DISCUSSION PAPER SERIES

DP16245
(v. 2)

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INTERNATIONAL MACROECONOMICS AND FINANCE

ISSN 0265-8003

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Discussion Paper DP16245
First Published 10 June 2021
This Revision 10 October 2021

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www.cepr.org

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- International Macroeconomics and Finance

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Global Risk and the Dollar

Abstract

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JEL Classification: F31, F41, F44

Keywords: Us dollar, safe-haven currencies, risk shocks, trade channel, financial channel, Bayesian proxy structural VAR, minimum relative entropy, counterfactual, monetary policy

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Acknowledgements
The views expressed in the paper are those of the authors and do not necessarily reflect those of the ECB or the Eurosystem and should not be reported as such. We thank our discussants Xiang Li, Boris Fisera and Alejandro Vicondoa, as well as Michele Piffer, seminar and conference participants at the ECB, the IM/TCD Workshop on International Capital Flows and Financial Globalisation, the 25th ICMAIF, the 11th RCEA Money-Macro-Finance Conference, the 21st IWH-CIREQ-GW Macroeconometric Workshop, the 13th FIW Research Conference on International Economics, the 2021 IAAE Annual Conference, and University of Namur for useful comments. We are grateful to Silvia Miranda-Agrippino for sharing data on global factors in risky asset prices and capital flows.

Global risk and the dollar *

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Abstract

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1 Introduction

According to received wisdom the dollar appreciates when global risk goes up. The Global Financial Crisis (GFC) and the COVID-19 pandemic provide striking examples. We illustrate this in Figure 1, as we display the expected volatility index (VIX) for the S&P 500 jointly with the broad dollar index: both rise strongly at the height of the GFC (left panel) and the early stage of the pandemic (right panel). In fact, this co-movement is a general pattern in the data.[1] At a theoretical level, it can be rationalized on the ground that some US assets are particularly safe and/or liquid (Farhi & Gabaix 2016; Bianchi et al. 2021; Jiang et al. 2021a). At an empirical level, however, the fundamental drivers of the co-movement are not firmly established.

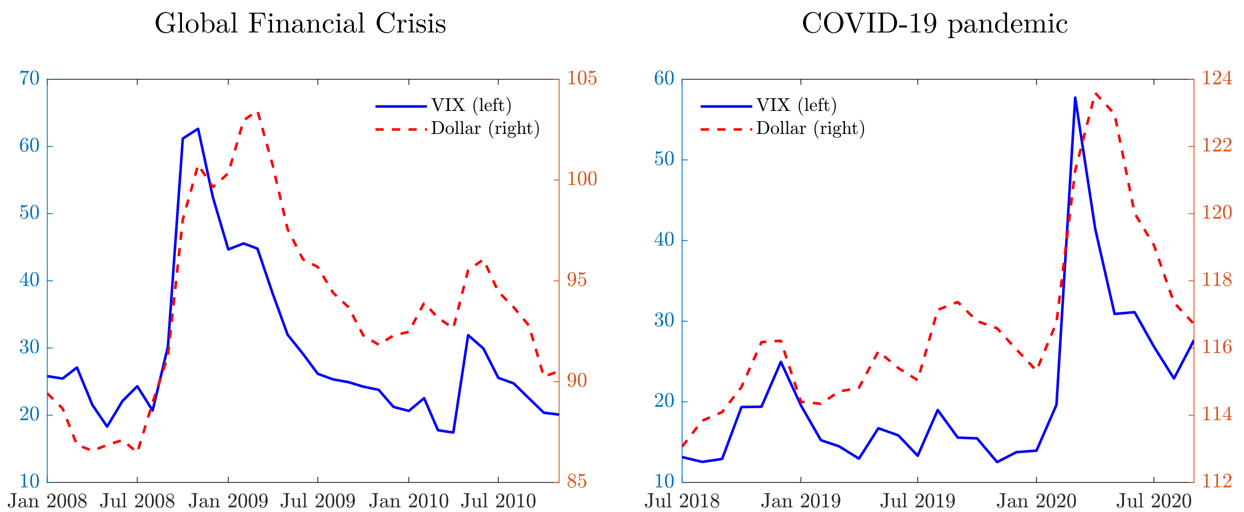
Against this background, we explore systematically how exogenous innovations to global risk transmit to the world economy. We give particular consideration to the response of the dollar and how it shapes the international transmission of global risk. Our key result is that global risk shocks induce dollar appreciation, which amplifies their contractionary effects through tighter financial conditions.

In order to estimate the effects of global risk shocks, we rely on high-frequency surprises in the price of gold—the ultimate safe asset—as external instruments in a Bayesian proxy vector-autoregressive model (Arias et al. 2018, forthcoming). As predicted by theory, we find that global risk shocks are followed by an appreciation of the dollar and other safe-haven currencies, 'flight-to-safety' as foreign holdings of US Treasury securities increase, an uptick in the US Treasury premium of Du et al. (2018) and Jiang et al. (2021b), an increase in the dollar liquidity buffers of banks and the share of dollar-denominated 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 (Fernández-Villaverde et al. 2015; Baker et al. 2016; Basu & Bundick 2017). US net exports contract, suggesting that the dollar appreciation induces expenditure switching (Gopinath et al. 2020). And global financial conditions tighten: equity prices drop, spreads increase, and cross-border bank credit contracts, in line with theory (Bruno & Shin 2015). These patterns conform well with the notion of a global financial cycle and an 'exorbitant duty' of the US in the global financial system (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.

