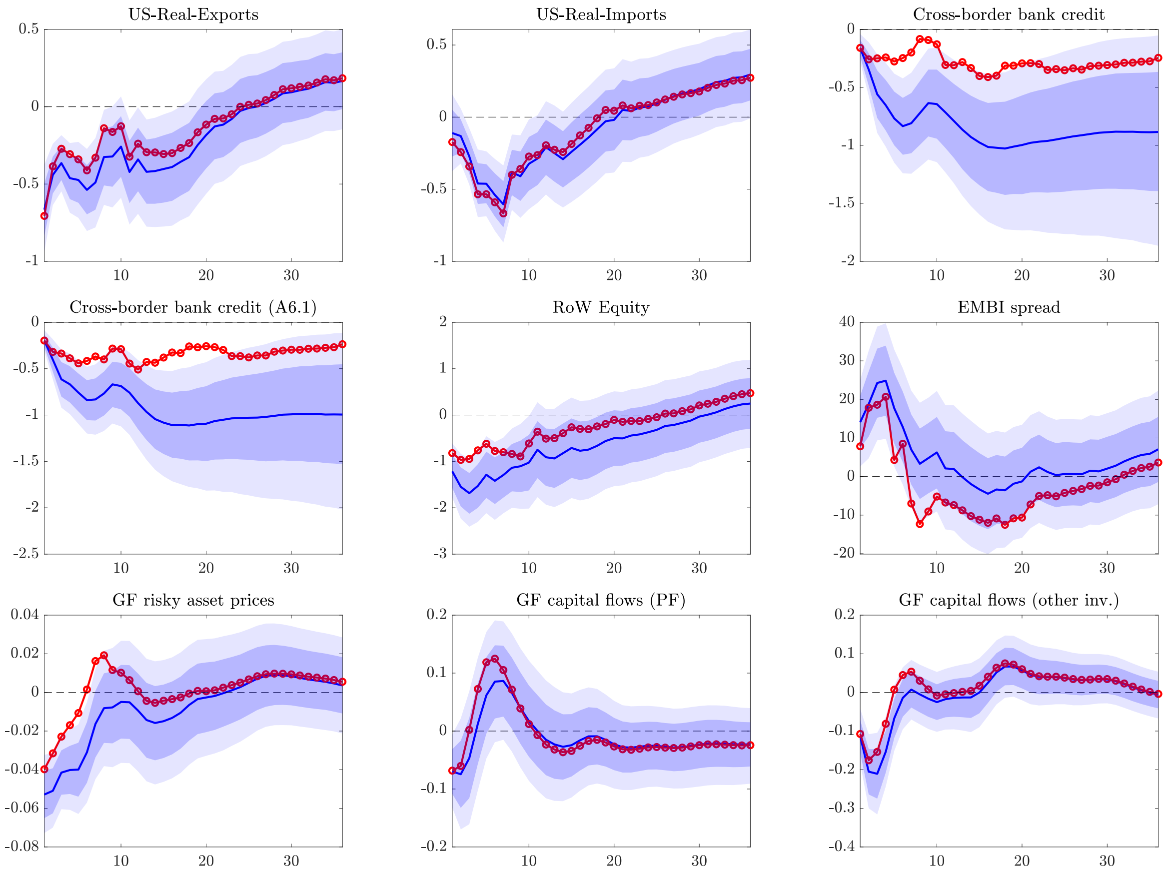
[16] A possible concern is that the strong effect of the dollar appreciation on cross-border credit in the baseline might simply reflect mechanical valuation effects due to non-dollar denominated credit flows being recorded in dollar. Figure

Figure 5: Baseline and MRE-based counterfactual responses to a global risk shock



Note: See the notes to Figure 2. The red circled lines depict point-wise means of the counterfactual posterior distribution obtained from the MRE approach.

Figure 6: Baseline and MRE-based counterfactual responses of US trade and cross-border bank credit to a global risk shock



Note: See the notes to Figures 4 and 5.

prices and—although less pronounced—in 'other investment' flows fall by less in the counterfactual.

We consider two extensions which illustrate the dollar's special role for the international transmission of global risk shocks. First, we show that suppressing appreciation of other safe-haven currencies such as the yen instead of the dollar in the counterfactual is inconsequential for the effects of global risk shocks (Appendix F). Second, consistent with Ivashina et al. (2015), we show that in the counterfactual without dollar appreciation especially *dollar*-denominated cross-border credit drops by less (Appendix G).

C.14 presents the results for the effects of a global risk shock for two alternative cross-border bank credit variables, namely for dollar-denominated cross-border bank credit—which accounts for about half of aggregate cross-border bank credit—and exchange-rate-adjusted total cross-border bank credit. Even if compared to Figure 6 the reduction in the drop in the counterfactual is indeed less pronounced for these two alternative cross-border bank credit variables, the reduction is still substantial, and in any case larger than for US exports and imports.

4.3 What if US monetary policy stabilized the dollar?

The MRE approach is purely data driven and thus agnostic as to what prevents dollar appreciation in the counterfactual. We therefore consider an alternative approach based on a structural policy experiment. A natural candidate is a monetary policy intervention, and, given the special status of the dollar in the global financial system, in particular an intervention by the US Fed.

Incidentally, such a policy experiment also speaks to one of the observations in Figure 1 that motivates this paper. In particular, during the COVID-19 pandemic the Fed provided unprecedented emergency liquidity to a number of countries. It is widely believed that these measures were helpful to prevent a global financial crisis (see, for instance, Choi et al. 2021). Theoretically, this liquidity provision can be conceived as increasing the supply of safe assets by crediting RoW central banks with dollar reserves. This, in turn, reduces the convenience yield and thereby depreciates—or dampens appreciation pressures on—the dollar (Jiang et al. 2021a). Theory thus predicts that emergency liquidity provision mitigated dollar appreciation in early-2020. Indeed, the dollar appreciation in early-2020 in the right panel of Figure 1 was relatively short-lived compared to our estimates for the typical risk event in the sample shown in Figure 2. This suggests that the Fed's emergency liquidity provision may have mitigated dollar appreciation as a side effect of the attempt to avert a global financial crisis. In turn, according to our results above, the dampening of dollar appreciation should have contributed to mitigating the pandemic's global fallout.

Against this background, we assume that after a global risk shock the Fed steps in unexpectedly to stabilize the dollar. We implement this policy experiment through a sequence of US monetary policy shocks. Specifically, we follow Antolin-Diaz et al. (2021) and consider a sequence of shocks that materializes over the impulse response horizon so as to offset the effects of the global risk shock on the dollar (see Appendix H for technical details).[17] Figure 7 shows that in this counterfactual US monetary policy as reflected by the 1-year Treasury bill rate is loosened much more than in the baseline, and the slowdown in RoW real activity is muted substantially. At the same time, there are non-trivial pressures on US consumer prices and overshooting in US real activity. Given the Fed's domestic mandate, this may be an important practical obstacle to such a policy response.[18]

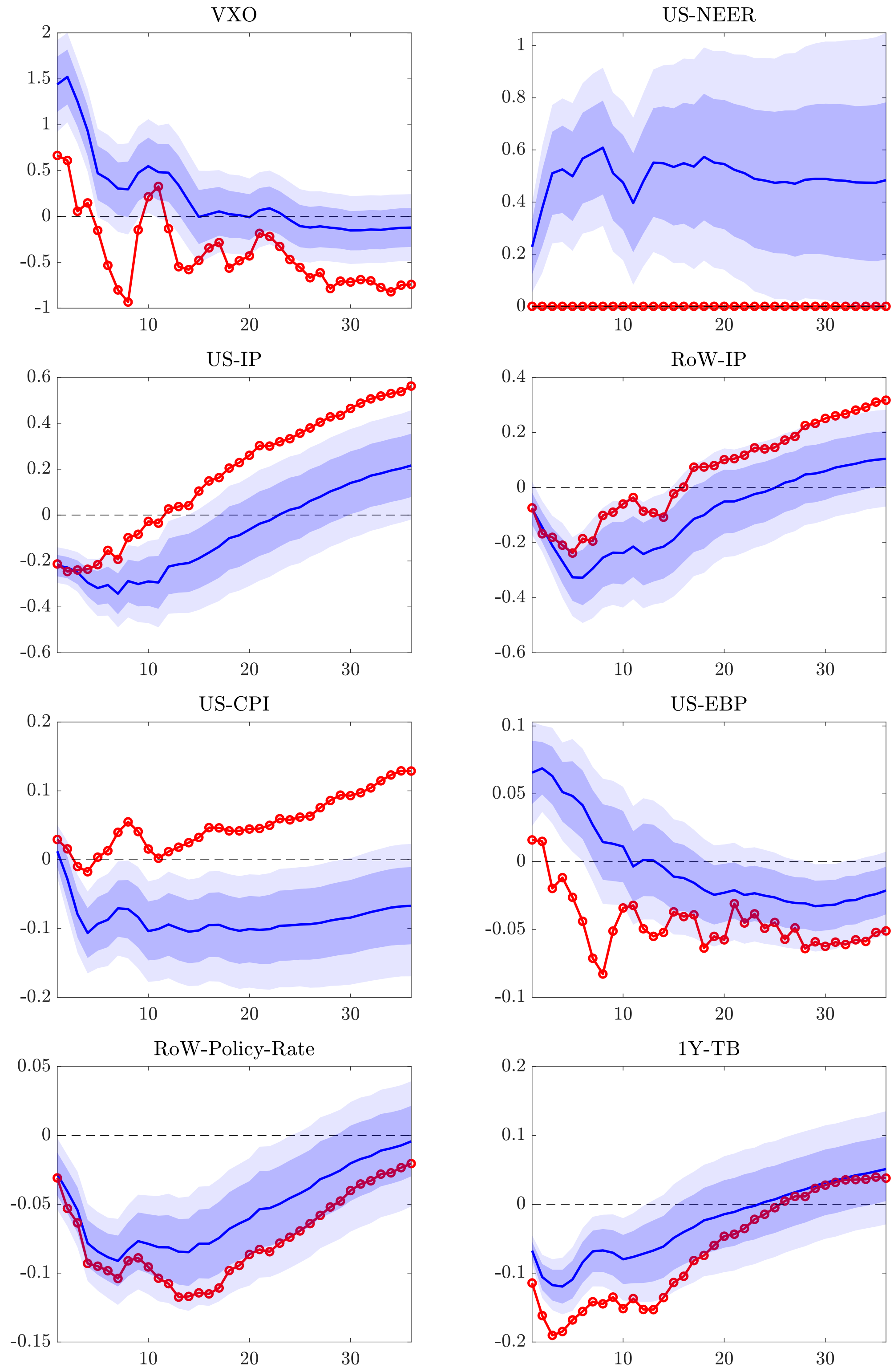
5 A DCP^2 model

The results of our time-series analysis square well with the received wisdom according to which the dollar shapes the international transmission of shocks through financial channels and trade. In what follows, we offer a structural perspective and put forward a two-country model for the world economy. The model—consistent with our empirical analysis—allows for a special role for the dollar

[17]The impulse responses to a US monetary policy shock are shown in Figure C.15. They are consistent with existing evidence for the domestic effects in the US (Gertler & Karadi 2015; Miranda-Agrippino & Ricco 2021) and for spillovers to the RoW (Georgiadis 2016; Degasperi et al. 2020; Miranda-Agrippino & Rey 2020).

[18]The results for the 'modesty statistic' of Leeper & Zha (2003) indicate that the offsetting shocks are not unusually large or persistent, see Figure C.16. Also the q-divergence of Antolin-Diaz et al. (2021) does not indicate that the distribution of shocks in the counterfactual is notably different from the baseline.

Figure 7: Baseline and counterfactual responses to a global risk shock based on US monetary policy shocks as offsetting shocks



Note: The circled red lines depict the counterfactual responses to a global risk shock.

by bringing together a dominant-currency paradigm (DCP) both in trade and cross-border banking. We refer to this framework as the 'DCP^2 model'.[19]

We first outline the structure of the model. We keep the description short and do not discuss explicitly many features that have been introduced elsewhere. We calibrate the model and show that the impulse responses to a global risk shock match the time-series evidence shown in Figure 2. In a last step, we show that DCP in *both* trade and cross-border banking are necessary for matching the empirical impulse responses. In particular, dollar pricing of trade is critical to produce a muted response of the US trade balance; and dollar cross-border credit is critical to generate the large tightening in RoW financial conditions and the contraction in RoW output.

5.1 Model structure

The model consists of two economies, the US, denoted by F, and a RoW block denoted E. The two economies are linked through trade and cross-border bank credit. The model features standard real and nominal frictions such as sticky prices and wages, habit formation in consumption and investment adjustment costs alongside Gertler & Karadi (2011)-type financial frictions adapted to allow for cross-border bank lending. These frictions are largely symmetric in the two economies with the exception of the invoicing of trade and foreign-currency lending. In particular, US banks provide dollar-denominated loans to RoW. Global risk shocks raise the perceived riskiness of these loans—consistent with the notion which underlies our empirical analysis. For the sake of brevity we relegate most of the equations to Appendix I and only state the core equations related to DCP in trade and cross-border credit.

5.1.1 Households

Each economy is populated by households and firms indexed on the unit interval. A fraction s of agents resides in the Row, the remaining fraction in the US. Following Erceg et al. (2000) we assume that within each country households are symmetric with the exception of the wage they receive and labor they supply. As in Gertler & Karadi (2011) we further assume that within each household a fraction $1 - f$ of members are workers, while the remaining fraction f are bankers. Workers supply labor, make consumption decisions and provide deposits to local banks, while bankers intermediate funds from households to firms and accumulate equity. To ensure that bankers do not end up with enough equity to fund all investments without having to rely on domestic deposits, we assume that every period a banker has to close its bank with probability $1 - \theta_b$ and transfer the accumulated equity back to the household. A corresponding number of workers randomly become bankers every period, keeping the ratio of workers to bankers fixed.

[19] Typically these features are studied in isolation. An exception is Akinci & Queralto (2018). Their focus is different from ours, however, and they only explores dominant currency pricing and dominant credit in a robustness analysis.

5.1.2 Financial intermediaries in the RoW

To account for the pattern that a large fraction of cross-border credit in the data is denominated in dollar we assume that financial intermediaries differ across countries. In the RoW, banks fund their long-term assets with equity, domestic short-term deposits, or dollar loans from US banks. Consider a generic bank j in RoW and let $S_{E,j,t}$ be its claims on RoW capital and $N_{E,j,t}$ its equity, $Q_{E,t}$ the price of such a claim expressed relative to the price of the RoW final consumption good $P^C_{E,t}$, $B_{E,j,t}$ the real value of RoW currency deposits held by RoW households, and $B^F_{E,j,t}$ the amount of dollar loans. The bank's balance sheet identity in real terms then reads as

$$Q_{E,t}S_{E,j,t} = B_{E,j,t} + RER^F_{E,t}B^F_{E,j,t} + N_{E,j,t}. \tag{13}$$

$RER^F_{E,t} = \mathcal{E}^F_{E,t}P^C_{F,t}/P^C_{E,t}$ represents the real exchange rate in terms of relative consumer prices, and $\mathcal{E}^F_{E,t}$ is the nominal dollar exchange rate defined as the price of a dollar in units of RoW currency.

Claims on RoW capital earn the nominal return $R_{K,E,t}$, deposits are paid the predetermined nominal rate $R_{E,t-1}$, and the endogenous rate $R^F_{E,b,t-1}$ is paid to US banks for cross-border loans. The law of motion for a RoW bank's real equity is then given by:

$$\begin{aligned} N_{E,j,t} = \frac{1}{1+\pi^C_{E,t}}\Big[\Big(R_{E,k,t} - \Xi^F_{E,j,t-1}\Big(\mathcal{E}^F_{E,t}/\mathcal{E}^F_{E,t-1}\Big)R^F_{E,b,t-1} - (1-\Xi^F_{E,t-1})R_{E,t-1})\Big) \\ Q_{E,t-1}S_{E,j,t-1} + R_{E,t-1}N_{E,j,t-1}\Big], \end{aligned} \tag{14}$$

with $\Xi^F_{E,j,t} = RER^{F^{dl}}_{E,t}B^F_{E,j,t}/Q_{E,t}S_{E,j,t}$ as the share of domestic investments funded by cross-border loans. All else equal, the RoW bank's net worth deteriorates if the dollar appreciates—and the more strongly so the more it relies on cross-border loans to fund its domestic lending.

The objective of a banker is to maximize expected terminal wealth, discounted with the household's real stochastic discount factor $\Theta_{E_{t,t+s}}$. Thus, the value function of the bank is

$$V_{E,j,t} = \max\ \mathbb{E}_t\sum_{s=0}^{\infty}(1-\theta_B)\Theta_{E_{t,t+s}}N_{E,j,t+1+s}. \tag{15}$$

In order to allow for the existence of a spread and to put a limit on the borrowing possibilities of RoW banks, we follow Gertler & Karadi (2011) and assume that bankers may divert a fraction $\delta_{E,B,t}$ of total assets $Q_{E,t}S_{E,j,t}$. Since the depositors know of this possibility, they are only willing to deposit with the bank if the expected terminal wealth $V_{E,j,t}$ of the bank is larger than the gain from absconding with the fraction of current periods assets. This gives rise to the incentive constraint:

$$V_{E,j,t} \geq \delta_{E,B,t}Q_{E,j,t}S_{E,j,t}. \tag{16}$$

While the solution to Equation (15) pins down the optimal leverage ratio and thereby the total amount of the bank's assets, the optimal liability composition is indeterminate up to a first-order approximation. Therefore, we follow Akinci & Queralto (2019) and Aoki et al. (2018) and impose that the fraction of assets that can be diverted by the banker is a function of the current-period portfolio liability composition. More specifically, we assume that $\delta_{E,B,t}$ varies over time as a function of the dollar funding share $\Xi_{E,j,t}$ according to

$$\delta_{E,B,t} = \overline{\delta}_{E,B}\left(1 + \epsilon_B \Xi_{E,j,t} + \frac{\kappa_B}{2}\Xi^{F^2}_{E,j,t}\right), \tag{17}$$

with $\epsilon_B < 0$ and $\kappa_B > 0$. This specification implies that in terms of leverage, there is an interior optimum of dollar funding. This accounts for the observation that in the data—especially EME—savers (that is the creditors of RoW banks) have a preference for liability dollarization of their bank in terms of offering dollar deposits, as this allows savers to insure against domestic currency depreciations ($\epsilon_B < 0$) (see Bocola & Lorenzoni 2020; Gopinath & Stein 2021). Yet, as liability dollarization exposes the bank to exchange rate risk, it is reasonable to assume that creditors do not favor complete dollarization of liabilities. Therefore, we set $\kappa_B > 0$ as in Akinci & Queralto (2018).

The first-order condition for the optimal choice of the share of cross-border loans $\Xi^F_{E,j,t}$ yields a modified UIP condition

$$\begin{aligned}\mathbb{E}_t\Big(\Omega_{E,t,t+1}\Big[R_{E,t} - \left(\mathcal{E}^F_{E,t+1}/\mathcal{E}^F_{E,t}\right) R^F_{E,b,t}\Big]\Big)& \\ = \left(\frac{\delta_{E,B,t}}{\delta'_{E,B,t}} - \Xi^F_{E,t}\right)^{-1} &\mathbb{E}_t\Big(\Omega_{E,t,t+1}\left[R_{E,k,t+1} - R_{E,t}\right]\Big),\end{aligned} \tag{18}$$

where $\Omega_{E,t,t+1}$ is the real stochastic discount factor of RoW households. Equation (18) points to two sources of UIP deviations. First, RoW banks might not be able to borrow at the US riskless rate $R^F_{E,b,t} \neq R_{F,t}$. Second, the time-varying incentive compatibility constraint $\delta_{E,B,t}$ endogeneously links the UIP deviation to the RoW credit spread and thereby the quality of RoW banks' balance sheets. In particular, to the extent that the term multiplying the credit spread in Equation (18) is negative, a rise in the RoW credit spread—arguably a proxy for risk aversion in the financial sector—will cause negative deviations from the UIP condition and an immediate appreciation of the dollar.[20]

5.1.3 Financial intermediaries in the US

US banks differ from RoW banks in two respects. First, US banks act as *lenders* rather than borrowers in the global dollar market, and so cross-border loans $B^{F^\star}_{E,j,t}$ appear on the asset side of the balance sheet. The balance sheet identity of a US bank j deflated by the price of the US

[20]The intuition is the following: Given a deteriorated quality of RoW banks' balance sheets (which come with a rise in the RoW credit spread), it becomes increasingly costly to have dollar loans on the balance sheet. RoW banks would like to exploit the arbitrage opportunities arising from the increased credit spread, but the dollar loans on their balance sheet tighten the incentive constraint at the margin. In order for RoW banks to be willing to at least hold some dollar loans, the RoW currency has to appreciate in the following period such that not only interest rate payments to US banks—which have to be paid in dollar—are equal to the domestic borrowing costs $R_{E,t}$ when converted into RoW currency, but RoW banks also need to be compensated for the forgone profits of the tightened balance sheet constraint. This additional expected depreciation of the dollar is partly achieved by a stronger contemporaneous appreciation.

consumption good reads as

$$Q_{F,t}S_{F,j,t} + B^{F^\star}_{E,j,t} = B_{F,j,t} + N_{F,j,t}, \tag{19}$$

with $S_{F,j,t}$, $B^{F^\star}_{E,j,t}$, $B_{F,j,t}$ and $N_{F,j,t}$ as the total amount of claims on capital, cross-border loans, domestic deposits and net worth, respectively. Furthermore, the law of motion for a bank's net worth is given by

$$\begin{aligned} N_{F,j,t} = & \frac{1}{1+\pi^C_{F,t}}\Big[(R_{K,F,t} - R_{F,t-1})Q_{F,t-1}S_{F,j,t-1} \\ & + (R^F_{E,b,t-1} - R_{F,t-1})B^{F^\star}_{E,j,t-1} + R_{F,t-1}N_{F,j,t-1}\Big]. \end{aligned} \tag{20}$$

Thus, a US bank is shielded from exchange rate movements as both its liabilities and assets are denominated in dollar, while a RoW bank's net worth is exposed to exchange rate risk.

The second difference between US and RoW banks arises from the incentive compatibility constraint. In particular, the fraction of assets $\delta_{F,B}$ that a US bank can divert is constant. But as in the closed-economy model of Karadi & Nakov (2021), the creditors of the bank may attach a different 'risk weight' Γ_t to the different assets (see Coenen et al. 2018, for a similar interpretation). Therefore, creditors of a US bank require that its expected terminal wealth satisfies

$$V_{F,j,t} \geq \delta_{F,B}(Q_{F,t}S_{F,j,t} + \Gamma_t B^{F^\star}_{E,j,t}). \tag{21}$$

The optimal portfolio choice of the US bank imposes a no-arbitrage relation between the returns on cross-border and domestic lending given by

$$\Gamma_t \mathbb{E}_t\Big(\Omega_{F,t,t+1}\big[R_{k,F,t+1} - R_{F,t}\big]\Big) = \mathbb{E}_t\left(\Omega_{F,t,t+1}\Big[R^F_{E,b,t} - R_{F,t}\Big]\right). \tag{22}$$

We assume the risk weight Γ_t evolves according to

$$\Gamma_t = \overline{\Gamma}(1 + \epsilon_{\Gamma,t}), \tag{23}$$

where $\epsilon_{\Gamma,t}$ follows an AR(1) process, $\epsilon_{\Gamma,t}$ represents the global risk shock, that is a shock to the relative 'riskiness' of global lending activities of US banks, which ceteris paribus increases the returns required for US banks to engage in cross-border lending as demand for these 'risky' assets falls.[21] Recall that the counterpart for US banks' cross-border lending are leveraged RoW banks, whereas domestic lending is extended to non-financial firms. Therefore, as US banks are the only global financial intermediaries, a shift in relative riskiness between cross-border and domestic lending represents an increase in global investors' risk aversion. As such $\epsilon_{\Gamma,t}$ roughly corresponds to our definition of the global risk shock in the VAR, i.e. an exogenous increase in the demand for safe assets (which implies a fall in the demand for risky assets).

[21] Our risk shock is similar in spirit to the 'safety shock' in Kekre & Lenel (2021)

5.1.4 Production and pricing of final goods

The second key element in our model is DCP in bilateral trade between the US and the RoW, following the seminal work of Gopinath et al. (2020). This means that the prices of both US and RoW exports are generally sticky in dollar. In our model we introduce a refinement. In particular, Boz et al. (2022) document that a large share of trade among countries in the RoW—even beyond commodities—is also priced in dollar. This feature of the world economy may be consequential for the effects of dollar appreciation in the context of a global risk shock, and hence we incorporate it in the model. Because all non-US countries are implicitly aggregated into the RoW in the model we introduce DCP in third-country trade, assuming that a share of *domestic* sales of RoW firms is priced in dollar. For this purpose, we consider a multi-layered production structure as in Georgiadis & Schumann (2019) and laid out in Figure C.17 in the Appendix.

At the top layer, the RoW final consumption and investment good $Y_{E,t}^C$ is put together by a continuum of firms that operate under perfect competition, see Figure C.17. They combine a RoW final domestic good $Y_{E,t}^E$ and a RoW final import good $Y_{F,t}^E$ employing a constant elasticity of substitution (CES) technology

$$Y_{E,t}^C = \left[n_E^{\frac{1}{\psi_f}} Y_{E,t}^{E\,\frac{\psi_f - 1}{\psi_f}} + (1 - n_E)^{\frac{1}{\psi_f}} Y_{F,t}^{E\,\frac{\psi_f - 1}{\psi_f}} \right]^{\frac{\psi_f}{\psi_f - 1}}. \tag{24}$$

The parameter n_E governs the share of domestically produced goods and thereby the degree of home bias.[22] The parameter ψ_f corresponds to the elasticity of substitution between the final domestic and import good. As aggregation within the auxiliary sectors takes place using a CES production function, the first-order conditions are standard for almost all stages of the bundling process and therefore only provided in Appendix I.

RoW final domestic good We assume that markets are partly segmented and firms set different prices in different markets depending on demand conditions. And, as explained above, in order to incorporate in the model the dollar pricing in third-country trade observed in the data we assume that a fraction of RoW firms $1 - \gamma_E^E$ set their prices for domestic sales in dollar. As in Gopinath et al. (2020), we assume that firms cannot choose their pricing currency, but are assigned to it exogenously and do not change it over time.

The firms that put together the RoW final domestic good $Y_{E,t}^E$ shown on the right side in Figure C.17 operate under perfect competition and combine inputs $\widetilde{Y}_{E,t}^E$ and $\widehat{Y}_{E,t}^E$ using a CES technology. The inputs are produced by two branches of firms that also operate under perfect competition and combine RoW retail goods. The firms in the first branch combine RoW retail goods $\widehat{Y}_{E,t}^E(i)$ priced in dollar (DCP goods) into the RoW final DCP good $\widehat{Y}_{E,t}^E$; analogously, the firms in the second branch combine RoW retail goods $\widetilde{Y}_{E,t}^E(i)$ priced in the producer's currency (PCP goods) into the

[22] The home bias parameter is adjusted in order to take into account differences in country size as in Sutherland (2005). In particular, given a degree of general trade openness op_E and the relative size of the RoW s, n_E takes the value $n_E = 1 - op_E(1 - s)$ with a similar adjustment for the US counterpart.

RoW final PCP good $\widetilde{Y}^E_{E,t}$. RoW retail goods producing firms buy and repackage RoW intermediate goods, operate under monopolistic competition and serve the RoW as well as the US market; for simplicity Figure C.17 only shows their domestic sales. The share of RoW retail goods producing firms whose domestic sales prices are sticky in dollar is given by $(1-\gamma^E_E)$. Therefore, $(1-\gamma^E_E)$ also reflects the degree to which movements in the dollar exchange rate cause fluctuations in the RoW aggregate producer-price index $P^E_{E,t}$.

Imports As shown on the left side in Figure C.17, the RoW import good $Y^E_{F,t}$ is produced analogously to the RoW final domestic good $Y^E_{E,t}$. In particular, the RoW final import good producers use inputs from two branches of firms that operate under perfect competition and aggregate goods from US retail goods producers. The latter operate under monopolistic competition and set prices that are either sticky in the producer's currency (PCP goods) or in the importer's currency (LCP goods). Likewise, we assume that when exporting a fraction $(1-\gamma^E_F)$ of RoW and $(1-\gamma^F_E)$ of US retailers face prices that are sticky in the currency of the importer, while the prices of the remaining firms are sticky in the producer's currency. While not shown in Figure C.17, notice for future reference that for RoW exports to the US the resulting US import price index expressed in dollar is (up to a first-order approximation) a weighted average of the RoW DCP and PCP good bundle. As a fraction $(1-\gamma^E_F)$ of RoW exporters sets their prices in dollar, the importance of the exchange rate $\mathcal{E}^F_{E,t}$ for movements in the import price depends on γ^E_F. The larger γ^E_F the more a nominal appreciation of the dollar *ceteris paribus* causes a fall in the US import price index and thereby an increase in the demand for import goods.

Retail goods firm pricing There exists a continuum of firms that operate under monopolistic competition and use intermediate goods to produce a retail good that is eventually sold to the specialized branches farther up. Each retail firm sells its product in the domestic and foreign markets; as mentioned above, for simplicity we only show sales to RoW in Figure C.17. When selling in the RoW (i.e. domestic) market, a fraction γ^E_E of RoW retail goods producing firms sets prices in RoW currency, while the remaining $(1-\gamma^E_E)$ share of firms sets their prices in dollar. A similar setting exists in the market for US imports, with γ^E_F indicating the fraction of RoW firms that price their exports in the producer's currency. Regardless of the pricing currency, all firms use the same production technology and face the same factor costs. Because firms are subject to Calvo-style price-setting frictions and can only change their price with a probability $(1-\theta^E_p)$ each period, the mark-up of a firm whose price is sticky in dollar fluctuates with the exchange rate. As each firm sets its price in the domestic RoW and US import markets optimally and as in each market firms may be subject to different pricing currencies, the profit functions differ across firms and markets as documented in Appendix I. As standard in Calvo-style price setting, firms choose their optimal reset price given demand and their pricing currency while taking into account that they might not be able to reset their price in the future. For instance, the optimal price $\hat{P}^E_{E,t}(i)$ for a DCP firm i for its sales in the RoW market is determined by

$$\max_{\hat{P}^E_{E,t}(i)} \mathbb{E}_t \sum_{s=0}^{\infty} \theta_p^{E^s} \Theta_{E_{t,t+s}} \Big[\mathcal{E}^E_{E,t} \hat{P}^E_{E,t}(i) \hat{Y}^E_{E,t}(i) - MC_{E,t} \hat{Y}^E_{E,t}(i) \Big]. \tag{25}$$

5.1.5 Monetary policy

The monetary authority sets the domestic short-term nominal interest rate. The reaction function corresponds to a standard Taylor-rule with inertia, where the central bank reacts to deviations of final (consumption) good inflation $\pi^C_{E,t}$ and real GDP $Z_{E,t}$ from steady state.

5.2 Calibration

We generally allow the parameter values to differ across the US and the RoW (see Table I.4 in the Appendix). For parameters that govern conventional model elements to the extent possible we use estimates from the literature. In particular, for US parameters we rely on Justiniano et al. (2010). For the RoW it is more difficult to find suitable estimates, as it reflects an aggregate of countries. Since the euro area accounts for roughly one quarter of the RoW in the data in terms of output, we use the estimates in Coenen et al. (2018) for many of the the RoW parameters. We next discuss the calibration of the parameters that govern DCP in trade and cross-border credit.

Regarding DCP in trade we first calibrate the relative country size s such that the steady-state share of US real GDP in global output is roughly 25%, which corresponds to the average share in the data when measuring GDP at current dollars over the period from 1990–2019. Given the country sizes, we set the general RoW openness vis-à-vis the US (op_E) such that the steady-state share of imports from the US in the aggregate RoW bundle $(1-\eta_E)$ is roughly 5.1%, again in line with the data. In the same vein we also choose the US trade openness (op_F) such that the share of imports in the US bundle $(1-\eta_F)$ is roughly 14%. We set the share of RoW firms that face sticky prices when exporting to the US $(1-\gamma^E_F)$ to 93%, in line with invoicing shares documented in Gopinath (2015). Based on the calculations in Georgiadis & Schumann (2019) we assume that US exporters almost exclusively face sticky prices in dollar and set γ^F_E to 3%. We set the share of intra-RoW *trade* that is subject to DCP $(1-\gamma^E_E)$ to 9%, which implies that 37.5% of intra-RoW *exports* are priced in dollar as indicated by the invoicing data in Boz et al. (2022).[23]

The parameters which govern the currency denomination of cross-border credit determine the strength of the balance sheet constraints for the Row and US banks as well as their preferences for cross-border loans. For the US, we specify that the local lending leverage ratio ($\phi^F_{F,B,ss} = Q_{F,ss}K_{F,ss}/N_{F,ss}$) has to equal 4 in steady state, which is the value proposed in the closed-economy model of Gertler & Karadi (2011). Furthermore, we impose that in the steady state the domestic credit spread ($S_{F,ss} = R_{K,F,ss} - R_{F,ss}$) equals 200 basis points, which roughly corresponds to the average of the

[23]We first calculate the fraction of intra-RoW trade (global exports without US imports and exports) over global non-US GDP and then take the yearly average from 1990-2019 ($\approx$ 24%). Next, we use the average share of global exports invoiced in dollar as calculated in Boz et al. (2022) and subtract the fraction of US trade in global trade to arrive at 37.5%. Multiplying the two numbers we arrive at about 9%.

GZ-spread developed in Gilchrist & Zakrajsek (2012). These two assumptions imply the values for the start-up fund parameter (ω_B^F) and the fraction of assets that the banker can divert ($\delta_{B,F}$) shown in Table I.4. We follow Akinci & Queralto (2019) and impose an average planning horizon of an international bank of six years, which implies a value for θ_B^F of 0.95. Lastly, we set the steady-state relative risk weight $\overline{\Gamma}$ to 1, which implies that *a priori* cross-border credit is perceived equally risky as domestic credit.

For the parameters of RoW banks we follow Coenen et al. (2018) and specify a steady-state leverage ratio of 6. Furthermore, we impose that the steady-state RoW capital spread is also 200 basis points. We assume that the ratio of dollar liabilities to total assets for non-US banks ($\Xi_{E,s}^F$) equals 0.18, which roughly corresponds to the average liability structure of non-US, internationally active banks in the BIS Locational Banking Statistics. These targets pin down the values for the start-up funds parameter (ω_B^E), the share of funds that a bank could divert in the absence of foreign borrowing ($\overline{\delta}_{B,E}$), and the linearity parameter (ϵ_B) in the incentive constraint which ensures that the RoW Bank is willing to take on some dollar loans in steady state. Lastly, we set the tightening parameter (κ_B) in the incentive constraint such that the derivative of the UIP deviation with respect to the share of dollar funding (Ξ_E^F) evaluated at the steady state equals the implied value in the specification of Akinci & Queralto (2019).

5.3 Confronting the model with the evidence

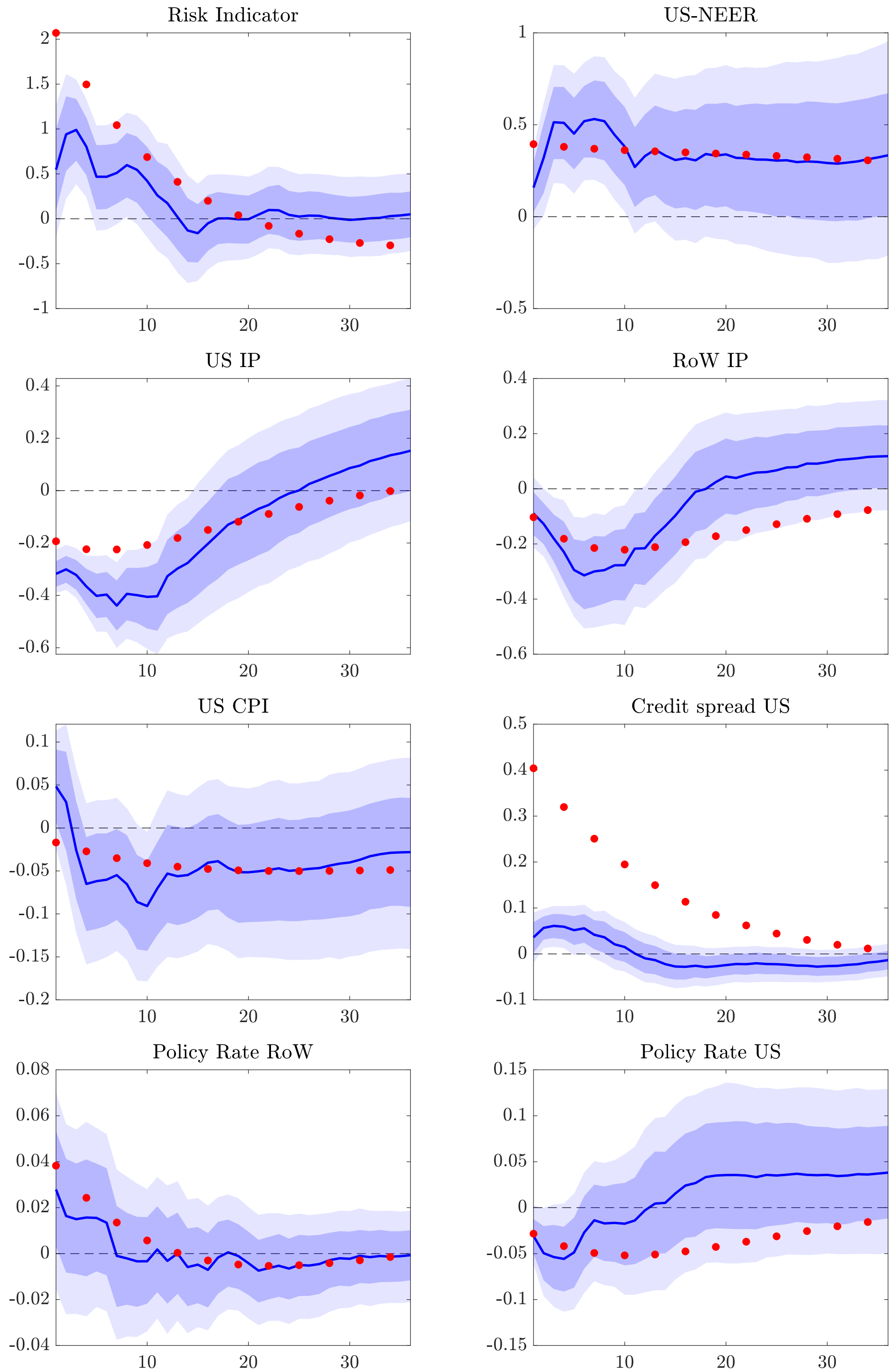
Figure 8 compares the theoretical impulse responses to a global risk shock from the DCP^2 model (red dots) to our empirical estimates from the BPSVAR model in Figure 2 (blue solid lines).[24] Remarkably, for most variables the responses in the DCP^2 model are even quantitatively similar to those estimated in the data—even though they have not been used as calibration targets. And those variables whose impulse responses we do not match very well correspond only broadly to their empirical counterparts, namely the VIX and the US risk-adjusted leverage ratio as well as the EBP and the domestic US credit spread.[25]

Given the empirical success of the model, we use it to shed additional light on the transmission of global risk shocks. For this purpose, we present additional impulse responses in Figure C.18 in the Appendix. A direct effect of the global risk shock is that RoW banks are perceived to be more risky and US banks' balance sheet constraints tighten. In order to compensate for this tightening, US

[24]We make several adjustments to render the impulse responses comparable. First, the DCP^2 model is calibrated to quarterly data, while the BPSVAR is estimated on monthly data. Therefore, the DCP^2 model impulse responses are only plotted for the first month of each quarter in Figure 8. Second, while the DCP^2 model features real GDP, the BPSVAR model is estimated on industrial production, which is about two and a half times more volatile in quarterly data. We therefore multiply the real GDP response in the DCP^2 model by 2.5. Third, in order to account for differences in the size of the shock, we re-scale the responses of the DCP^2-models that the dollar exchange rate appreciates by as much (on average over 12 quarters) as in the BPSVAR model. Finally, because in the baseline version of the BPSVAR model we include only the AE policy rate, in Figure C.18 we show the response of the RoW policy rate estimated from a modified BPSVAR model.