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 dollar appreciation is absent. The contractionary effects of dollar appreciation that materialize through tighter financial conditions thus dominate the 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; as predicted by theory, we find this holds in particular for *dollar*-denominated cross-border credit (Ivashina et al. 2015).

[1] The t-value in a regression of changes in the VIX on changes in the dollar exchange rate over the period 01/1990-12/2020 is 5.8. The t-value is 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 exchange rate and the VIX



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

In more detail, we estimate a Bayesian proxy structural vector-autoregressive (BPSVAR) model as proposed by Arias et al. (2018, forthcoming). We use monthly observations for the period 1990-2019 and include eight variables in the baseline specification: 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 addition, we consider 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 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, we use the intra-daily change in the gold price as recorded on narratively selected dates related to global risk events (Piffer & Podstawski 2018; Engel & Wu 2018; Ludvigson et al. forthcoming). In order to explore the effects of a policy experiment in which the Federal Reserve stabilizes the dollar in an extension to our baseline analysis, we use interest-rate changes around Federal Open Market Committee announcements as an additional external instrument to identify US monetary policy shocks (Gertler & Karadi 2015; Jarociński & Karadi 2020).

The BPSVAR framework of Arias et al. (2018, forthcoming) 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 inference, especially when the external instruments are weak (Caldara & Herbst 2019; Montiel Olea et al. forthcoming). 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 may seem controversial (Angelini et al. 2019; Alessandri et al. 2020; Redl 2020; Carriero et al. forthcoming).

We find that a one-standard-deviation global risk shock appreciates the dollar by about 0.5%. Other currencies commonly labelled as safe-havens such as the Japanese yen and the Swiss franc also appreciate; other 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 convenience yield of US over foreign government bonds. Foreign holdings of US Treasury securities increase by up to 1%, indicating 'flight-to-safety' capital flows. US and RoW industrial production exhibit hump-shaped and rather synchronized contraction; the recessionary impact is strongest after about six months, with US and RoW industrial production falling 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 the dominant-currency paradigm the maximum contraction in US exports occurs on impact, while it is delayed for US imports (Gopinath et al. 2020). Global financial conditions reflected in global factors of risky asset prices and capital flows tighten. More 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 the concept of 'minimum relative entropy' (MRE) 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—hence the label. We apply the MRE approach to construct impulse responses for a counterfactual in which the dollar is unresponsive to a global risk shock but which is otherwise as similar as possible to the baseline.

In the 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 which operate via the "financial channel" dominate the expansionary effects via expenditure switching in the "trade channel". And indeed, while US net exports only fall by somewhat less in the counterfactual, global financial conditions tighten much less. Moreover, we find that the dollar plays a special role: in the counterfactual the global risk shock is associated with a weaker drop in especially *dollar*-denominated 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 explore a policy experiment in which the Federal Reserve deviates from its past behaviour and stabilizes the dollar in the face of a global risk shock. Our analysis of this policy experiment is motivated by the unprecedented emergency liquidity the Federal Reserve provided to many economies through various facilities during the COVID-19 pandemic. It is widely believed that this policy was crucial for preventing a global financial crisis (see Cetorelli et al. 2020). Theoretically, Federal Reserve swap lines can be conceived as increasing the supply of safe dollar 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). Technically, 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 exchange rate (e.g. Bachmann & Sims 2012; Epstein et al. 2019). We refer to this as a 'structural shock counterfactual' (SSC; Antolin-Diaz et al. 2021). We find that by adopting a more accommodative stance that prevents dollar appreciation, US monetary policy would mitigate substantially the contractionary effects of a global risk shock; however, such a scenario entails substantial price pressures and overshooting in real activity in the US.

The rest of the paper is organised as follows. In Section 2 we relate our paper to existing literature and spell out in detail our contribution. Section 3 outlines the BPSVAR framework, discusses our identification assumptions and priors. Section 4 presents our results for the effects of global risk shocks on exchange rates and the world economy. In section 5 we zoom in on the role of the dollar on the basis of a counterfactual, and in section 6 we carry out a policy experiment. Finally, Section 7 concludes.