[25]In the DCP^2 model, the credit spread is the return from borrowing funds at the policy rate and lending those funds to a firm which transfers all returns from production as well as the remaining value of the acquired capital back to the lender. In the data, the EBP reflects the component of the corporate bond credit spread of senior unsecured bonds over synthetic risk-free securities that exactly mimic the cash flows of the bonds, net of expected default risk.

Figure 8: Responses to a global risk shock in the DCP^2 and BPSVAR models



Note: Solid blue lines show BPSVAR model responses reproduced from Figure 2. Red dots show impulse responses of the DCP^2 model. Footnote 24 provides further details on the scaling of impulse responses.

banks charge higher cross-border credit spreads from RoW banks. In turn, this induces the dollar to appreciate so that in the modified UIP condition in Equation (18) borrowing costs in RoW currency are equalized for cross-border credit and domestic deposits, up to the endogenous UIP deviation.[26]

In the US, the increase in the riskiness of banks requires them to raise the domestic credit spread and to reduce domestic lending in order to comply with their incentive constraint. The widening of credit spreads reduces investment and output contracts. The decline in the demand for capital through investment leads to a fall in the price of capital, which reduces the net worth of banks and raises their leverage ratio. This triggers a financial accelerator mechanism. The dollar appreciation furthermore leads to an increase in RoW and—although mitigated by DCP in trade—a decline in US import prices. This causes expenditure switching away from US goods and thereby a decline in the US trade balance, further amplifying the contraction in output. The drop in US output and US import prices induce a fall in the price of the US consumption good, which prompts US monetary policy to loosen.

In the RoW, banks' net worth drops as the local-currency value of their dollar liabilities and the cross-border credit spread rise. This, too, triggers a financial accelerator effect: Domestic borrowing costs rise, investment and output decline, the price of capital and net worth fall further, and the leverage ratio rises. Moreover, as the local-currency value of dollar-denominated cross-border credit increases faster than RoW banks can reduce their lending, the incentive constraint of RoW banks further tightens giving rise to a UIP premium, which further appreciates the dollar, triggering another round of financial amplification.Finally, because imports from the US as well as a non-trivial share of intra-RoW transactions are subject to DCP, the appreciation causes the price of the consumption good to rise and monetary policy to tighten.

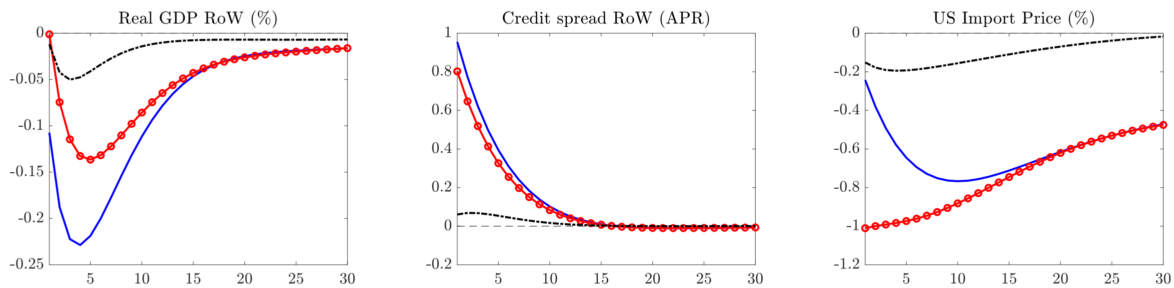
Compared to the US, the increase in the domestic credit spread through the financial accelerator in the RoW is larger, putting more downward pressure on output; moreover, monetary policy in the RoW is tightened rather than loosened. At the same time, the expenditure switching that manifests in the deterioration of the US trade balance somewhat mitigates the slowdown in RoW real activity. As a result, the global risk shock causes a fairly symmetric contraction in real activity in the US and the RoW.

5.4 The square in the DCP^2 model is crucial

The distinct feature of the model is that it incorporates DCP in credit *and* trade—hence, the name, DCP^2. It turns out that both are crucial for the model's success in matching the time-series evidence. We illustrate this as we show impulse responses for two alternative versions of the model. First, we assume cross-border credit is denominated in RoW currency instead of dollar, that is there is no dollar dominance in cross-border credit (black dash-dotted lines). Second, we assume prices are all sticky in the producer's currency instead of dollar, that is there is no dollar dominance in trade

[26] The intuition is that investors would borrow RoW currency, exchange it for dollar and then lend the proceeds to RoW banks to exploit the increase in returns on dollar borrowing. The additional demand for dollar would cause an appreciation. As all investors know about this arbitrage opportunity, in equilibrium the dollar appreciates and is then expected to depreciate to preclude these riskless returns.

Figure 9: Responses to a global risk shock w/ and w/o DCP^2



Note: blue solid lines show the impulse responses of baseline model, the red lines with circles from an alternative model in which DCP in trade is replaced by PCP, and the black dash-dotted lines from an alternative model in which cross-border credit is denominated in RoW currency instead of dollar.

(both across borders and within the RoW; red dotted lines). Figure 9 shows result for variables that exemplify the role of dollar dominance in trade and cross-border credit for the transmission of the global risk shock to RoW output.

Comparing the blue solid lines for the baseline and the red dotted (no dollar dominance in trade) and the black dash-dotted (no dollar dominance in credit) lines suggests that a global risk shock induces a large output contraction in the RoW only when there is a dollar dominance in *both* cross-border credit *and* trade. The spike in the RoW credit spread and the sharp decline in output through powerful cross-border financial spillovers occurs only with a dollar dominance in cross-border credit. In particular, in this case, dollar appreciation results in an exchange-rate valuation effect that raises the local-currency value of RoW banks' dollar liabilities, which reduces their net worth and induces them to tighten lending conditions (Bruno & Shin 2015; Akinci & Queralto 2019). On the other hand, without a dollar dominance in trade, the dollar appreciation in response to the global risk shock reduces US import prices (Gopinath et al. 2020), which triggers expenditure switching away from US towards RoW goods and hence dampens the contraction in the RoW. Moreover, without dollar dominance in trade, RoW consumer prices do not rise due to intra-RoW transactions are priced in dollar. Hence, RoW monetary policy is tightened less, which also reduces contractionary pressures on output (Mukhin 2022; Zhang 2022).

6 Conclusion

In this paper we document that global risk shocks cause a slowdown in world real activity and an appreciation of the US dollar. Other key financial indicators adjust in line with predictions of recent theoretical work. Overall, our results lend empirical support to theoretical models that rationalize the special role of the dollar and US assets in the international monetary system (Farhi & Gabaix 2016; Bianchi et al. 2021; Jiang et al. 2021a; Kekre & Lenel 2021).

In order to understand the implications of the dollar's dominance for the international transmission of global risk we run counterfactual experiments. They show that without dollar appreciation, the slowdown in economic activity outside of the US would be much weaker, reflecting the contractionary effect of the financial channel in the face of a dollar appreciation. Absent the appreciation, the burden of adjustment would fall disproportionately on the US. Our results thus illustrate how the dollar's dominance shapes the international adjustment mechanism. This raises interesting normative questions about the design of the international financial architecture. These are, however, beyond the scope of the present paper.

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A Online Appendix - Advantages of the BPSVAR framework over the traditional frequenstist external instruments SVAR framework

The BPSVAR framework has several appealing features relative to traditional frequentist external instrument SVAR models that render it particularly well-suited for the purpose of estimating the effects of global risk and US monetary policy shocks on the world economy. First, it requires relatively weak additional identifying assumptions when more than one structural shock is to be identified by proxy variables. In this case, the shocks are only set identified as rotations of the structural shocks $\boldsymbol{Q}\boldsymbol{\epsilon}_t^*$ with orthonormal matrices $\boldsymbol{Q}$ also satisfy the relevance and exogeneity conditions in Equations (2a) and (2b). Therefore, additional restrictions are needed in order to point-identify the structural shocks in $\boldsymbol{\epsilon}_t^*$. In the frequentist external instruments VAR model these additional restrictions are imposed on the contemporaneous relationships between the *endogenous variables* $\boldsymbol{y}_t$ reflected in $\boldsymbol{A}_0^{-1}$ (Mertens & Ravn 2013; Lakdawala 2019). However, Arias et al. (2021) show that relaxing this type of additional identifying assumptions can change the results profoundly. Instead, the BPSVAR framework allows us to impose the additional identifying assumptions on the contemporaneous relationships between the *structural shocks* $\boldsymbol{\epsilon}_t^*$ *and proxy variables* $\boldsymbol{m}_t$ reflected in $\boldsymbol{V}$ in the relevance condition in Equation (2a). For example, we can impose the restriction that a particular structural shock does not affect a particular proxy variable. Restrictions on the contemporaneous relationships are arguably weaker for structural shocks and proxy variables in $\boldsymbol{V}$ than for the endogenous variables in $\boldsymbol{A}_0^{-1}$.

Second, the BPSVAR framework allows coherent and exact finite sample inference, even in settings in which the proxy variables are weak instruments and only set rather than point identification is achieved with a combination of sign, magnitude and zero restrictions (see Moon & Schorfheide 2012; Caldara & Herbst 2019; Arias et al. 2021). In particular, frequentist external instruments VAR models are estimated in a two-step procedure (Mertens & Ravn 2013; Gertler & Karadi 2015): (i) estimate the reduced-form VAR model; (ii) regress the reduced-form residuals on the proxy variable to obtain the structural parameters. This two-step procedure is inefficient, as the estimation of the reduced-form VAR model in (i) is not informed by the proxy variable. In contrast, the BPSVAR model considers the joint likelihood of the endogenous variables and the proxy variables based on Equation (3), so that the proxy variables inform the estimation of both reduced-form and structural parameters. The BPSVAR framework also facilitates inference, as the joint estimation captures all sources of uncertainty. Furthermore, as long as the prior distribution is proper, in a Bayesian setting inference is straightforward even when the instruments are weak (Poirier 1998). By contrast, frequentist external instruments VAR models require an explicit theory to accommodate weak instruments (Montiel Olea et al. 2021), either to derive the asymptotic distributions of the estimators or to ensure satisfactory coverage in bootstrap algorithms.[27]

[27] To the best of our knowledge, there is no consensus yet on how to conduct inference in frequentist external instruments VAR models, even in a setting with only a single proxy variable (Jentsch & Lunsford 2019).

Third, from Equation (7) it can be seen that the BPSVAR framework is relatively flexible in that it allows for the proxy variables to be serially correlated and to be affected by lags of the endogenous variables as well as by measurement error. This is a useful feature as it has been shown that some widely-used proxy variables are serially correlated and/or contaminated by measurement error (Miranda-Agrippino & Ricco 2021). In these cases, it is typically proposed to cleanse the proxy variables in an additional step preceding the analysis in the VAR model, exacerbating issues regarding efficiency and coherent inference.

And fourth, the BPSVAR model allows us to incorporate a prior belief about the strength of the proxy variables as instruments based on the notion that "researchers construct proxies to be relevant" (Caldara & Herbst 2019, p. 165). In particular, consider the 'reliability matrix' $\boldsymbol{R}$ derived in Mertens & Ravn (2013) given by

$$\boldsymbol{R} = \left(\boldsymbol{\Gamma}_{0,2}^{-1'}\boldsymbol{\Gamma}_{0,2} + \boldsymbol{V}\boldsymbol{V}'\right)^{-1}\boldsymbol{V}\boldsymbol{V}'. \tag{A.1}$$

Intuitively, $\boldsymbol{R}$ indicates the share of the total variance of the proxy variables that is accounted for by the structural shocks $\boldsymbol{\epsilon}_t^*$ (see Equation (7)). Specifically, the minimum eigenvalues of $\boldsymbol{R}$ can be interpreted as the share of the variance of (any linear combination of) the proxy variables explained by the structural shocks $\boldsymbol{\epsilon}_t^*$ (Gleser 1992).

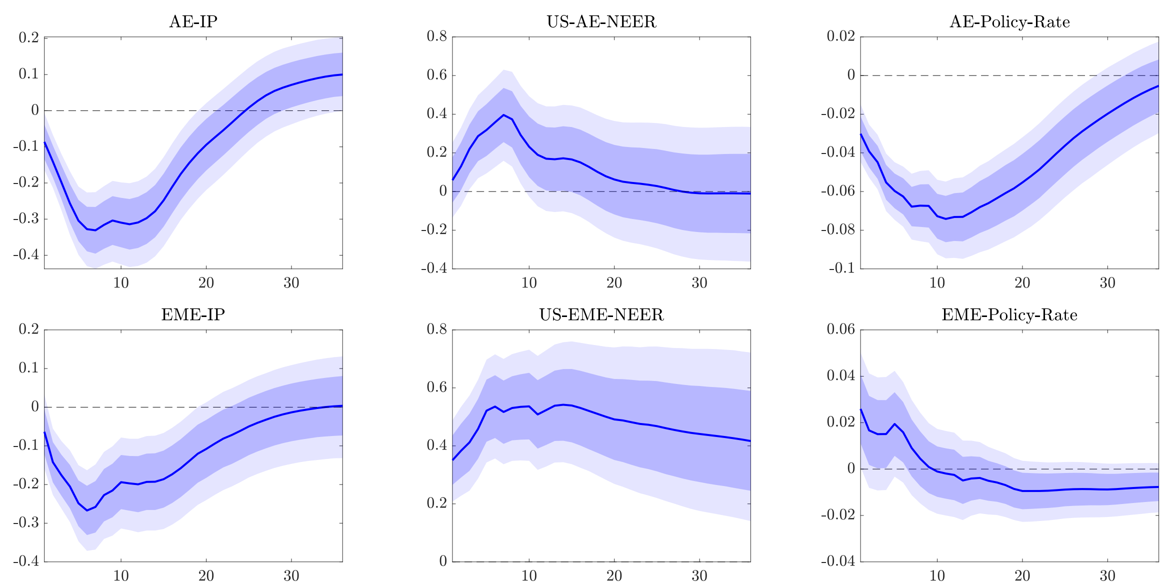
B Online appendix - Effects in AEs and EMEs

A large body of work discusses differences in the severity of financial frictions across AEs and EMEs (Caballero et al. 2008; Mendoza et al. 2009; Coeurdacier et al. 2015). In turn, related work illustrates that these differences give rise to greater sensitivity of EMEs to variation in global risk and the dollar exchange rate (Banerjee et al. 2016; Aoki et al. 2018; Akinci & Queralto 2019). Figure B.1 presents results for a specification extended simultaneously by AE and EME industrial production, consumer prices, NEERs, and policy rates.

The results suggest that the contractionary effects of global risk shocks are very similar across AEs and EMEs. At the same time, the dollar appreciates much more strongly against EME than against AE currencies. And also monetary policy responses are starkly different: While interest rates fall in AEs, they actually rise in EMEs. This is reminiscent of 'fear-of-floating' (Calvo & Reinhart 2002). In particular, in the context of monetary policy spillovers, it has been documented that small open economies tend to mirror core central banks actions' in order to limit exchange rate depreciation despite the ensuing contractionary effects (Ahmed et al. 2021; Corsetti et al. 2021).

The responses of EME interest rates and exchange rates in Figure B.1 are consistent with the analysis of Kalemli-Özcan (2019), namely that the attempt to prevent depreciation against the dollar by tightening monetary policy is self defeating as it induces currency risk premia to rise, eventually resulting in an even larger depreciation.

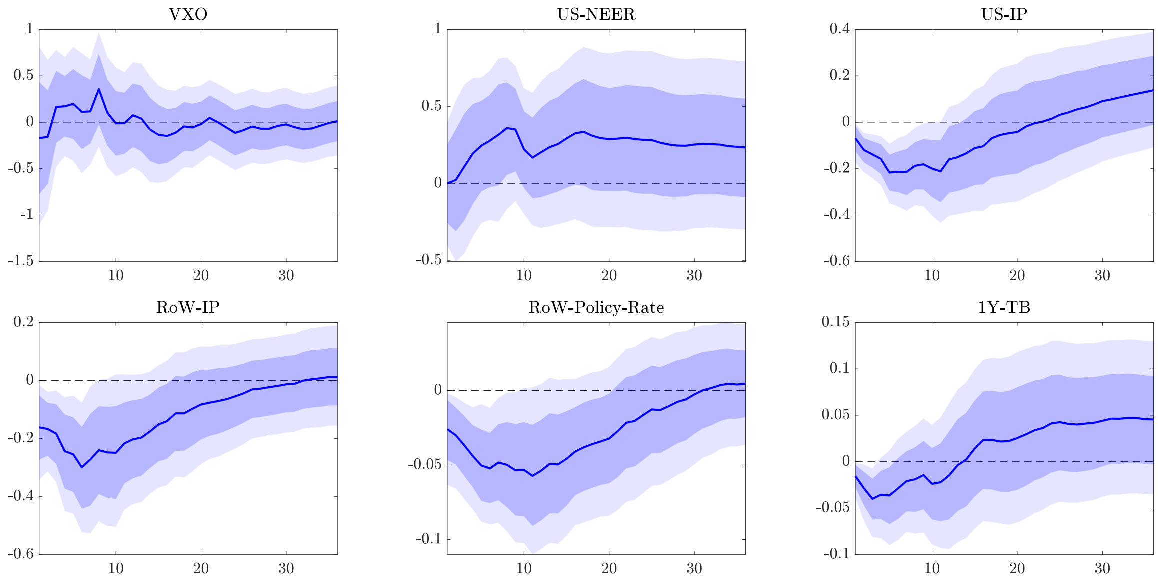
Figure B.1: Impulse responses for AEs and EMEs to a global risk shock



Note: The figure presents the impulse responses to a one-standard deviation global risk shock. Due to the larger dimensionality of the VAR model we use informative Minnesota-type priors and optimal hyperpriors/prior tightness as suggested by Giannone et al. (2015) in the estimation. See also the notes to Figure 2.

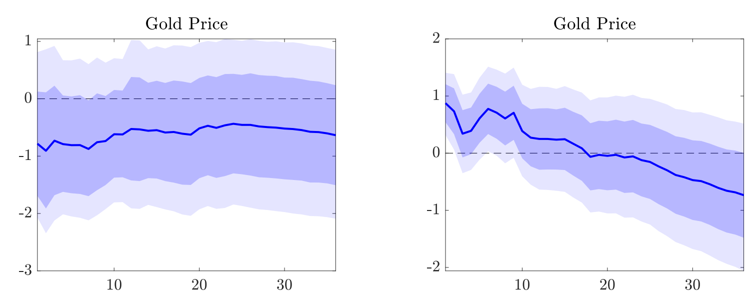
C Online appendix - Additional figures

Figure C.1: Impulse responses to a global demand shock



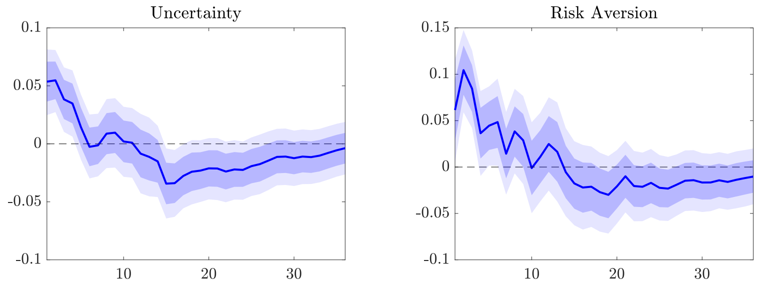
Note: The figure presents the impulse responses to a one-standard deviation global demand shock identified based on sign restrictions. See also the notes to Figure 2. Impulse responses of US CPI and the EBP are omitted to save space.

Figure C.2: Impulse responses of the gold price to global demand (left panel) and global risk (right panel) shocks



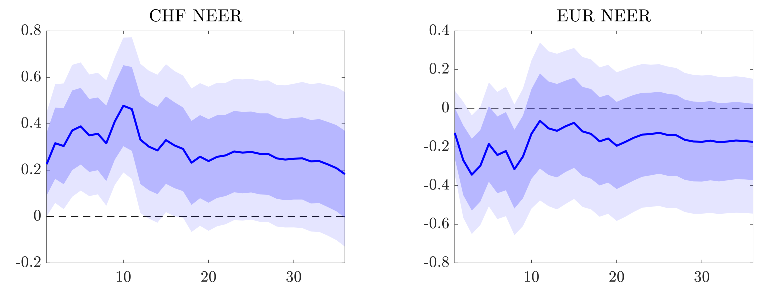
Note: See the notes to Figure C.1.

Figure C.3: Impulse responses of risk aversion and uncertainty a global risk shock



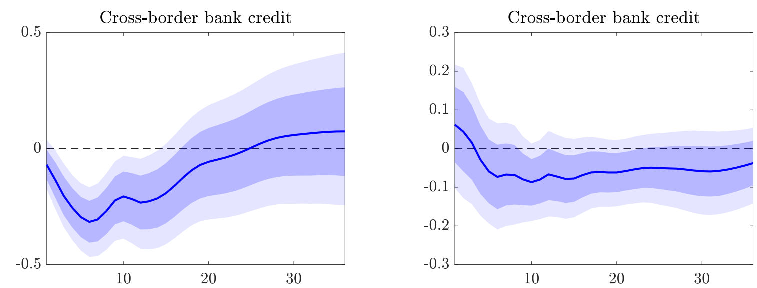
Note: See notes to Figure 2. Responses are obtained from re-estimating the baseline BPSVAR model with the vector $\boldsymbol{y}_t$ augmented with one additional variable at a time.

Figure C.4: Impulse responses of Swiss franc and euro exchange rates to global risk shock



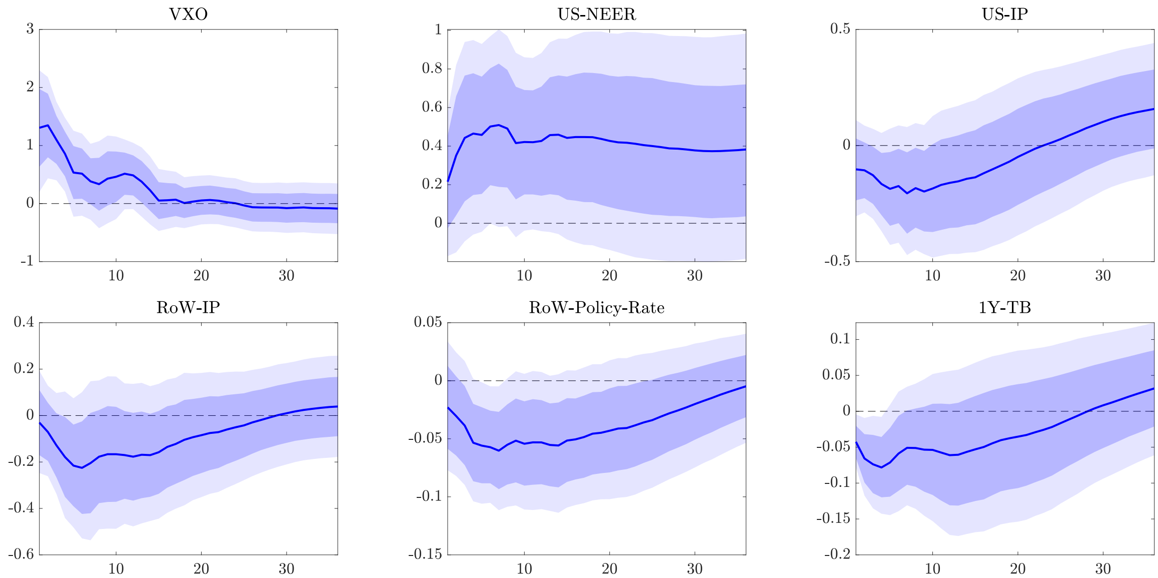
Note: See notes to Figure 2. Responses are obtained from re-estimating the baseline BPSVAR model with the vector $\boldsymbol{y}_t$ augmented with one additional variable at a time.

Figure C.5: Impulse responses of cross-border bank credit and international debt securities to a global risk shock before 2007 (left) and from 2009 (right)



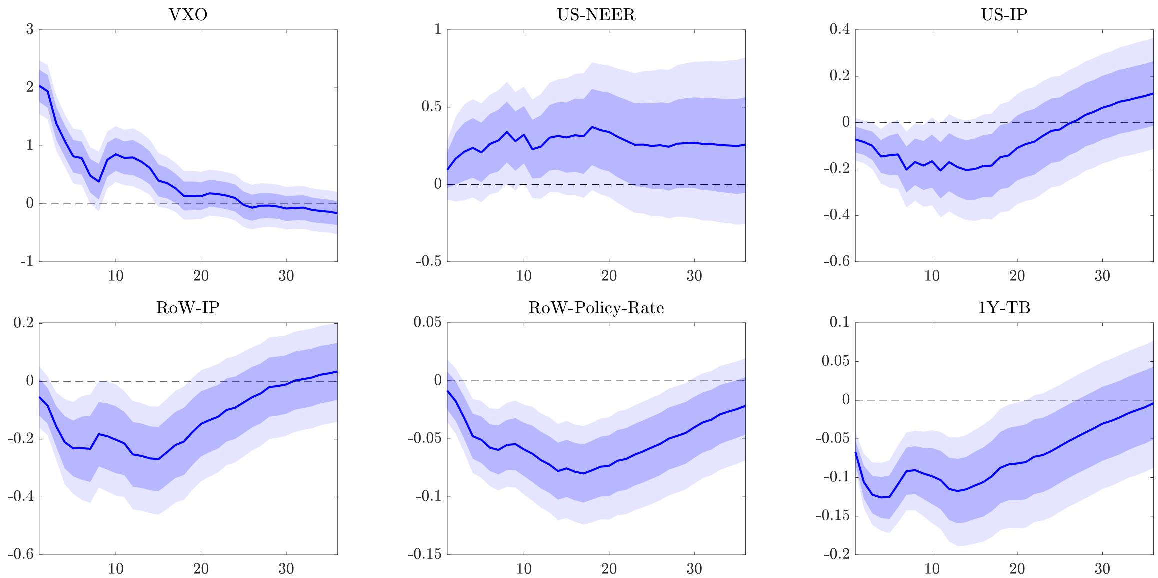
Note: The figure presents the impulse responses of cross-border bank credit to non-US borrowers for the pre and the post-GFC samples. See the notes to Figure 2.

Figure C.6: Impulse responses to a global risk shock when allowing the gold price surprises to be correlated with all structural shocks



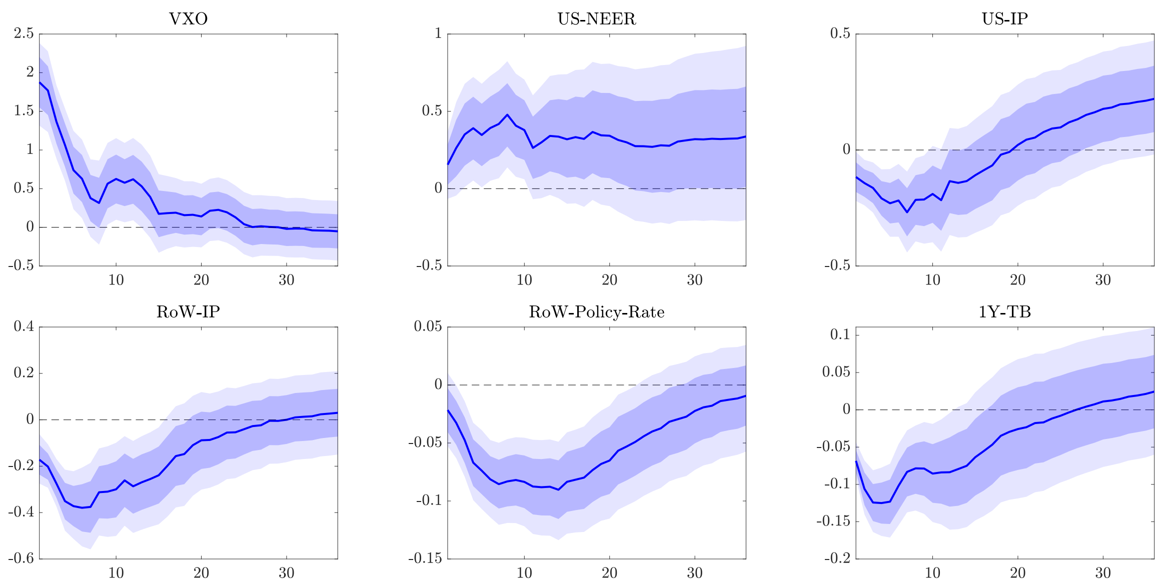
Note: The figure presents the impulse responses to a one-standard deviation global risk shock based on an alternative identification scheme in which the gold price surprises are allowed to be correlated with all structural shocks, imposing only that the correlation is strongest with the global risk shock. See also the notes to Figure 2. Impulse responses of US CPI and the EBP are omitted to save space.

Figure C.7: Impulse responses to a global risk shock when considering intra-daily surprises in 30-year Treasury yields instead of the gold price as proxy variable



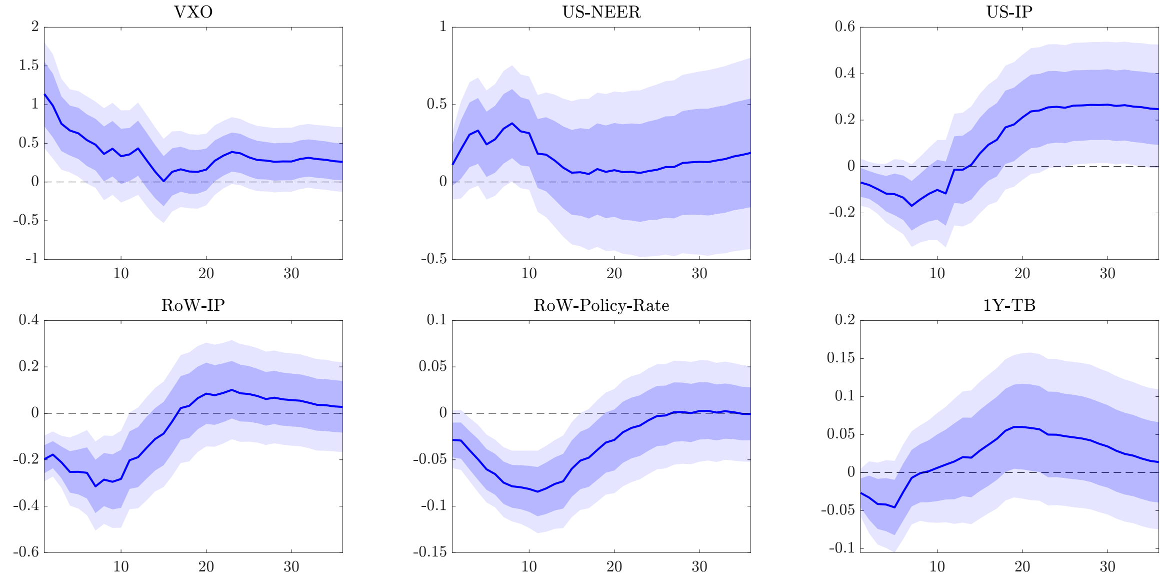
Note: See notes to Figure 5. The results are obtained from a BPSVAR model with intra-daily 30-year Treasury yield surprises as proxy variable. Impulse responses of US CPI and the EBP are omitted to save space.

Figure C.8: Impulse responses to a global risk shock when considering intra-daily surprises in the US dollar-euro exchange rate instead of the gold price as proxy variable



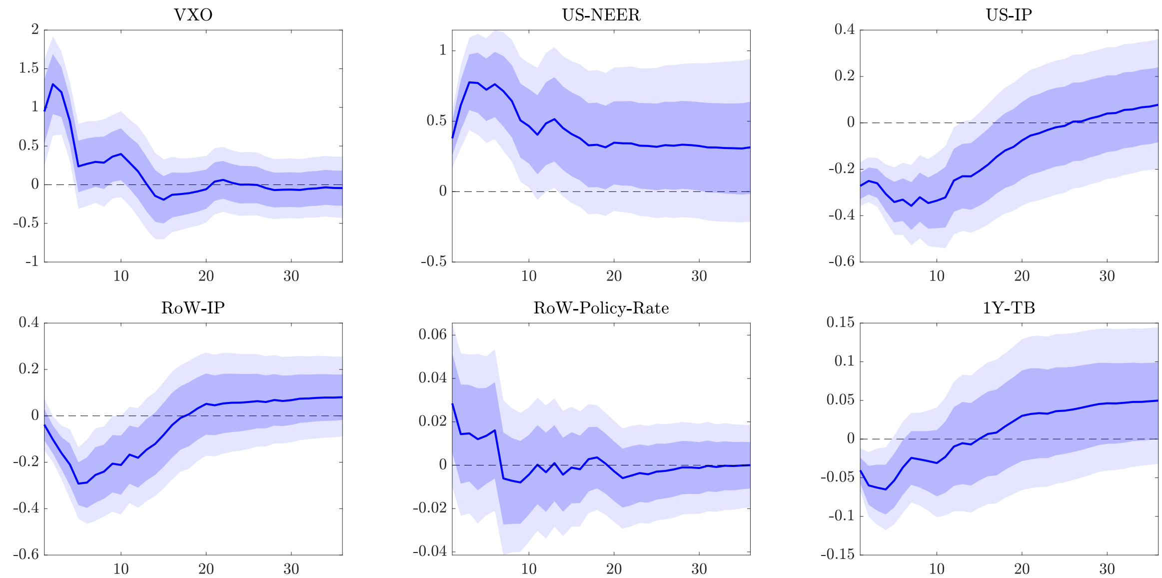
Note: See notes to Figure 5. The results are obtained from a BPSVAR model with intra-daily US dollar-euro exchange rate surprises as proxy variable. Impulse responses of US CPI and the EBP are omitted to save space.