2 Related literature

Our paper relates to a number of earlier contributions. First, it speaks to recent theoretical work on the special role of the US dollar exchange rate and US assets in the international monetary system. For example, Farhi & Gabaix (2016) consider a model in which economies differ in the resilience of their non-tradables sector's productivity to rare disasters. A high-resilience economy's productivity drops less when a disaster strikes, and hence its exchange rate reflecting the relative price of its non-tradeables to the world numeraire appreciates. Maggiori (2017) introduces heterogeneity in risk-bearing capacity rooted in differences in financial development in a two-country model and shows that this can account for the special role of the US and the dollar in the international monetary system, although it produces the 'reserve-currency paradox' that the dollar appreciates in response to a global risk shock. Jiang et al. (2021a) incorporate a demand for safe dollar assets that induces a non-pecuniary convenience yield and thereby drives the dollar exchange rate in a two-country overlapping generations model: An increase in the demand for safe dollar assets during a global crisis raises the convenience yield, which appreciates the dollar, thereby deteriorates the net worth of foreign borrowers subject to currency mismatches, and eventually elicits a financial accelerator-driven global contraction. Similar predictions emerge from the model of Kekre & Lenel (2021), which combines demand for safe dollar assets as in Jiang et al. (2021a) and heterogeneity in risk-bearing capacity as in Maggiori (2017). Bianchi et al. (2021) build a model for a dollar-dominated international financial system in which banks maintain a buffer of liquid dollar assets to insure against liquidity risk. Risk shocks increase banks' demand for dollar assets and thereby raise their liquidity premium, which appreciates the dollar exchange rate. Our contribution is to explore the empirical relevance of the mechanisms spelled out in these models, especially the responses of the dollar exchange rate, capital flows, and convenience yields to global risk shocks. 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 the literature that studies the empirical relationship between global risk and the dollar exchange rate. Lustig et al. (2014) document that a global dollar factor drives currency returns and that the dollar exchange rate features a positive and countercyclical safety premium. Verdelhan (2018) shows that a carry and a dollar factor account for a substantial share of the variation in bilateral exchange rates, and argues the dollar factor may reflect global macroeconomic risk. Lilley et al. (forthcoming) find that global risk measures perform as in-sample predictors of the dollar exchange rate. Engel & Wu (2018) and Jiang et al. (2021b) document that the convenience yield has explanatory and predictive power for the dollar exchange rate. Hassan et al. (2021) construct an elaborate measure of global risk based on textual analysis applied to earnings calls of thousands of publicly listed firms around the world and find it is associated with dollar appreciation. And Avdjiev, Du, et al. (2019) document that dollar appreciation is associated with larger deviations from covered interest parity and contractions of cross-border dollar-denominated bank credit. We complement this literature by moving from forecasting and reduced-form regressions to showing that identified, exogenous innovations to global risk cause dollar appreciation, adjustments in relative convenience yields, and contractions in cross-border credit flows.

Third, our paper contributes to empirical work on the role of financial channels in the global transmission of risk—or, alternatively labelled, uncertainty—shocks. Carriere-Swallow & Cespedes (2013) consider US uncertainty shocks given by changes in the VIX that exceed some pre-specified threshold in small-open economy VAR models for 40 countries and find that emerging market economies (EMEs) experience larger spillovers than advanced economies due to more pervasive credit constraints. Bhattarai et al. (2020) estimate contractionary spillovers to 15 EMEs from US uncertainty shocks represented by innovations to the VIX ordered last in a recursive VAR model. Cesa-Bianchi et al. (2018) estimate the effects of international credit supply shocks given by innovations to US broker-dealer leverage ordered first in a panel mean-group VAR model and find that they cause a decline in cross-border banking flows. And Epstein et al. (2019) identify global risk shocks as innovations to the US corporate bond spread in small open-economy panel VAR models and find that bank credit plays a key role in their global transmission. Relative to these studies, we zoom in on and quantify the role of the dollar exchange rate 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.

Finally, our findings on the role of the dollar for financial spillovers are based on aggregate data and thus complement existing evidence based on micro data. Shim et al. (2021) document that firms in 10 EMEs whose non-financial sectors hold more debt in foreign currency reduce their leverage relatively more after home-currency depreciation. Using firm-level data for 18 major economies, Banerjee et al. (2020) find that exchange rate depreciation dampens corporate investment through firm leverage and foreign-currency debt. Avdjiev, Bruno, et al. (2019) study firm-level data for 32 EMEs and find that dollar appreciation is associated with declines in real investment, which is stronger for firms that are more dependent on external financing and that are located in countries with floating exchange rates. And using micro data for Mexico, Bruno & Shin (2021) provide evidence that dollar appreciation reduces exports in particular for firms that rely on dollar liquidity to finance working capital. There is also evidence that dollar appreciation tightens financial conditions in the US (Niepmann & Schmidt-Eisenlohr 2017; Meisenzahl et al. 2019). Our analysis allows us to assess the net effect of dollar appreciation—contrasting trade and financial channels—in the context of global risk shocks on the aggregate economy. Moreover, we consider exogenous drivers of dollar appreciation rather than reduced-form regressions.

3 Empirical strategy

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

3.1 General framework

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

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

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 specification we include additional lags and deterministic terms, but omit them in Equation (1) 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 (i) correlated with the k unobserved structural shocks of interest $\boldsymbol{\epsilon}_t^*$, and (ii) orthogonal to the remaining unobserved structural shocks $\boldsymbol{\epsilon}_t^o$. Formally, the identifying assumptions are

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

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

and are known as the relevance and the exogeneity condition, respectively.

In order to operationalize these identifying assumptions the model in Equation (1) is augmented with equations for the k proxy variables $\boldsymbol{m}_t$. In particular, define $\tilde{\boldsymbol{y}}_t' \equiv (\boldsymbol{y}_t', \boldsymbol{m}_t')$, denote by $\tilde{\boldsymbol{A}}_\ell$ coefficient matrices of dimension $\tilde{n} \times \tilde{n}$ with $\tilde{n} = n + k$ and by $\tilde{\boldsymbol{\epsilon}} \equiv (\boldsymbol{\epsilon}_t', \boldsymbol{v}_t')' \sim N(\mathbf{0}, \boldsymbol{I}_{n+k})$. The augmented model is then given by

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

In order to preclude that augmenting the model in Equation (1) with equations for the proxy variables alters the dynamics of the endogenous variables, restrictions are imposed on the matrices $\tilde{\boldsymbol{A}}_\ell$ such that

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

Because the inverse of $\tilde{\boldsymbol{A}}_0$ is given by

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

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

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

The first equality is obtained using Equation (7) and because the structural shocks $\boldsymbol{\epsilon}_t$ are by assumption orthogonal to $\boldsymbol{y}_{t-1}$ and $\boldsymbol{v}_t$. The second equality is due to 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 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 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 Equation (6) it can be seen that this implies that the last k structural shocks impact the proxy variables contemporaneously. Arias et al. (forthcoming) develop an algorithm that estimates $\boldsymbol{A}_0$ 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 $\boldsymbol{\epsilon}_t^*$.