Figure C.9: Impulse responses to a global risk shock when considering changes in the Geopolitical Risk Index of Caldara & Iacoviello (2022) instead of gold price surprises as proxy variable



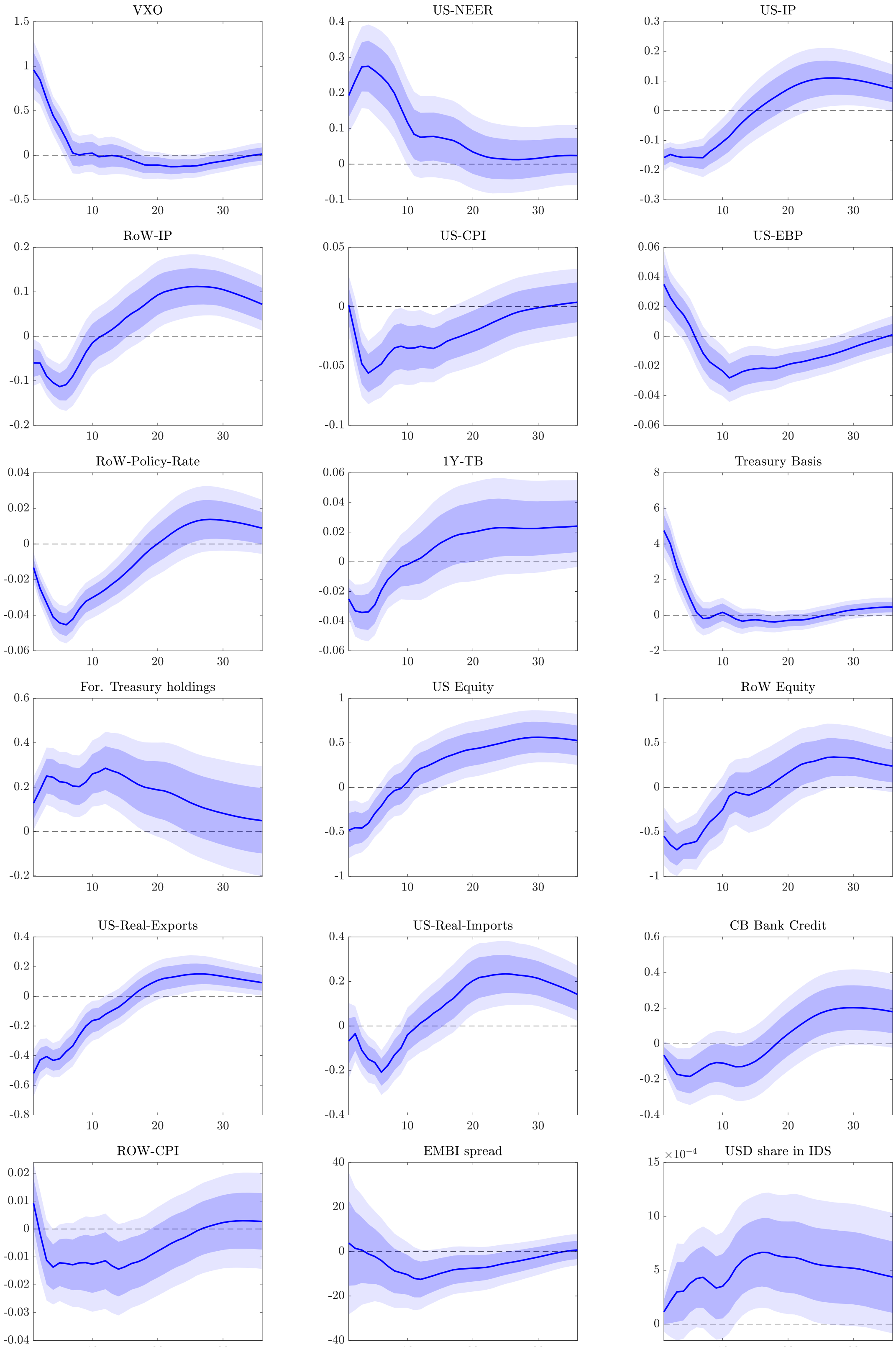
Note: See notes to Figure 5. The results are obtained from a BPSVAR model with monthly changes in the Geopolitical Risk index of Caldara & Iacoviello (2022) as proxy variable. Impulse responses of US CPI and the EBP are omitted to save space.

Figure C.10: Impulse responses to a global risk shock when considering only risk events associated with positive gold price surprises



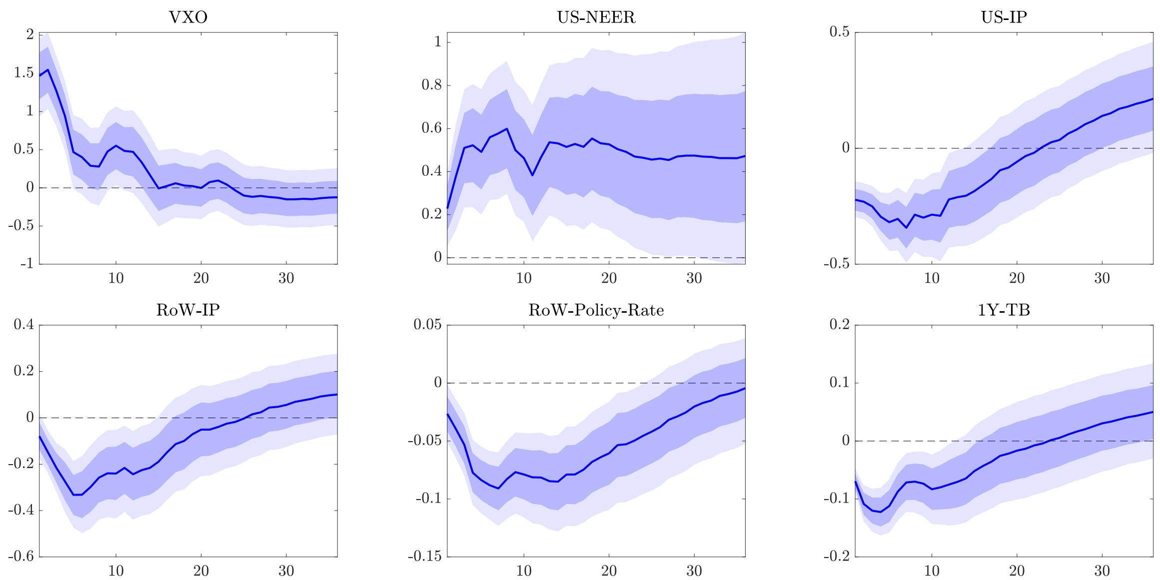
Note: See notes to Figure 5. The results are obtained from a BPSVAR model with only positive gold price surprises. Impulse responses of US CPI and the EBP are omitted to save space.

Figure C.11: Impulse responses to a global risk shock from a large BPSVAR model



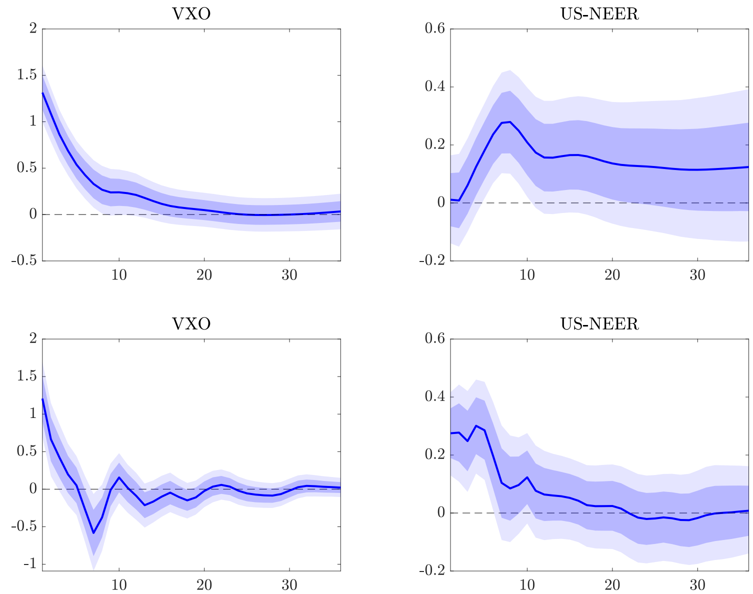
Note: See notes to Figure 5. The model is estimated with informative Minnesota-type priors and optimal hyperpriors/prior tightness as in Giannone et al. (2015). We do not include the liquidity ratio in the VAR model because it is only available for a substantially shorter sample period (see Table D.1).

Figure C.12: Impulse responses to global risk shock when no relevance threshold is imposed



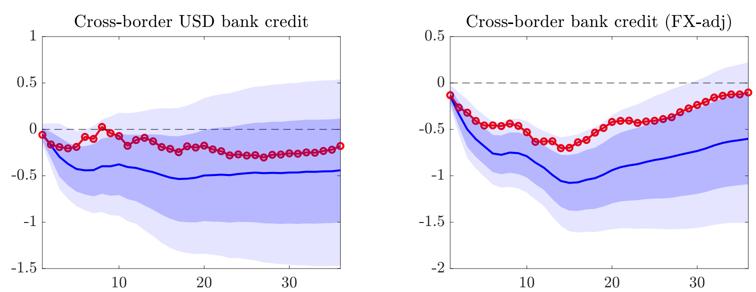
Note: The figure presents the impulse responses to a one-standard deviation global risk shock based on an alternative identification scheme in which we do not impose any relevance threshold. See also the notes to Figure 2. Impulse responses of US CPI and the EBP are omitted to save space.

Figure C.13: Impulse responses to a global risk shock for the sample periods 1990-2006 (top row) and 2007-2019 (bottom row)



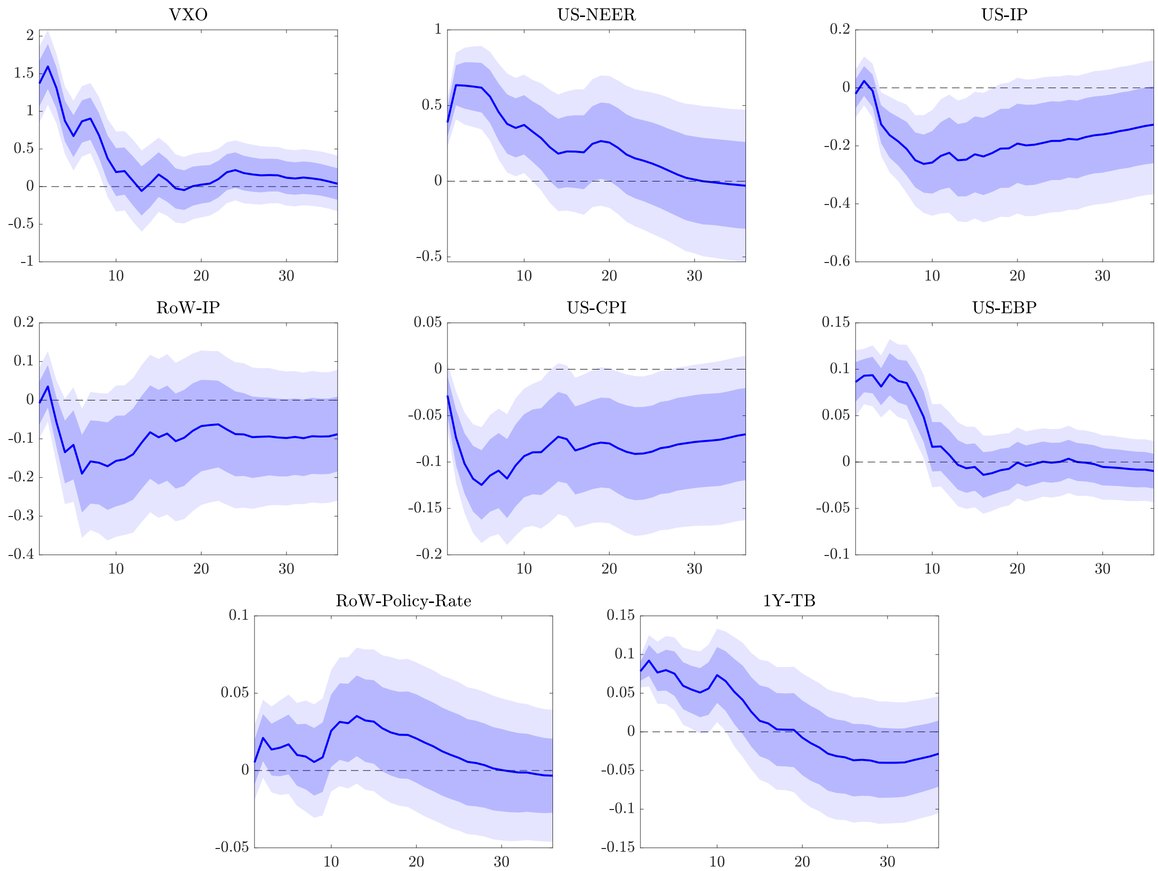
Note: The figure presents the impulse responses to a one-standard deviation global risk shock for the sample periods from 1990-2007 (top row) and 2007 to 2019 (bottom row). Due to the short sample period estimation uses informative Minnesota-type priors and optimal hyperpriors/prior tightness as suggested by Giannone et al. (2015). See also the notes to Figure 2.

Figure C.14: Baseline and MRE-based counterfactual responses of alternative cross-border bank credit variables



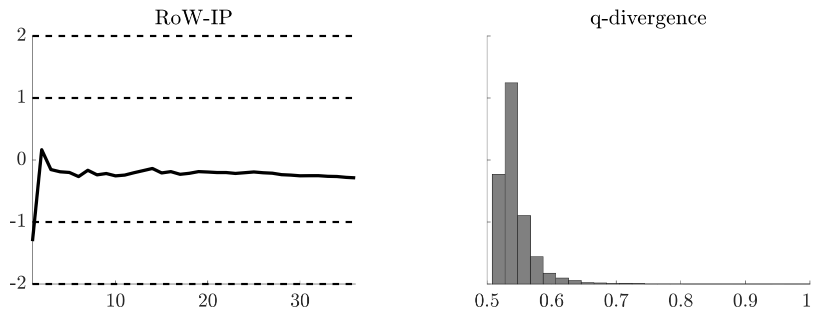
Note: See the notes to Figure 5. The left-hand side panel depicts the responses for US dollar instead of total cross-border bank credit and the right-hand side panel for the exchange rate-adjusted total cross-border bank credit.

Figure C.15: Responses to a contractionary US monetary policy shock



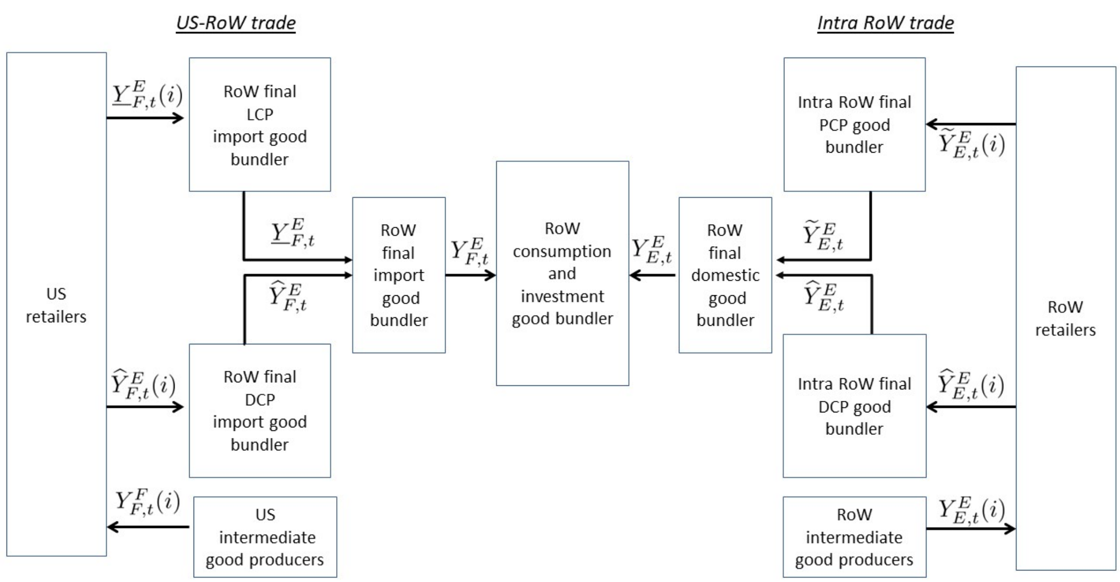
Note: The figure presents the impulse responses to a one-standard deviation US monetary policy shock. See the notes to Figure 2.

Figure C.16: 'Modesty' statistic of Leeper & Zha (2003) and q-divergence of Antolin-Diaz et al. (2021) for the SSC



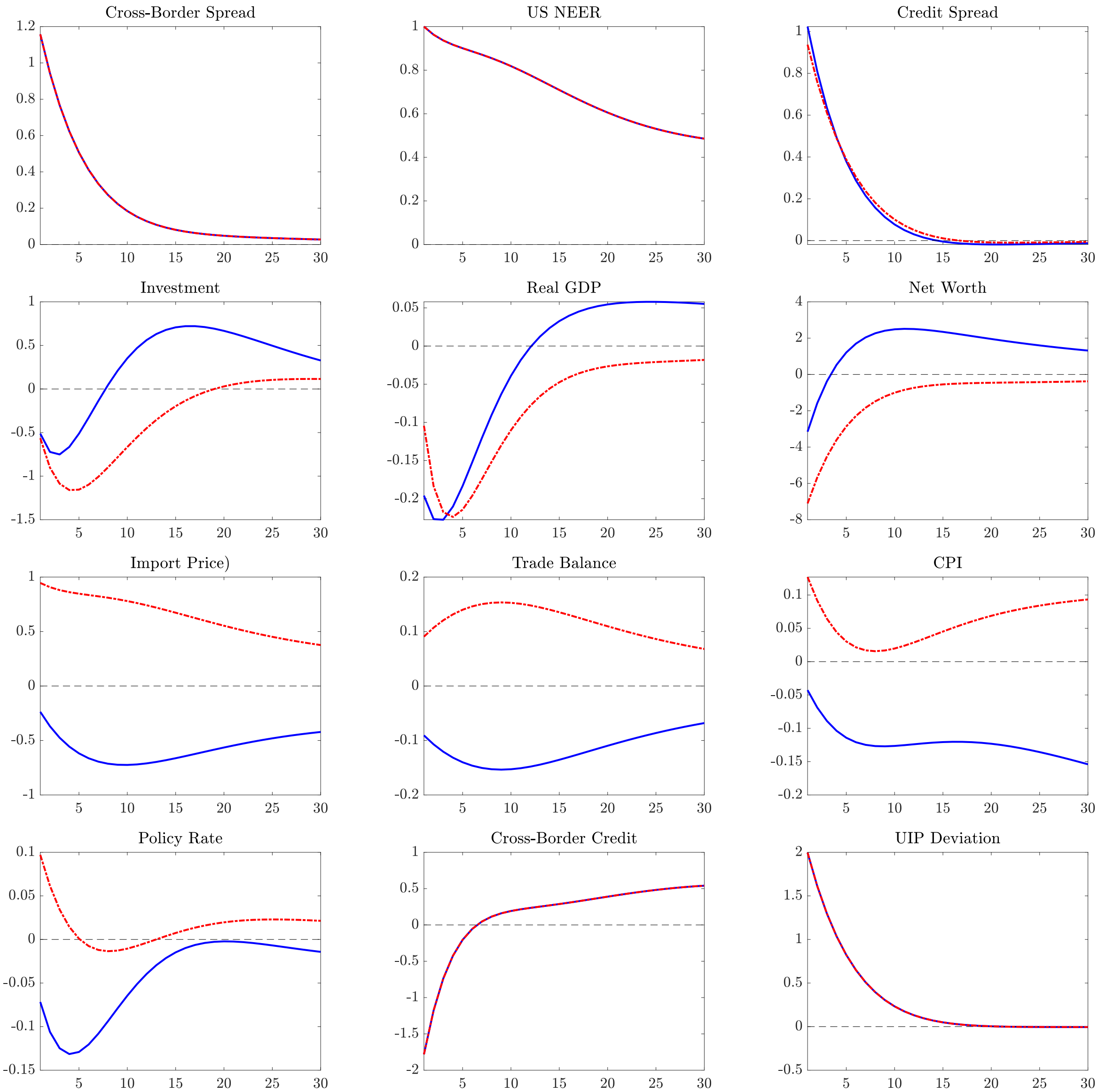
Note: The left-hand side panel shows the 'modesty statistic' of Leeper & Zha (2003) for the required US monetary policy shocks that are needed to impose the counterfactual path of the US dollar NEER (point-wise mean). The offsetting shocks represent 'modest' policy interventions—meaning it would be unlikely to induce agents to adjust their expectations formation—if the statistic is smaller than two in absolute value; the test statistic is distributed as standard normal under the null of 'modest' policy interventions. The right-hand side panel shows the distribution of the q-divergence of Antolin-Diaz et al. (2021). The q-divergence indicates how unlikely a conditional forecast is in terms of comparing the implied distributions of shocks with their unconditional distributions, translated into a comparison of the binomial distributions of a fair and a biased coin.

Figure C.17: Multi-layered production structure in the DCP^2 model for the RoW consumption and investment good



Note: The figure lays out the multi-layered production structure in the DCP^2 model, focusing on the RoW consumption and investment good.

Figure C.18: Impulse responses to a global risk shock in the DCP^2 model for the US (solid blue lines) and the RoW (red dotted lines)



Note: The figure shows the DCP^2 model impulse responses to a shock to the relative 'risk weight' Γ_t creditors of US banks' attach to their cross-border lending activities (see Equations (I.12) and (23)) interpreted as a global risk shock. The blue solid lines depict the impulse responses for the US and the red dash-dotted lines for the RoW. US NEER, investment, real GDP, net worth, import prices, CPI, and cross-border credit are plotted in %-deviations from the steady state, and the remaining variables in absolute percentage-point deviations from the steady state.

D Online appendix - Additional tables

Table D.1: Data description

Variable	Description	Source	Coverage
US 1-year TB rate	1-year Treasury Bill yield at constant maturity	US Treasury/Haver	1990m1 - 2019m12
US IP	Industrial production excl. construction	FRB/Haver	1990m1 - 2019m12
US CPI	US consumer price index	BLS/Haver	1990m1 - 2019m12
US EBP		Favara et al. (2016)	
US dollar NEER	Nominal broad trade-weighted Dollar index	FRB/Haver	1990m1-2019m12
VXO	CBOE market volatility index VXO	Wall Street Journal/Haver	1990m1 - 2019m12
RoW IP	Industrial production, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m12
RoW CPI	Consumer price index	Dallas Fed Global Economic Indicators/Haver (Martínez-García et al. 2015)	1990m1 - 2019m12
RoW policy rate	Short-term official/policy rate, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m12
Yen, euro, Swiss franc, British pound NEER	Nominal broad effective exchange rate	J.P. Morgan/Haver	1990m1-2019m12
US real exports	Exports of goods and services (chnd. 2012$)	BEA/Haver	1990q1-2019q2, interpolated to monthly frequency
US real imports	Imports of goods and services (chnd. 2012$)	BEA/Haver	1990q1-2019q2, interpolated to monthly frequency
Non-US USD cross-border bank credit	Banks' external liabilities in USD of banks owned by the world less externalliabilities in USD of banks owned by US nationals	BIS Locational Banking Statistics, Table A7/Haver	1990q1-2019q2, interpolated to monthly frequency
Non-US non-USD cross-border bank credit	Banks' external liabilities in non-USD of banks owned by the world lessexternal liabilities in non-USD of banks owned by US nationals	BIS Locational Banking Statistics, Table A7/Haver	1990q1-2019q2, interpolated to monthly frequency
EMBI spread	EMBI Brady bonds sovereign spread	JP Morgan Emerging Markets Bond Indexes /Haver	1990m1-2019m12
International debt securities	Debt securities issued outside of the resident's home market	BIS International Debt Issuance Statistics/Haver	1990q1-2019q4, interpolated to monthly frequency
AE and EME IP	Industrial production, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m12
AE and EME CPI	Consumer price index, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m12
AE and EME policy rate	Short-term official/policy rate, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m12
US dollar AE NEER	Nominal broad trade-weighted Dollar index against AEs	FRB/Haver	1990m1-2019m12
US dollar EME NEER	Nominal broad trade-weighted Dollar index against EMEs	FRB/Haver	1990m1-2019m12
US Treasury premium	Defined as the deviation from covered interest parity between US and G10 government bond yields	Du et al. (2018)	1991m4-2019m12
Foreign Treasury security holdings		Treasury International Capital (TIC) System/Haver	1990q1-2000q1, 2000m1-2019m12, interpolated to monthly frequency for 1990m1-2000m2
Commercial banks' Treasury and agency securities	Used for calculation of liquidity ratio	FRB/Haver	1990m1-2019m12
Total reserve balances with Federal Reserve banks	Used for calculation of liquidity ratio	FRB/Haver	1990m1-2019m12
Total demand deposits	Used for calculation of liquidity ratio	FRB/Haver	1990m1-2019m12
Financial commercial paper outstanding	Used for calculation of liquidity ratio	FRB/Haver	2001m1-2019m12
S&P 500	S&P 500 Composite	S&P/Haver	1990m1 - 2019m12
MSCI World excl. US	MSCI world excluding US	MSCI/Bloomberg	1990m1 - 2019m12
Risk aversion		Bekaert et al. (forthcoming)	1990m1 - 2019m12
Uncertainty		Bekaert et al. (forthcoming)	1990m1 - 2019m12
Global factor in risky asset prices		Miranda-Agrippino et al. (2020)	1990m1 - 2019m4
Global factor in capital flows		Miranda-Agrippino et al. (2020)	1990q1 - 2018q3, interpolated to monthly frequency

Notes: BLS stands for Bureau of Labour Statistics, FRB for Federal Reserve Board, BEA for Bureau of Economic Analysis, and BIS for Bank for International Settlements.

E Online appendix - Implementation of the MRE approach

The posterior distribution of the impulse responses $f(\cdot)$ is approximated by N draws obtained from a Bayesian estimation algorithm. Following the importance sampling procedure of Arias et al. (2018, 2021), the re-sampled draws from the BPSVAR for $\boldsymbol{y}_{T+1,T+h}$ constitute an unweighted and independent sample from the posterior distribution $f(\cdot)$ and as such are assigned a weight of $w_i = 1/N$, $i = 1, 2, \ldots, N$. The counterfactual posterior distribution $f^\star(\cdot)$ can be approximated by assigning different weights w_i^* to the draws from the baseline posterior.

The relative entropy (or distance) between the approximated posterior distributions is measured by

$$\mathcal{D}(f^*, f) = \sum_{i=1}^{N} w_i^\star log\left(\frac{w_i^\star}{w_i}\right). \tag{E.1}$$

The goal of the MRE approach is to determine the counterfactual weights $\boldsymbol{w}^*$ that minimise $\mathcal{D}(\cdot)$ subject to

$$w_i^\star \geq 0, \qquad \forall\, i = 1, 2, ..., N, \tag{E.2}$$

$$\sum_{i=1}^{N} w_i^* = 1, \tag{E.3}$$

$$\sum_{i=1}^{N} w_i^* g(\boldsymbol{y}_{T+1,T+h}^{(i)}) = \bar{\boldsymbol{g}}, \tag{E.4}$$

where $\boldsymbol{y}_{T+1,T+h}^{(i)}$ are the impulse responses to a global risk shock shock as defined in Section 4. Equations (E.2) and (E.3) reflect that the weights are probabilities, and Equation (E.4) that the counterfactual posterior distribution shall satisfy some constraint.

In particular, in our application for Equation (E.4) we have

$$\sum_{i=1}^{N} y_{\$,T+h}^{(i)} w_{i,h}^* = 0, \tag{E.5}$$

where $y_{\$,T+h}^{(i)}$ the impulse response of dollar exchange rate to a global risk shock at horizon h associated with the i-th draw. Notice that—consistent with the baseline posterior for which we report point-wise means in Figure 2 and elsewhere in the paper as well as with Giacomini & Ragusa (2014)—we apply the MRE approach separately at each impulse response horizon $T + 1, T + 2, \ldots, T + h$.

As shown by Robertson et al. (2005) and Giacomini & Ragusa (2014), the weights of the counterfactual posterior distribution $\boldsymbol{w}_h^*$ can be obtained numerically by tilting the weights of the baseline posterior distribution $\boldsymbol{w}_h$ using the method of Lagrange. In particular, the weights of the counterfactual posterior distribution are given by

$$w^*_{i,h} = \frac{w_{i,h} \exp\left[\lambda_h g(y^{(i)}_{\$,T+h})\right]}{\sum\limits_{i=1}^{N} w_{i,h} \exp\left[\lambda_h g(y^{(i)}_{\$,T+h})\right]}, \quad i = 1, 2, \ldots, N \tag{E.6}$$

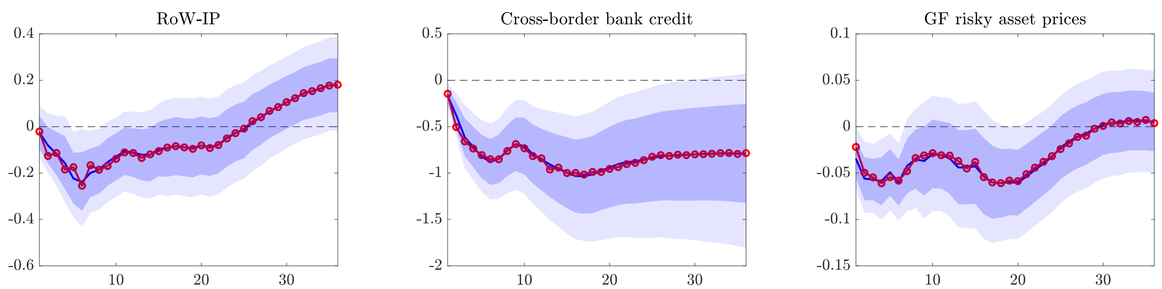
where λ_h is the Lagrange multiplier associated with the constraint $g(y^{(i)}_{\$,T+h}) = y^{(i)}_{\$,T+h} = 0$. It can be shown that the Lagrange multiplier can be obtained numerically as

$$\lambda_h = \arg\min_{\tilde{\lambda}_h} \sum_{i=1}^{N} w_{i,h} \exp\left\{\tilde{\lambda}_h \left[g(y^{(i)}_{\$,T+h})\right]\right\}. \tag{E.7}$$

F Online Appendix - The special role of the dollar exchange rate

As shown in Figure 3 global risk shocks not only appreciate the dollar but also other safe-haven currencies. And yet the dollar's role for the transmission is special. We illustrate this with a counterfactual in which we preclude an appreciation of the Japanese yen while imposing that the dollar appreciates as in the baseline. Figure F.1 shows that the responses of RoW industrial production, cross-border bank credit and the global factor in risky asset prices of Miranda-Agrippino & Rey (2020) to a global risk shock for this counterfactual are unchanged relative to the baseline. One explanation for this result is the uniqueness of US safe assets and the minuscule role of yen-denominated in global cross-border bank credit. Hence, the appreciation of the yen does not shape the transmission of global risk shocks to the RoW in the way the dollar does.

Figure F.1: Baseline and MRE-based counterfactual responses when Japanese yen instead of US dollar is unresponsive to a global risk shock



Note: See the notes to Figure 5. The red dotted lines depict the responses of rest-of-the-world industrial production (first panel), cross-border bank credit (middle panel) and the global factor in risky asset prices of Miranda-Agrippino & Rey (2020) in the counterfactual in which the Japanese yen is constrained to not respond to the global risk shock, while the response of the dollar is constrained to be identical to the baseline.

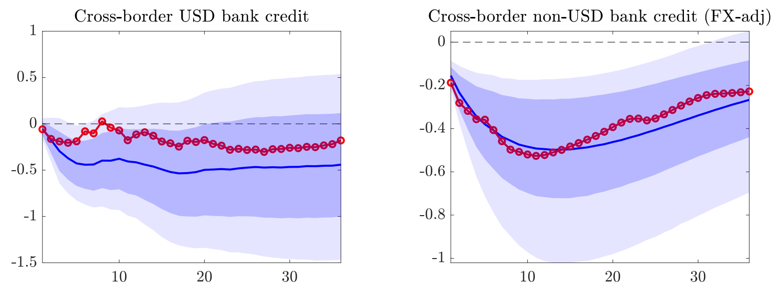
G Online Appendix - The special role of dollar-denominated cross-border bank credit

Ivashina et al. (2015) present a model in which globally active banks cut dollar lending more than euro lending in response to a shock to their credit quality. In particular, motivated by the data, in their model globally active banks raise *unsecured* dollar funding through wholesale markets in the US and euro funding through *insured* retail deposits in Europe. In this setting, the supply of dollar funding is more sensitive to credit quality—for example global risk—shocks than the supply of euro funding. In principle, banks could borrow in euros and swap them into dollars to make up for the dollar funding shortfall in response to a credit quality shock, but this is precluded by deviations from covered interest parity (CIP) when there is limited capital to take the other side of the swap trade. As a result, a credit quality shock induces globally active banks to cut dollar lending more than euro lending. Indeed, Avdjiev, Du, et al. (2019) document a 'triangular' relationship in that a (i) stronger dollar goes hand in hand with (ii) larger CIP deviations and (iii) contractions of dollar cross-border bank credit.

In the context of our paper, dollar appreciation in response to a global risk shock constitutes a credit quality shock as it deteriorates the net worth of borrowers with currency mismatches. The prediction that emerges against the background of the model of Ivashina et al. (2015) and the segmentation of funding markets across currencies due to CIP deviations is that dollar-denominated cross-border bank credit should drop more in response to a global risk shock than euro-denominated cross-border bank credit, as well as cross-border bank credit in other currencies such as in Japanese yen, for which funding also stems from secured deposits.

To test this prediction Figure G.1 shows effects of global risks shocks in the baseline and the counterfactual separately for dollar and non-dollar-denominated, exchange rate adjusted cross-border bank credit. Our findings are consistent with the prediction from the model of Ivashina et al. (2015) and the findings in Avdjiev, Du, et al. (2019). In particular, the weakening of the effect of a global risk shock in the counterfactual is much larger for dollar than for non-dollar-denominated cross-border bank credit.

Figure G.1: Baseline and MRE counterfactual responses of dollar and non-dollar/exchange rate-adjusted cross-border credit to a global risk shock



Note: See the notes to Figure 5. Because the data are only available from 2002, the BPSVAR model with non-dollar/exchange rate-adjusted cross-border credit is estimated with informative Minnesota-type priors and optimal hyperpriors/prior tightness as suggested by Giannone et al. (2015) and—to obtain a stable model—six instead of twelve lags.

H Online appendix - Implementation of structural shock counterfactuals

Iterate forward the VAR model in Equation (1) to obtain

$$\boldsymbol{y}_{T+1,T+h} = \boldsymbol{b}_{T+1,T+h} + \boldsymbol{M}'\boldsymbol{\epsilon}_{T+1,T+h}, \tag{H.1}$$

where $\boldsymbol{b}_{T+1,T+h}$ represents the autoregressive component of the system that is due to initial conditions as of period T, and the $nh \times nh$ matrix $\boldsymbol{M}$ the effects of the structural shocks; $\boldsymbol{M}$ is a function of the structural VAR parameters $\boldsymbol{\psi} \equiv vec(\boldsymbol{A}_0, \boldsymbol{A}_1)$. Assume again for simplicity of exposition but without loss of generality that the VAR model is stationary and in steady state in period T so that $\boldsymbol{b}_{T+1,T+h} = \boldsymbol{0}$. In this setting, the impulse response to a global risk shock again coincides with the forecast $\widetilde{\boldsymbol{y}}_{T+1,T+h}$ conditional on $\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}$ with $\widetilde{\epsilon}^u_{T+1} = 1$, $\widetilde{\epsilon}^u_{T+s} = 0$ for $s > 1$ and $\widetilde{\epsilon}^\ell_{T+s} = 0$ for $s > 0$, $\ell \neq u$. In contrast to the MRE-based counterfactual, under SSC the implied counterfactual VAR model as reflected in $\boldsymbol{M}$ in Equation (H.1) is unchanged relative to the baseline. Instead, in order for the impulse response $\widetilde{\boldsymbol{y}}_{T+1,T+h}$ to satisfy the counterfactual constraints additional shocks in $\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}$ are allowed to materialize over periods $T+1, T+2, \ldots, T+h$. Intuitively, these additional shocks are chosen such that they offset the effect of the global risk shock on the dollar exchange rate.

Antolin-Diaz et al. (2021) describe how to implement SSC in terms of a conditional forecast $\widetilde{\boldsymbol{y}}_{T+1,T+h}$ with constraints on the paths of the endogenous variables represented by

$$\overline{\boldsymbol{C}}\widetilde{\boldsymbol{y}}_{T+1,T+h} = \overline{\boldsymbol{C}}\boldsymbol{M}'\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\overline{\boldsymbol{f}}_{T+1,T+h}, \overline{\boldsymbol{\Omega}}_f), \tag{H.2}$$

where $\overline{\boldsymbol{C}}$ is a $k_o \times nh$ selection matrix, $\overline{\boldsymbol{f}}_{T+1,T+h}$ is a $k_o \times 1$ vector and $\overline{\boldsymbol{\Omega}}_f$ a $k_o \times k_o$ matrix, as well as constraints on the structural shocks represented by

$$\boldsymbol{\Xi}\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\boldsymbol{g}_{T+1,T+h}, \boldsymbol{\Omega}_g), \tag{H.3}$$

where $\boldsymbol{\Xi}$ is a $k_s \times nh$ selection matrix, $\boldsymbol{g}_{T+1,T+h}$ a $k_s \times 1$ vector, and $\boldsymbol{\Omega}_g$ a $k_s \times k_s$ matrix.[28] Antolin-Diaz et al. (2021) show how to obtain the SSA solution

$$\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\boldsymbol{\mu}_\epsilon, \boldsymbol{\Sigma}_\epsilon), \tag{H.4}$$

that satisfies the counterfactual constraint in Equation (H.2) and the constraint on the structural shocks in Equation (H.3). The SSC impulse response is then given by $\widetilde{\boldsymbol{y}}_{T+1,T+h} = \boldsymbol{M}'\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}$.

[28] For example, if the dollar exchange rate is ordered last in the vector of endogenous variables $\boldsymbol{y}_t$ in the VAR model and we impose in the counterfactual that it remains at its baseline along the impulse response horizon without any uncertainty, then in Equation (H.2) we have

$$\overline{\boldsymbol{C}} = \boldsymbol{I}_h \otimes \boldsymbol{e}_n', \quad \overline{\boldsymbol{f}}_{T+1,T+h} = \boldsymbol{0}_{h\times 1}, \quad \overline{\boldsymbol{\Omega}}_f = \boldsymbol{0}_{h\times h},$$

where $\boldsymbol{e}_i$ is $n \times 1$ denotes vector of zeros with unity at the i-th position.