3.2 Appealing features of the BPSVAR 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. (forthcoming) 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. forthcoming). 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. forthcoming), either to derive the asymptotic distributions of the estimators or to ensure satisfactory coverage in bootstrap algorithms.[2]

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

$$\boldsymbol{R} = \left(\boldsymbol{\Gamma}_{0,2}^{-1\prime}\boldsymbol{\Gamma}_{0,2} + \boldsymbol{V}\boldsymbol{V}'\right)^{-1}\boldsymbol{V}\boldsymbol{V}'. \tag{9}$$

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).

To sum up, the BPSVAR framework of Arias et al. (forthcoming) is particularly appealing for our purposes as it allows us to: (i) avoid recursiveness assumptions between the endogenous variables, (ii) jointly identify multiple structural shocks, and (iii) carry out coherent inference when identification is achieved by multiple and possibly weak proxy variables.

3.3 Empirical specification

Our point of departure is the US VAR model of Gertler & Karadi (2015) which includes among the 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 (see for example 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).[3] We use monthly data for the time period from February 1990 to December 2019. We assume 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 B.1.

[2]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).

[3]We use AE instead of RoW policy rates as the latter exhibit extreme spikes reflecting periods of hyperinflation in some EMEs. We consider 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 from 1990.

3.4 Identification

We think of global risk shocks as events 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), differences in the risk-bearing capacity of economies' financial systems (Maggiori 2017), or frictions in interbank markets (Bianchi et al. 2021). In the data, it has been documented that both US and non-US investors impute a non-pecuniary convenience yield to especially US Treasury securities (Krishnamurthy & Vissing-Jorgensen 2012; Jiang et al. 2021b). During episodes of global turmoil a 'flight-to-safety' to US Treasury securities raises their relative convenience yield.

3.4.1 Proxy variables

Our proxy variables are constructed on the basis of high-frequency data in the spirit of work on the identification of monetary policy shocks (see Gertler & Karadi 2015, and references therein). Specifically, we draw on Piffer & Podstawski (2018) and consider the intra-daily changes in the price of gold—the ultimate safe asset—in a narrow window around auctions on narratively selected days as proxy variable for global risk shocks. Piffer & Podstawski (2018) first extend the list of exogenous risk events compiled by Bloom (2009). Second, they calculate the change in the price of gold between the last auction before and the first auction after the news about the risk event were released to markets.[4] Among the dates identified in Piffer & Podstawski (2018) we consider those labelled as 'global' and 'US' risk events; we include those labelled as 'European' and 'other' risk events in a robustness check.

For the US monetary policy shock we follow Gertler & Karadi (2015) and use the change of the 3-month Federal Funds Futures rate in a narrow time window around FOMC announcements as a proxy variable. We purge these interest rate surprises from central bank information effects 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.[5]

[4] 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 VIX (Bekaert et al. forthcoming).

[5] We aggregate the daily gold price and interest rate surprises to monthly frequency as in Gertler & Karadi (2015). In particular, we first create a cumulative daily surprise series, then, second, take monthly averages of these series, and, third, obtain monthly average surprises as the first difference of this series. 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)).

3.4.2 Identifying assumptions

Define $\boldsymbol{\epsilon}_t^* \equiv (\epsilon_t^r, \epsilon_t^{mp})'$, where ϵ_t^r denotes the unobserved global risk shock and ϵ_t^{mp} the unobserved US monetary policy shock. Furthermore, define $\boldsymbol{m}_t \equiv (p_t^{\epsilon,r}, p_t^{\epsilon,mp})'$ as the vector containing the observed proxy variables for the global risk and the US monetary policy shock, that is, the gold price and the (cleansed) Federal Funds futures surprises.

Our identifying assumptions are given by

$$E[\boldsymbol{\epsilon}_t^* \boldsymbol{m}_t'] = \begin{pmatrix} E[p_t^{\epsilon,r} \epsilon_t^r] & E[p_t^{\epsilon,mp} \epsilon_t^r] \\ E[p_t^{\epsilon,r} \epsilon_t^{mp}] & E[p_t^{\epsilon,mp} \epsilon_t^{mp}] \end{pmatrix} = \boldsymbol{V}, \tag{10a}$$

$$E[\boldsymbol{\epsilon}_t^o \boldsymbol{m}_t'] = \begin{pmatrix} E[p_t^{\epsilon,r} \boldsymbol{\epsilon}_t^o] & E[p_t^{\epsilon,mp} \boldsymbol{\epsilon}_t^o] \end{pmatrix} = \boldsymbol{0}. \tag{10b}$$

First, in the relevance condition in Equation (10a) we assume that global risk shocks drive the selected 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. (forthcoming) 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 (10b), Piffer & Podstawski (2018) document that gold price surprises are uncorrelated with a range of measures of non-risk shocks.[6]

It is worthwhile emphasizing that we do *not* include the VXO among the endogenous variables in the BPSVAR model in order to identify global risk shocks. Our identification of global risk shocks is unrelated to the choice of endogenous variables in $\boldsymbol{y}_t$; identification rests only on the assumptions about the relationship between structural shocks and proxy variables in Equations (10a) and (10b). The reason for including the VXO among the endogenous variables in $\boldsymbol{y}_t$ is to explore the adjustment of a measure of global risk in response to an exogenous shock. It is also worthwhile remarking that while the VXO reflects future volatility in the *US* stock market, it includes a large global component and is very strongly correlated with other economies' analogues (Londono & Wilson 2018).

Second, in the relevance condition in Equation (10a) we assume that US monetary policy shocks drive the Federal Funds futures surprises on FOMC announcement days, $E[p_t^{\epsilon,mp} \epsilon_t^{mp}] \neq 0$ (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 (10b), it seems plausible that in a narrow time window around FOMC

[6] The exogeneity condition for the gold price surprises might be questioned as on some of the dates also non-risk shocks may have materialized. However, note that the events considered by Bloom (2009), Piffer & Podstawski (2018) as well as Bobasu et al. (2021) are very diverse, meaning that even if on each and every event it was not only a global risk shock that 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 been in part a financial as well as a global risk shock, the 9/11 attacks or the launch of Operation Desert Storm were arguably no financial shocks. Therefore, we believe it is reasonable to assume that the only structural shock that has been *systematically* related to gold price surprises across the narratively dates are global risk shocks. Recall also that the BPSVAR framework allows the proxy variables to be affected by measurement error (see Equation (7)).

announcements monetary policy shocks are the only systematic drivers of Federal Funds futures surprises, especially after these have been purged from central bank information effects.

As discussed in Section 3.1, when multiple proxy variables are used to identify multiple structural shocks, the relevance and exogeneity conditions are not sufficient for point identification. In this case, additional restrictions need to be imposed on $\boldsymbol{V}$ in Equation (10a). A natural idea is to impose that $\boldsymbol{V}$ is a diagonal matrix, implying that Federal Funds futures surprises on FOMC announcement days are not driven by global risk shocks and that the narratively selected gold price surprises are not driven by US monetary policy shocks. Technically, this implies overidentifying restrictions, which cannot be implemented by the estimation algorithm of Arias et al. (forthcoming). 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, $E[p_t^{\epsilon,mp}\epsilon_t^r] = 0$. Note that this assumption is implicitly maintained in the literature on the effects of monetary policy shocks (Gertler & Karadi 2015; Caldara & Herbst 2019; Jarociński & Karadi 2020). Moreover, the assumption is mild in the context of our analysis, since we purge the Federal Funds futures surprises of central bank information effects. Nevertheless, below we consider a robustness check in which we relax the assumption $E[p_t^{\epsilon,r}\epsilon_t^\ell] = 0$ for $\ell \neq r$, by replacing the corresponding zero restrictions in Equations (10a) and (10b) with restrictions on their relative magnitude.[7]

For consistency we follow Caldara & Herbst (2019) and Arias et al. (forthcoming) 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 US monetary policy and global risk shocks, respectively; this is weaker than the relevance threshold of $\gamma = 0.2$ used by Arias et al. (forthcoming), 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.

Finally, note that the use of the BPSVAR framework allows us to avoid imposing potentially controversial recursiveness assumptions in order to identify the global risk shock. In particular, it is often assumed that only risk shocks have a contemporaneous effect on the corresponding measure in the VAR model (Bloom 2009; Jurado et al. 2015; Baker et al. 2016; Basu & Bundick 2017). However, this assumption seems restrictive; 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); our results below based on identification assumptions that do not impose restrictions on the contemporaneous relationship between endogenous variables confirm this finding.[8]

[7]Note that when two proxy variables are used to identify two structural shocks, a single additional zero restriction on $\boldsymbol{V}$ in Equation (10a) 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).

[8]Other approaches to overcome the limitations of recursive identification in the context of global risk/uncertainty shocks exploit heteroskedasticity across regimes or over time for identification (Angelini et al. 2019; Carriero et al. forthcoming), 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 Results

In this section we first present results for the effects of global risk shocks on the variables in the baseline specification of the model. Then, we present impulse responses of variables we add to the baseline specification to lend further credibility to our identification scheme and speak to the various mechanisms put forth in the theoretical literature. Following this, we consider the impulse responses of a second set of additional variables to flesh out the transmission of global risk shocks to the RoW through trade and financial conditions. We also explore cross-sectional differences by presenting separately effects for AEs and EMEs. Finally, we report results for several robustness checks.