I Additional model details

I.1 Households and unions

In each period a household consumes a non-traded final good subject to habit formation in consumption. Furthermore each households is a monopolistic supplier of a differentiated labor service $L_{E,t}(h)$ and sells this to a perfectly competitive union that transforms it into an aggregate labor supply using a constant elasticity of substitution (CES) technology. Households satisfy demand for labor given the wage rate $W_{E,t}$, with wage setting being subject to frictions à la Calvo. The period-by-period utility function is given by

$$U(C_{E,t}, L_{E,t}) = \frac{1}{1-\sigma^c}(C_{E,t} - h_E C_{E,t-1})^{1-\sigma^c} - \frac{\kappa_{E,w}}{1+\varphi} L_{E,t}^{1+\varphi}. \tag{I.1}$$

with $\sigma^c, \varphi, h_E, \kappa_{E,w}$ as the intertemporal elasticity of substitution, the inverse Frisch elasticity of labor supply, the habit formation parameter and an exogenous labor scale parameter respectively. Households maximize utility subject to the following budget constraint

$$\frac{B^n_{E,t}}{P^C_{E,t}} + C_{E,t} = \frac{B^n_{E,t-1} R_{E,t-1}}{P^C_{E,t}} + \frac{W_{E,t}(h) L_{E,t}(h) + IS_{E,t}(h)}{P^C_{E,t}} + \frac{\Pi^C_{E,t}}{P^C_{E,t}} + \frac{\Pi^R_{E,t}}{P^C_{E,t}},$$

where we chose the final consumption and investment good price $P^C_{E,t}$ as the numeraire. $R_{E,t-1}$ is the predetermined domestic risk-free rate paid on nominal deposits with domestic banks $B^n_{E,t}$. $IS_{E,t}$ furthermore denotes an income stream from domestic state-contingent securities ensuring that all households will choose the same consumption and savings plans, despite temporarily receiving different wages due to the assumption of Calvo-type wage setting. Lastly $\Pi^C_{E,t}$ and $\Pi^R_{E,t}$ represent nominal profits from domestic (RoW) capital producing and retail firms respectively. The first-order condition of the household with respect to the choice of consumption is given by

$$\Lambda_{E,t} = (C_{E,t} - h_E C_{E,t-1})^{-\sigma_c} - \beta h_E \mathbb{E}_t[(C_{E,t+1} - h C_{E,t})^{-\sigma_c}] \tag{I.2}$$

with $\Lambda_{E,t}$ as the marginal utility of consumption. The intertemporal optimality conditions for the individual holdings of deposits with the local bank reads as

$$\Lambda_{E,t} = \mathbb{E}_t\Big[\beta \Lambda_{E,t+1} \frac{R_{E,t}}{1+\pi^C_{E,t+1}}\Big]. \tag{I.3}$$

where $\pi^C_{E,t+1}$ corresponds to the net inflation rate of the final consumption good. The working part of the household also sells its differentiated labor services $L_{E,t}(h)$ to a competitive union, which combines the differentiated labor services into a composite labor good using CES technology. Lastly the union leases the combined labor service to the intermediate good firms at the aggregate nominal wage rate $W_{E,t}$. The worker optimally chooses its wage given labor demand by the union taking into account that wage setting is subject to frictions à la Calvo, meaning that in each period they face a constant probability $(1-\theta_w)$ of being able to adjust their nominal wage. As such the aggregate real wage index evolves as

$$w_{E,t}^{1-\psi_w} = (1-\theta_w)\tilde{w}_{E,t}^{1-\psi_w} + \theta_w(1+\pi_{E,t}^C)^{\psi_w - 1} w_{E,t-1}^{1-\psi_w} \tag{I.4}$$

with $\tilde{w}_{E,t}$ as the optimal reset wage and $w_{E,t}$ as the economy wide real wage.

I.2 RoW financial intermediaries

Recall that the objective of the banker is to maximize its expected terminal wealth

$$V_{E,j,t} = max\ \mathbb{E}_t \sum_{s=0}^{\infty} (1-\theta_B)\Theta_{E_{t,t+s}}(N_{E,j,t+1+s}). \tag{I.5}$$

subject to the incentive compatibility constraint

$$V_{E,j,t} \geq \delta_{E,B,t}(Q_{E,j,t}S_{E,j,t}). \tag{I.6}$$

It can be shown that the value function of a bank is linear in its components and applying a guess and verify procedure the solution to the bankers problem can be characterized by the following set of equations.

$$v_{E,t} = \mathbb{E}_t\Big(\Omega_{E,t,t+1}(R_{E,k,t+1} - R_{E,t})\Big) \tag{I.7}$$

$$_{E,t} = \mathbb{E}_t\Big(\Omega_{E,t,t+1}(\Big[\frac{\mathcal{E}_{E,t+1}^F}{\mathcal{E}_{E,t}^F}\Big]^{dl} R_{E,b,t}^F - R_{E,t})\Big) \tag{I.8}$$

$$n_{E,t} = \mathbb{E}_t\Big(\Omega_{E,t,t+1}R_{E,t}\Big) \tag{I.9}$$

$$Q_{E,t}S_{E,j,t} = \frac{n_{E,t}}{\delta_{E,B,t} - (v_{E,t} - u_{E,t}\Xi_{E,j,t}^F)} N_{E,j,t} = \phi_{E,B,t}N_{E,j,t}. \tag{I.10}$$

$$\Omega_{E,t,t+1} = \mathbb{E}_t\Big(\frac{\Theta_{E,t,t+1}}{(1+\pi_{E,t+1}^c)}\Big[(1-\theta_B) + \theta_B((v_{E,t+1} - u_{E,t+1}\Xi_{E,j,t+1}^F)\phi_{E,B,t+1} + n_{E,t+1})\Big]\Big) \tag{I.11}$$

Equations I.7, I.8, I.9 represent the discounted excess returns from borrowing and lending domestically, the discounted excess costs of borrowing in US-$ instead of acquiring domestic deposits and the discounted marginal value of an additional unit of equity. Equation I.11 is the bankers "augmented" real stochastic discount factor, which accounts for marginal value of funds internal to the financial intermediary and the fact that the bank may have to close with a probability of $1-\theta_B$. Lastly I.10 shows that total lending is restricted to be a multiple of existing net worth, with $\phi_{E,B,t}$ as the optimal leverage ratio, which is common across all RoW banks. It is increasing in the marginal value of equity as a rise in $n_{E,t}$ causes the bank to be more profitable in the future, thereby raising $V_{E,j,t}$ and thus the value of $Q_{E,t}S_{E,j,t}$ a bank can acquire before reaching the constraint outlined in equation I.6. It is furthermore increasing in the total excess returns from domestic lending $(v_{E,t} - u_{E,t}\Xi_{E,j,t}^F)$, where the fraction $\Xi_{E,j,t}^F$ accounts for the fact that lending to domestic firms is partly financed by US-$ loans obtained from US banks. Lastly the leverage ratio falls in the fraction of assets that the banker can divert $\delta_{E,B,t}$, as the moral hazard problem becomes more severe and the incentive constraint in equation I.6 tightens. While the equations above pin down the optimal leverage ratio and thereby the total amount of assets of the banker as well as the solutions to the value function of the bank, the optimal liability composition is yet to be determined.

I.3 US financial intermediaries

Recall that the creditors of a US bank require that the expected terminal wealth of the banker j satisfies

$$V_{F,j,t} \geq \delta_{F,B}(Q_{F,t}S_{F,j,t} + \Gamma_t B^{F\star}_{E,j,t}). \tag{I.12}$$

Defining $\xi^F_{E,j,t} = \frac{B^{F\star}_{E,j,t}}{RER^{F(1-dl)}_{E,t} Q_{F,t}S_{F,j,t}}$ as the asset ratio of interbank loans to domestic investments it can be shown that the coefficients of the value function of the banker $V_{F,j,t}$ are given

$$v_{F,t} = \mathbb{E}_t\Big(\Omega_{F,t,t+1}(R_{k,F,t+1} - R_{F,t})\Big) \tag{I.13}$$

$$u^F_{E,b,t} = \mathbb{E}_t\Big(\Omega_{F,t,t+1}(\Big[\frac{\mathcal{E}^F_{E,t}}{\mathcal{E}^F_{E,t+1}}\Big]^{1-dl} R^F_{E,b,t} - R_{F,t})\Big) \tag{I.14}$$

$$n_{F,t} = \mathbb{E}_t\Big(\Omega_{F,t,t+1}(R_{F,t})\Big) \tag{I.15}$$

$$(1 + \Gamma_t\xi^F_{E,j,t})Q_{F,t}S_{F,j,t} = \frac{n_{F,t}}{\delta_{F,B} - v_{F,t}} = \phi_{F,B,t}N_{F,j,t} \tag{I.16}$$

$$\Omega_{F,t,t+1} = \Big(\frac{\Theta_{F,t,t+1}}{(1+\pi^c_{F,t+1})}\Big[(1-\theta^F_B) + \theta^F_B\Big(\frac{v_{F,t+1} + u^F_{E,b,t+1}\xi^F_{E,j,t+1}}{1+\Gamma_t\xi^F_{E,j,t+1}}\phi_{F,B,t+1} + n_{F,t+1}\Big)\Big]\Big). \tag{I.17}$$

While $v_{F,t}, n_{F,t}$ and $\Omega_{F,t,t+1}$ are slightly different versions of their RoW counterparts touched up on the previous section, $u^F_{E,b,t}$ and $\phi_{F,B,t}$ need a bit of elaboration. In particular $u^F_{E,b,t}$ represents the discounted excess return from cross-border interbank lending. Furthermore the optimal leverage ratio $\phi_{F,B,t}$ takes into account the fact that the relative "risk weight" Γ_t, which depositors attach to the different kinds of assets, may be different from one. Lastly it can be shown that the optimal portfolio choice of the US bank imposes a no-arbitrage relation between the returns on cross-border interbank lending and the returns on domestic activities given by

$$\Gamma_t\mathbb{E}_t\Big(\Omega_{F,t,t+1}(R_{k,F,t+1} - R_{F,t})\Big) = \mathbb{E}_t\Big(\Omega_{F,t,t+1}(\Big[\frac{\mathcal{E}^F_{E,t}}{\mathcal{E}^F_{E,t+1}}\Big]^{1-dl} R^F_{E,b,t} - R_{F,t})\Big). \tag{I.18}$$

I.3.1 Intermediate good firms

In each economy there exists a continuum of perfectly competitive intermediate goods firms that sell their output to domestic retailers. We assume that at the end of period t but before the realization of shocks the intermediate good firm acquires capital for use in next period's production. To do so, the intermediate good firm i issues $S_{E,i,t}$ claims equal to the number $K_{E,i,t}$ of units of capital acquired, and prices each claim at the real price of a unit of capital $Q_{E,t}$.[29] The production function is

$$Z_{E,i,t} = \Big(U_{E,i,t}K_{E,i,t-1}\Big)^{\alpha} L_{E,i,t}^{(1-\alpha)}, \tag{I.19}$$

with $Z_{E,i,t}$ the amount of output produced by the individual RoW intermediate good firm in period t, $L_{E,i,t}$ the labor used in production, and $U_{E,i,t}$ the employed utilization rate of capital.

Cost minimization yields the standard equations for the optimal amount of production inputs

$$MC^{r}_{E,t} = \frac{w_{E,t}^{1-\alpha}\tau_{E,t}(U_{E,t})'^{\alpha}}{(1-\alpha)^{(1-\alpha)}\alpha^{\alpha}}. \tag{I.20}$$

$$\frac{w_{E,t}}{\tau_{E,t}(U_{E,t})'} = \frac{1-\alpha}{\alpha}\frac{(U_{E,t}K_{E,t-1})}{L_{E,t}}, \tag{I.21}$$

where $MC^{r}_{E,t}$ denote the real marginal costs of the intermediate good firms deflated by the RoW final good price $P^{C}_{E,t}$ and $\tau_{E,t}(U_{E,t})'$ as the derivative of the adjustment cost function, which maps a change in utilization rate into a change in the depreciation rate[30]. The optimal choice of capital gives the resulting gross nominal returns on capital, which are transferred to the bank in exchange for funding

$$R_{K,E,t} = (1+\pi^{c}_{E,t})\frac{\Big(MC^{r}_{E,t}\alpha\frac{Z_{E,t}}{K_{t-1}}\Big) + (Q_{E,t} - \tau_{E,t}U_{E,t})}{Q_{E,t-1}}. \tag{I.22}$$

I.4 Capital producers

Capital producing firms buy and refurbish depreciated capital from the intermediate goods firm at price $P^{C}_{E,t}$ and also produce new capital using the RoW final good, which consists of domestically produced and imported retail goods, as an input. Furthermore we assume that they face quadratic adjustment costs on net investment[31] and that profits, which arise outside of the steady state, are distributed lump sum to the households. The optimal choice of investment yields the familiar *Tobins Q* relation for the evolution of the relative price of capital

$$\begin{aligned} Q_{E,t} = 1 &+ \frac{\Psi}{2}\Big(\frac{In_{E,t}+Iss_E}{In_{E,t-1}+Iss_E} - 1\Big)^2 + \Psi\Big(\frac{In_{E,t}+Iss_E}{In_{E,t-1}+Iss_E} - 1\Big)\frac{In_{E,t}+Iss_E}{In_{E,t-1}+Iss_E} \\ &- \beta\frac{\Lambda_{E_{t+1}}}{\Lambda_{E_t}}\Psi\Big(\frac{In_{E,t+1}+Iss_E}{In_{E,t}+Iss_E} - 1\Big)\Big(\frac{In_{E,t+1}+Iss_E}{In_{E,t}+Iss_E}\Big)^2 \end{aligned} \tag{I.23}$$

alongside the law of motion for capital

$$K_{E,t} = K_{E,t-1} + In_{E,t} \tag{I.24}$$

[29] As the market for claims is frictionless, arbitrage requires that the value of capital installed and used in next period's production has to equal the value of claims on capital ($Q_{E,t}S_{E,t} = Q_{E,t}K_{E,t}$).

[30] The adjustment cost function is given by $\tau_{E,t}(U_{E,t}) = \tau_{E,ss,scale} + \zeta_{E,1}\frac{U_t^{1+\zeta_2}}{1+\zeta_2}$ with $\tau_{E,ss,scale}$ as an exogenous scale parameter in order to normalize utilization in the steady state.

[31] Following Gertler & Karadi (2011) we assume that adjustment costs are only present when changing net investment in order for the optimal choice of the utilization rate to be independent from fluctuations in the relative price of capital $Q_{E,t}$

I.5 Goods bundling and pricing

I.5.1 Final consumption and investment good

They combine a final domestically produced good $Y^E_{E,t}$ and a final import good $Y^E_{F,t}$ into a combined final good, employing the following CES technology

$$Y^C_{E,t} = \Big[n_E^{\frac{1}{\psi_f}} Y^{E\,\frac{\psi_f-1}{\psi_f}}_{E,t} + (1-n_E)^{\frac{1}{\psi_f}} Y^{E\,\frac{\psi_f-1}{\psi_f}}_{F,t}\Big]^{\frac{\psi_f}{\psi_f-1}}. \tag{I.25}$$

The parameter n_E governs the share of domestically produced goods and thereby the degree of home bias in the assembling process[32]. The parameter ψ_f on the other hand corresponds to the elasticity of substitution between the final domestic and import good.

Taking the prices of the domestic final good $P^E_{E,t}$ and the price of the final import good expressed in domestic currency $(\mathcal{E}^F_{E,t} P^E_{F,t})$[33] as well as total demand from consumers and capital producers as given, the optimal demand for goods produced domestically and abroad is governed by

$$Y^E_{E,t} = n_E \Big(\frac{P^E_{E,t}}{P^C_{E,t}}\Big)^{-\psi_f} Y^C_{E,t} \tag{I.26}$$

$$Y^E_{F,t} = (1-n_E)\Big(\frac{\mathcal{E}^F_{E,t} P^E_{F,t}}{P^C_{E,t}}\Big)^{-\psi_f} Y^C_{E,t}. \tag{I.27}$$

Lastly note that the three equations above imply that the price of the final consumption and investment good in the RoW $P^C_{E,t}$ is (up to first order) a weighted average of the prices of the final domestic and import good

$$P^c_{E,t} = \Big[n_E P^{E^{1-\psi_f}}_{E,t} + (1-n_E)(\mathcal{E}^F_E P^E_{F,t})^{1-\psi_f}\Big]^{\frac{1}{1-\psi_f}}. \tag{I.28}$$

I.5.2 RoW domestically produced and sold final good

Table I.1 provides an overview of the core equations and first order conditions for the multistage bundling process.

I.5.3 Import good bundling

Table I.2 provides an overview of the core equations and first order conditions for the multistage bundling process of the final import good.

[32] The home bias parameter is adjusted in order to take into account the differences in country size as in Sutherland (2005). In particular, given a degree of general trade openness op_E and the relative country size of the RoW s, the parameter n_E takes the value $n_E = 1 - op_E(1-s)$ with a similar adjustment for the US counterpart

[33] Note that because of the pricing-to-market assumption the price for US exports expressed in US-\$ $P^E_{F,t}$ will in general be different from the price charged for US goods sold in the US $P^F_{F,t}$.