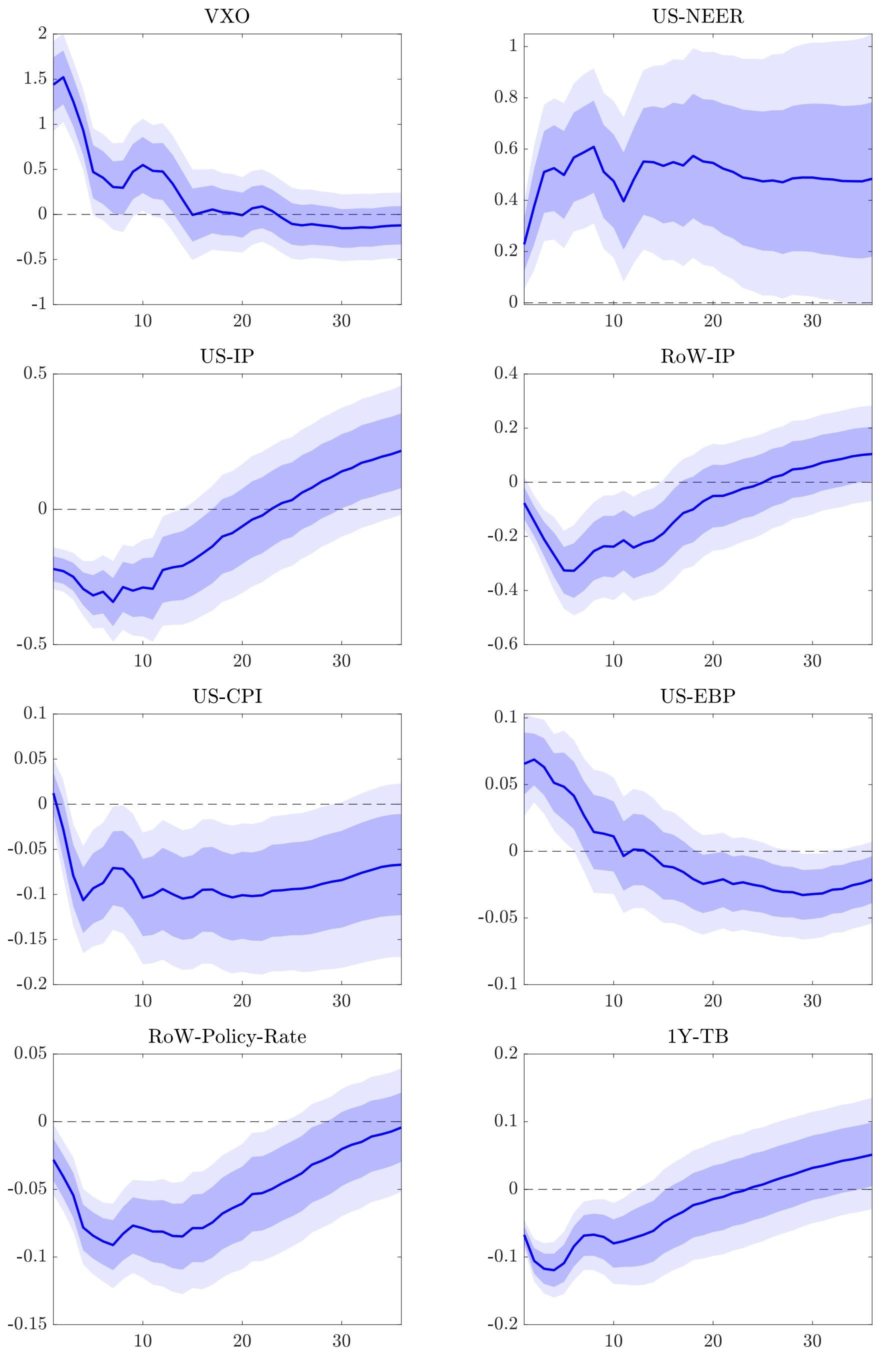
4.1 The effect of global risk shocks

We show the impulse responses to a global risk shock for the baseline specification in Figure 2. The solid lines represent the point-wise posterior means of the impulse responses, and shaded areas indicate 68% (dark) and 90% (light) equal-tailed point-wise credible sets. In each panel, time is measured in months along the horizontal axis and the deviation from the pre-shock level along the vertical axis. We consider a one-standard deviation global risk shock.

The response of the VXO is shown in the upper-left panel. It rises on impact and reaches a peak of about 1.5 index points one month after impact, and returns to the pre-shock level after about one year. The dollar appreciates on impact by about 0.2%, and reaches a maximum appreciation of 0.6% after about nine months (upper-right panel). The appreciation is persistent in that the dollar remains expensive relative to the baseline. That the dollar appreciates strongly and persistently in response to a global risk shock is the first key result of our paper. It is consistent with the prediction for the behaviour of the dollar exchange rate in response to risk and uncertainty shocks in 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).

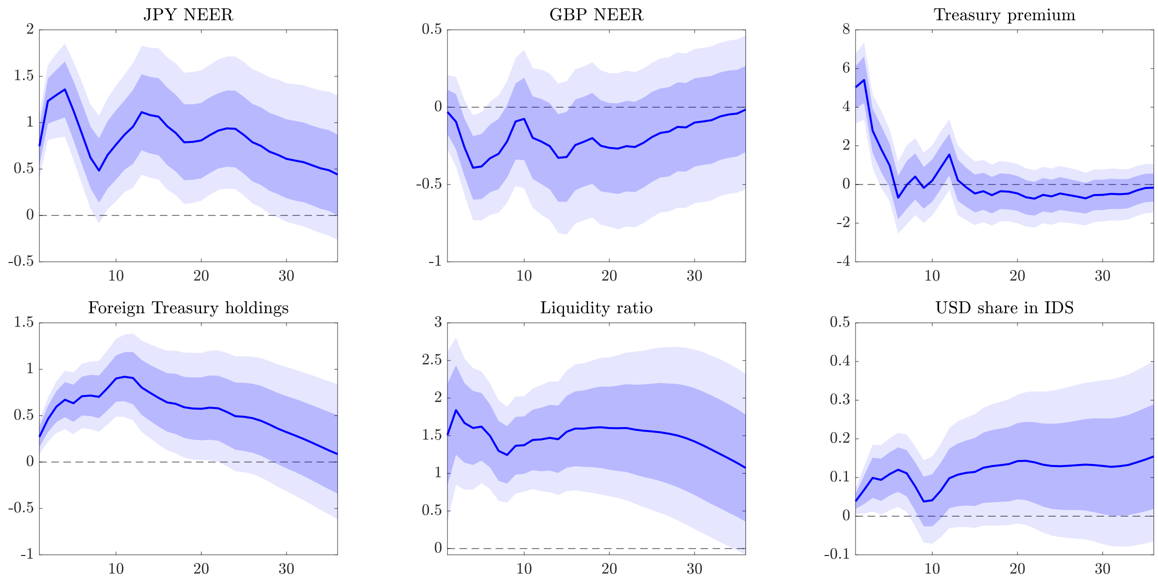
Industrial production in the US and the RoW, shown in the second row, contract in tandem. In both cases, there is a sharp contraction with a trough of almost 0.4% reached after 5-6 months. The contraction is more immediate and somewhat larger in the US than in the RoW, consistent with the notion that the reserve asset issuer bears the brunt of 'safety traps' (Kekre & Lenel 2021; Caballero et al. 2020). Economic activity recovers and reaches the pre-shock level after about 24 months, followed by some overshooting. This pattern of adjustment in economy activity is a well-established feature of risk and uncertainty shocks, and has been explored by various studies in a closed-economy context following the seminal analysis of Bloom (2009). Our results regarding the dollar appreciation and the simultaneous contraction in global real activity are also consistent with theoretical predictions in Jiang et al. (2021a) and Kekre & Lenel (2021). For example, the model of Jiang et al. (2021a) features a demand for safe dollar assets which gives rise to a non-pecuniary convenience yield on dollar assets and financial accelerator effects. An increase in the demand for safe dollar assets due to a global crisis raises the convenience yield and appreciates the dollar. This tightens financing conditions for foreign borrowers subject to currency mismatches as it reduces their net worth, which

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.

Figure 3: Impulse responses of additional variables



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. Because data on the liquidity ratio is only available from 2001 we use informative Minnesota-type priors and optimal hyperpriors/prior tightness as suggested by Giannone et al. (2015) in the estimation.

amplifies the recession.

Consistent with the contractionary effects of the global risk shock US consumer prices (third row left panel) fall below their pre-shock level and remain persistently reduced by about 0.1%. The US external bond premium rises by about 0.06% on impact and remains elevated for almost one year (third row right panel). Finally, monetary policy in the US and the RoW are loosened in tandem in response to a global risk shock, with a maximum decline in interest rates by about 0.1%.

4.2 Responses of additional variables

A key contribution of this paper is to establish the causal effect of global risk on the dollar exchange rate, how it transmits to global financial conditions and eventually the world business cycle. In our analysis, we pull together several strands 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. Specifically, we next present the impulse responses of 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). Technically, we modify the baseline specification by including one additional variable at a time. In this way we keep the dimensionality of the VAR model limited; we consider a large VAR model in which we include all additional variables simultaneously 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 the currencies of other safe-haven economies to also appreciate (Farhi & Gabaix 2016). Besides the dollar, the Japanese yen is typically considered a safe-haven currency as well (Ranaldo & Söderlind 2010; De Bock & de Carvalho Filho 2015). Indeed, the first panel in Figure 3 documents that the yen appreciates in response to the shock we identify. In contrast, as shown in the second panel the British pound depreciates; Figure A.1 in the Online Appendix documents that results for the Swiss franc and the euro as alternative safe-haven and non-safe-haven currencies are very similar. This pattern is consistent with the 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 based on 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' relative convenience yield, which is in turn triggered by a drop in the supply/increase in the demand for safe and liquid dollar assets during a global crisis. The top-right panel in Figure 3 depicts the impulse response of the Treasury premium of Du et al. (2018)—or, inversely defined, the Treasury basis of Jiang et al. (2021b)—reflecting the currency-hedged difference between the convenience yields of US Treasury securities and other G10 countries' sovereign bonds. The third panel in Figure 3 shows that the US Treasury premium indeed increases sharply in response to a global risk shock.

Third, in Jiang et al. (2021a) the mechanism through which a global risk shock appreciates the dollar works through an increase in the convenience yield of US Treasury securities that is driven by a rise in the demand for safe and liquid US assets. The model of Maggiori (2017) also highlights 'flight-to-safety' by non-US financial intermediaries during global crises in terms of capital flows. Consistent with this prediction, the first panel in the second row of Figure 3 shows that foreign holdings of US Treasury securities indeed increase in response to the global risk shock.[9] This finding is also consistent with empirical work studying capital flows during risk-off periods (Habib & Stracca 2015) and specific prominent events such as the GFC (Noeth & Sengupta 2010). Interestingly, that 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, which increases global demand for dollar assets, raises the convenience yield, and thereby appreciates the dollar. Bianchi et al. (2021) provide evidence for this predicted correlation in regressions of the dollar exchange rate on measures of banks' liquidity

[9] 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 this risk-off event was marked by a short-lived 'dash for cash' in which investors actually liquidated US Treasury holdings (Haddad et al. forthcoming).

ratio defined as the sum of commercial banks' reserves held at the Federal Reserve and government securities relative to short-term funding through demand deposits and commercial paper. The middle panel in the second row in Figure 3 documents that our global risk shock indeed induces a positive correlation between dollar appreciation and the liquidity ratio, which rises by almost 2 percentage points.