Table I.1: RoW domestic sales bundling

Production function/Price index	Demand functions
RoW domestically produced final good	
$Y_{E,t}^{E} = \left[\gamma_E^{E\frac{1}{\psi_i}} \tilde{Y}_{E,t}^{E\frac{\psi_i-1}{\psi_i}} + (1-\gamma_E)^{E\frac{1}{\psi_i}} \hat{Y}_{E,t}^{E\frac{\psi_i-1}{\psi_i}} \right]^{\frac{\psi_i}{\psi_i-1}}$ $P_{E,t}^{E} = \left[\gamma_E^E \tilde{P}_{E,t}^{E1-\psi_i} + (1-\gamma_E^E)\left(\mathcal{E}_{E,t}^F \hat{P}_{E,t}^E \right)^{1-\psi_i} \right]^{\frac{1}{1-\psi_i}}$	$\tilde{Y}_{E,t}^{E} = \gamma_E^E \left(\frac{\tilde{P}_{E,t}^E}{P_{E,t}^E} \right)^{-\psi_i} Y_{E,t}^E$ $\hat{Y}_{E,t}^{E} = (1-\gamma_E^E)\left(\frac{\mathcal{E}_{E,t}^F \hat{P}_{E,t}^E}{P_{E,t}^E} \right)^{-\psi_i} Y_{E,t}^E$
RoW domestically sold PCP good	
$\tilde{Y}_{E,t}^{E} = \left[\left(\frac{1}{\gamma_E^E} \right)^{\frac{1}{\psi_i}} \int_0^{\gamma_E^E} \tilde{Y}_{E,t}^E(i)^{\frac{\psi_i-1}{\psi_i}} di \right]^{\frac{\psi_i}{\psi_i-1}}$ $\tilde{P}_{E,t}^{E} = \left[\frac{1}{\gamma_E^E} \int_0^{\gamma_E^E} \tilde{P}_{E,t}^E(i)^{1-\psi_i} di \right]^{\frac{1}{1-\psi_i}}$	$\tilde{Y}_{E,t}^{E}(i) = \frac{1}{\gamma_E^E} \left(\frac{\tilde{P}_{E,t}^E(i)}{\tilde{P}_{E,t}^E} \right)^{-\psi_i} \tilde{Y}_{E,t}^E$ $= \left(\frac{\tilde{P}_{E,t}^E(i)}{P_{E,t}^E} \right)^{-\psi_i} Y_{E,t}^E$
RoW domestically sold DCP good	
$\hat{Y}_{E,t}^{E} = \left[\left(\frac{1}{1-\gamma_E^E} \right)^{\frac{1}{\psi_i}} \left(\int_{\gamma_E^E}^1 \hat{Y}_{E,t}^E(i)^{\frac{\psi_i-1}{\psi_i}} di \right) \right]^{\frac{\psi_i}{\psi_i-1}}$ $\mathcal{E}_{E,t}^F \hat{P}_{E,t}^{E} = \left[\frac{1}{(1-\gamma_E^E)} \int_{\gamma_E^E}^1 (\mathcal{E}_{E,t}^F \hat{P}_{E,t}^E(i))^{1-\psi_i} di \right]^{\frac{1}{1-\psi_i}}$	$\hat{Y}_{E,t}^{E}(i) = \frac{1}{1-\gamma_E^E} \left(\frac{\mathcal{E}_{E,t}^F \hat{P}_{E,t}^E(i)}{\mathcal{E}_{E,t}^F \hat{P}_{E,t}^E} \right)^{-\psi_i} \hat{Y}_{E,t}^E$ $= \left(\frac{\mathcal{E}_{E,t}^F \hat{P}_{E,t}^E(i)}{P_{E,t}^E} \right)^{-\psi_i} Y_{E,t}^E$

Table I.2: US import good bundling

Production function/Price index	Demand functions
US final import goods	
$Y_{E,t}^{F} = \left[\gamma_F^{E\frac{1}{\psi_i}} \tilde{Y}_{E,t}^{F\frac{\psi_i-1}{\psi_i}} + (1-\gamma_F)^{E\frac{1}{\psi_i}} \hat{Y}_{E,t}^{F\frac{\psi_i-1}{\psi_i}} \right]^{\frac{\psi_i}{\psi_i-1}}$ $P_{F,t}^{E^I} = \left[\gamma_F^E \left(\frac{\tilde{P}_{E,t}^F}{\mathcal{E}_{E,t}^F} \right)^{1-\psi_i} + (1-\gamma_F^E) \hat{P}_{E,t}^{F1-\psi_i} \right]^{\frac{1}{1-\psi_i}}$	$\tilde{Y}_{E,t}^{F} = \gamma_F^E \left(\frac{\tilde{P}_{E,t}^F}{\mathcal{E}_{E,t}^F P_{F,t}^{E^I}} \right)^{-\psi_i} Y_{E,t}^F$ $\hat{Y}_{E,t}^{F} = (1-\gamma_F^E)\left(\frac{\hat{P}_{E,t}^F}{P_{F,t}^{E^I}} \right)^{-\psi_i} Y_{E,t}^F.$
US imported PCP good	
$\tilde{Y}_{E,t}^{F} = \left[\left(\frac{1}{\gamma_F^E} \right)^{\frac{1}{\psi_i}} \left(\int_0^{\gamma_F^E} \tilde{Y}_{E,t}^F(i)^{\frac{\psi_i-1}{\psi_i}} di \right) \right]^{\frac{\psi_i}{\psi_i-1}}$ $\frac{\tilde{P}_{E,t}^F}{\mathcal{E}_{E,t}^F} = \left[\frac{1}{\gamma_F^E} \int_0^{\gamma_F^E} \left(\frac{\tilde{P}_{E,t}^F(i)}{\mathcal{E}_{E,t}^F} \right)^{1-\psi_i} di \right]^{\frac{1}{1-\psi_i}}$	$\tilde{Y}_{E,t}^{F}(i) = \frac{1}{\gamma_F^E} \left(\frac{\tilde{P}_{E,t}^F(i)}{\tilde{P}_{F,t}^F} \right)^{-\psi_i} \tilde{Y}_{E,t}^F$ $= \left(\frac{\tilde{P}_{E,t}^F(i)}{\mathcal{E}_{E,t}^F P_{F,t}^{F^I}} \right)^{-\psi_i} Y_{E,t}^F$
US imported DCP good	
$\hat{Y}_{E,t}^{F} = \left[\left(\frac{1}{1-\gamma_F^E} \right)^{\frac{1}{\psi_i}} \left(\int_{\gamma_F^E}^1 \hat{Y}_{E,t}^F(i)^{\frac{\psi_i-1}{\psi_i}} di \right) \right]^{\frac{\psi_i}{\psi_i-1}}$ $\hat{P}_{E,t}^{F} = \left[\frac{1}{(1-\gamma_F^E)} \int_{\gamma_F^E}^1 \hat{P}_{E,t}^F(i))^{1-\psi_i} di \right]^{\frac{1}{1-\psi_i}}$	$\hat{Y}_{E,t}^{F}(i) = \frac{1}{1-\gamma_F^E} \left(\frac{\hat{P}_{E,t}^F(i)}{\hat{P}_{E,t}^F} \right)^{-\psi_i} \hat{Y}_{E,t}^F$ $= \left(\frac{\hat{P}_{E,t}^F(i)}{P_{F,t}^{E^I}} \right)^{-\psi_i} Y_{E,t}^F$

I.6 Retail good pricing

I.3. The optimal price choice of a DCP firm i for its sales in the RoW market, taking into account the fact that it may not be able to reset its US-\$ denominated price $\hat{P}_{E,t}^E(i)$, can be written as

$$\max_{\hat{P}_{E,t}^E(i)} \mathbb{E}_t \sum_{s=0}^{\infty} \theta_p^{E^s} \Theta_{E_{t,t+s}} \left[\mathcal{E}_{E,t}^E \hat{P}_{E,t}^E(i) Y_{E,t}^E(i) - MC_{E,t} Y_{E,t}^E(i) \right]. \tag{I.29}$$

Table I.3: Market and pricing paradigm specific profit functions of RoW firms

Type of firm and market	Profit function
RoW market PCP firm	$\tilde{\Pi}^E_{E,t}(i) = \tilde{P}^E_{E,t}(i)\tilde{Y}^E_{E,t}(i) - MC_{E,t}\tilde{Y}^E_{E,t}(i)$
RoW market DCP firm	$\hat{\Pi}^E_{E,t}(i) = \mathcal{E}^F_{E,t}\hat{P}^E_{E,t}(i)\hat{Y}^E_{E,t}(i) - MC_{E,t}\hat{Y}^E_{E,t}(i)$
US import market PCP firm	$\tilde{\Pi}^F_{E,t}(i) = \tilde{P}^F_{E,t}(i)\tilde{Y}^F_{E,t}(i) - MC_{E,t}\tilde{Y}^F_{E,t}(i)$
US import market DCP firm	$\hat{\Pi}^F_{E,t}(i) = \mathcal{E}^F_{E,t}\hat{P}^F_{E,t}(i)\hat{Y}^F_{E,t}(i) - MC_{E,t}\hat{Y}^F_{E,t}(i)$

It is possible to show that the optimal reset price of a firm that sets its price for the RoW market in US-\$, relative to the aggregate RoW DCP sales price index $\hat{P}^E_{E,t}$, is given by

$$\frac{\hat{P}^E_{E,t}(i)}{\hat{P}^E_{E,t}} = \hat{p}^E_{E,t} = \frac{\psi_i}{(\psi_i - 1)}\frac{\hat{x}^E_{E,1,t}}{\hat{x}^E_{E,2,t}}. \tag{I.30}$$

The auxiliary recursive variables $\hat{x}^E_{E,1,t}$ and $\hat{x}^E_{E,2,t}$ read as

$$\hat{x}^E_{E,1,t} = \Lambda_{E,t}\Big(\frac{\mathcal{E}^F_{E,t}\hat{P}^E_{E,t}}{P^E_{E,t}}\Big)^{-\psi_i} Y^E_{E,t}\frac{P^E_{E,t}}{P^C_{E,t}}MC^{rp}_{E,t} + \beta\theta_p\mathbb{E}_t\hat{x}^E_{E,1,t+1}(1+\hat{\pi}^E_{E,t+1})^{\psi_i} \tag{I.31}$$

$$\hat{x}^E_{E,2,t} = \Lambda_{E,t}\Big(\frac{\mathcal{E}^F_{E,t}\hat{P}^E_{E,t}}{P^E_{E,t}}\Big)^{-\psi_i} Y^E_{E,t}\Big(\frac{\mathcal{E}^F_{E,t}\hat{P}^E_{E,t}}{P^C_{E,t}}\Big) + \beta\theta^E_p\mathbb{E}_t\hat{x}^E_{E,1,t+1}(1+\hat{\pi}^E_{E,t+1})^{\psi_i - 1}, \tag{I.32}$$

with $MC^{rp}_{E,t}$ as marginal costs deflated in by the aggregate producer price $P^E_{E,t}$. It becomes apparent that not only does the exchange rate $\mathcal{E}^F_{E,t}$ impact the optimal DCP price setting decision as it determines the demand for DCP goods via the relative price $\frac{\mathcal{E}^F_{E,t}\hat{P}^E_{E,t}}{P^E_{E,t}}$, it also impacts the optimal reset price via the term $\frac{\mathcal{E}^F_{E,t}\hat{P}^E_{E,t}}{P^C_{E,t}}$, which translates the local currency revenues that a DCP firm makes from selling one unit of its good $\mathcal{E}^F_{E,t}\hat{P}^E_{E,t}$ into the unit of account that the firm's owners (households) care about $P^C_{E,t}$. Everything else equal, an appreciation of the US-\$ exchange rate, will cause the local currency revenues per unit of DCP good sold to rise, while the input costs, which are denominated in the RoW currency, remain roughly stable. Thus the mark-up rises above the optimal mark-up and a DCP good firm would like to lower its US-\$ price in response to an appreciation of the US-\$ over and above what the induced fall in RoW demand for the DCP good would dictate. It is easy to verify that when aggregating across intra RoW sales of RoW DCP firms the inflation rate of the aggregate RoW sales DCP price (expressed in US-\$) is given by

$$1 = (1-\theta_p)\hat{p}^{E^{1-\psi_i}}_{E,t} + \theta_p(1+\hat{\pi}^E_{E,t})^{(\psi_i - 1)}, \tag{I.33}$$

where $\hat{p}^E_{E,t}$ denotes the ratio of the optimal reset price relative to the aggregate price index. Very similar equations hold for the optimal price of RoW retail firms that set their prices in the US import market in US-\$ as well as, with slight adaptions, for PCP firms.

I.7 Market clearing and the aggregate budget constraint

Turning to the market clearing conditions, aggregate demand for the domestic consumption good $Y^C_{E,t}$ is given by the sum of individual demand from all sources that either consume the good or use it as an input in production

$$Y^C_{E,t} = C_{E,t} + I_{E,t} + \frac{\Psi}{2}\Big(\frac{In_{E,t} + Iss_E}{In_{E,t-1} + Iss_E} - 1\Big)^2 (In_{E,t} + Iss_E). \tag{I.34}$$

Aggregating across all intermediate and retail goods firms and imposing market clearing yields the aggregate production function of the economy

$$Z_{E,t} = (U_{E,t}K_{E,t-1})^{\alpha} L^{(1-\alpha)}_{E,t} = \delta^E_{E,t}Y^E_{E,t} + \delta^F_{E,t}Y^F_{E,t}, \tag{I.35}$$

with $\delta^E_{E,t}$ and $\delta^F_{E,t}$ as price dispersion terms which are zero up to a first order approximation. $Y^E_{E,t}$ corresponds to the aggregate domestic demand for the final *domestically produced* RoW good given by

$$Y^E_{E,t} = n_E\Big(\frac{P^E_{E,t}}{P^C_{E,t}}\Big)^{-\psi_f} Y^C_{E,t}, \tag{I.36}$$

with $Y^C_{E,t}$ as the households and firms demand for the final good. Furthermore the aggregate demand for RoW goods produced for exports reads as

$$Y^F_{E,t} = \frac{1-s}{s}(1 - n_F)\Big(\frac{\mathcal{E}^F_{E,t}P^F_{E,t}}{P^C_{F,t}}\Big)^{-\psi_f} Y^C_{F,t}, \tag{I.37}$$

where it it is important to note that variables are expressed in per capita terms and therefore, following Sutherland (2005), the relative population size has to be taken when aggregating across countries as indicated by the ratio $\frac{1-s}{s}$.

Imposing double-entry bookkeeping i.e. that some RoW bank's interbank market liability $B^F_{E,j,t}$ has to always be an asset of a some US bank $B^{F\star}_{E,j,t}$ and taking into account the fact that population sizes differ yields a market clearing condition for the US-\$ interbank loan market

$$\int_0^s B^F_{E,j,t}dj = \int_s^1 B^{F\star}_{E,j,t}dj. \tag{I.38}$$

This can be translated into a solution for the aggregate ratio of interbank lending to domestic funding $\xi^F_{E,t}$ as a function of the share of RoW investments funded by US-\$ loans $\Xi^F_{E,t}$ given by

$$\xi^F_{E,t} = \frac{\frac{s}{1-s}\Xi^F_{E,t}Q_{E,t}K_{E,t}}{RER^F_{E,t}Q_{F,t}K_{F,t}}. \tag{I.39}$$

After aggregating the joint budget constraints of bankers and households and consolidating profits from all types of retail firm sales and the capital producing firms, one arrives at the familiar open-economy budget constraint

$$RER^F_{E,t}\Big(\frac{R^F_{E,b,t-1}}{(1+\pi^C_{F,t})}\Big)B^F_{E,t-1} + Y^C_{E,t} = \frac{P^E_{E,t}}{P^C_{E,t}}Y^E_{E,t} + \frac{P^F_{E,t}}{P^C_{E,t}}Y^F_{E,t} + RER^F_{E,t}B^F_{E,t}. \tag{I.40}$$

I.8 Calibration

Table I.4: Parameter values used in the simulations

Param.	**Val.**	**Description**	**Source**
		Households	
h_E	0.620	Habit persistence in consumption RoW	CKSW(2018)
h_F	0.790	Habit persistence in consumption US	JPT(2010)
σ_c	1.002	Intertemporal elasticity of substitution	$\approx$ log utility
φ	2.000	Inverse Frisch elasticity of labor	CKSW(2018)
β	0.995	Discount factor	2% ann. real rate
		RoW financial intermediaries	
ω_B^E	0.004	Start up funds RoW	endogenous in SS
θ_B^E	0.950	Survival probability of Banks RoW	AQ(2019)
ϵ_B	-2.409	Linear parameter incentive constraint	endogenous in SS
κ_B	8.301	Squared parameter incentive constraint	$\approx$ AQ(2019)
$\overline{\delta}_{B,E}$	0.460	Constant in incentive constraint	endogenous in SS
		US financial intermediaries	
ω_B^F	0.005	Start-up funds parameter US	endogenous in SS
θ_B^F	0.950	Survival probability of Banks US	AQ(2019)
$\delta_{B,F}$	0.391	Share of assets that the US Banker can divert	endogenous in SS
$\overline{\Gamma}$	1	Risk weight of global interbank loans	$\approx$ CKSW(2018)
		Wage decision	
ψ_w	6.000	Elasticity of substitution labor services	20% wage mark up
θ_w^E	0.780	Calvo parameter wages RoW	CKSW(2018)
θ_w^F	0.840	Calvo parameter wages US	JPT(2010)
		International trade	
ψ_f	1.120	Trade price elasticity	CKSW(2018)
op_E	0.200	General trade openness RoW	$\eta_E \approx 0.95$
op_F	0.185	General trade openness US	$\eta_F \approx 0.86$
n	0.750	Share of RoW in global economy	$1 - \frac{GDP_{US}}{GDP_{RoW}}$
		Intermediate goods production	
α	0.333	Share of capital in production	AQ(2019)
ζ_2	5.800	Elasticity of depreciation wrt. to utilization	JPT(2010)
$\tau_{E,ss}$	0.020	Normalization parameter depreciation RoW	endogenous in SS

Table I.4 –			
Param.	**Val.**	**Description**	**Source**
ζ_1^E	0.035	Normalization of utilization parameter RoW	endogenous in SS
ζ_1^F	0.035	Normalization of utilization parameter US	endogenous in SS
$\tau_{F,ss}$	0.020	Normalization parameter depreciation US	endogenous in SS
		Retail good pricing	
ψ_i	6.000	Elasticity of substitution retail goods	20% mark up
θ_P^E	0.820	Calvo parameter retail firms RoW	CKSW(2018)
θ_P^F	0.840	Calvo parameter retail firms US	JPT(2010)
$\widehat{\gamma_E^E} = 1 - \gamma_E^E$	0.09	Share of RoW domestic sales DCP firms	37.5% intra RoW exp.
$\widehat{\gamma_F^E} = 1 - \gamma_F^E$	0.97	Share of RoW export to US DCP firms	≈ G(2015) invoicing
$\widetilde{\gamma_E^E} = 1 - \gamma_E^F$	0.05	Share of US export LCP firms	≈ G(2015) invoicing
		Capital goods production	
Ψ_E	5.770	Investment adjustment costs RoW	CKSW(2018)
Ψ_F	2.950	Investment adjustment costs US	JPT(2010)
		Monetary Policy	
$\rho_{E,r}$	0.930	RoW interest rate smoothing	CKSW(2018)
$\phi_{E,\pi}$	2.740	RoW Taylor Rule coefficient inflation	CKSW(2018)
$\phi_{E,z}$	0.030	RoW Taylor Rule coefficient output	CKSW(2018)
$\rho_{F,r}$	0.810	US interest rate smoothing	JPT (2010)
$\phi_{F,\pi}$	1.970	US Taylor Rule coefficient inflation	JPT(2010)
$\phi_{F,z}$	0.050	US Taylor Rule coefficient output	JPT(2010)
		Steady State targets	
$L_{E,ss}$	0.333	SS labor target RoW	GK(2011)[a]
U_{ss}	1.000	SS utilization rate target RoW and US	JPT(2010)
τ_{ss}	0.025	SS depreciation rate target RoW and US	JPT(2010)
$S_{E,ss}$	0.005	SS credit spread target RoW (quarterly)	≈ CKSW(2018)
$S_{F,ss}$	0.005	SS credit spread target US (quarterly)	≈ avg. GZ spread
$\phi_{E,b,ss}$	6.00	SS leverage ratio target, RoW banks	CKSW(2018)
$\phi_{F,b,ss}^F$	4.00	SS local leverage ratio target, US banks	GK(2011)
$\Xi_{E,ss}^F$	0.18	SS target US-$ liabilities over assets RoW	≈ LBS avg.

[a] GK(2011), JPT(2010), CKSW(2018), GZ(2012), AQ(2019), G(2015), represent abbreviations for Gertler & Karadi (2011), Justiniano et al. (2010), Coenen et al. (2018), Gilchrist & Zakrajsek (2012), Akinci & Queralto (2019) and Gopinath (2015) respectively.