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 reflected in the corporate basis correlate with a higher dollar share in their total corporate bond issuance. The last panel in 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 by about 0.1 percentage point in response to a global risk shock.

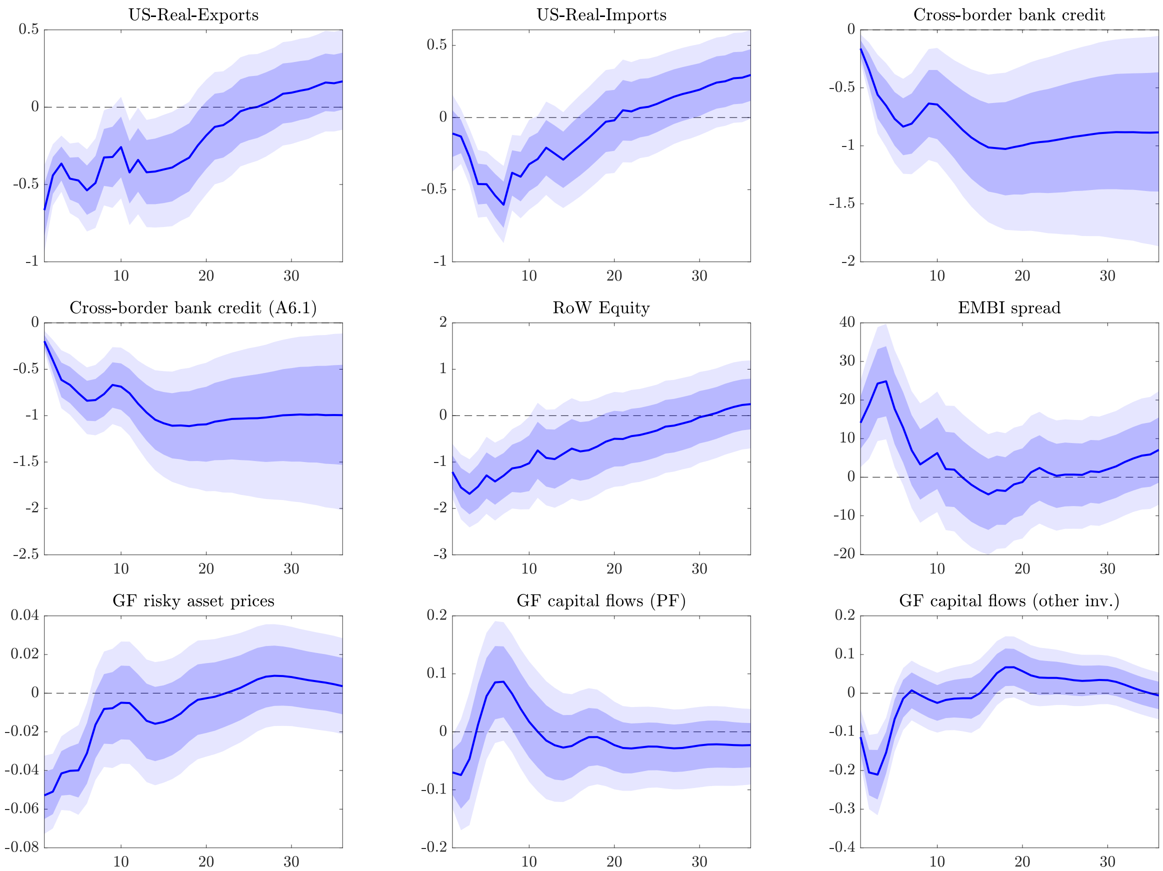
4.3 Trade and financial channels in the transmission of global risk shocks

We next explore in more detail the transmission of global risk shocks to the RoW focusing on variables that are particularly closely connected to exchange rate movements. Specifically, theory suggests that the dollar appreciation induced by global risk shocks transmits to the RoW through an expansionary trade and a contractionary financial channel (Bruno & Shin 2015; Jiang et al. 2021a; Obstfeld & Rogoff 1996; Gopinath et al. 2020).

The first two panels in Figure 4 show the impulse responses of US real exports and imports. Both decline in response to a global risk shock. Consistent with the notion of dominant-currency pricing the decline of exports is front loaded and stronger than that of imports: when both US import and export prices are sticky in dollar, a dollar appreciation induces expenditure switching only away from US exports but not imports; in contrast, the weaker and delayed decline of US imports tracks the hump-shaped contraction of economic activity in the US (Gopinath et al. 2020).

The remaining panels in Figure 4 present the impulse responses of various variables reflecting global financial conditions. Cross-border bank credit to non-US borrowers drops sharply and persistently by up to 1% in response to the global risk shock, whether measured in terms of cross-border liabilities (third panel in the first row) or cross-border claims (first panel in the second

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



Note: See notes to Figure 2 and Figure 3. The upper-right panel depicts the response of cross-border bank credit taken from the BIS Locational Banking Statistics Table A7 based on 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 from the BIS Locational Banking Statistics Table A6.1 based on residency principle (calculated as "Banks' external claims on all sectors in all countries" less "Banks' external claims on all sectors in the US").

row) of globally active banks.[10,11] The tightening in global financial conditions induced by global

[10]As in Bruno & Shin (2015) for cross-border bank credit we rely on data reported in Table A7 of the Locational Banking Statistics of the BIS. The data are originally available at quarterly frequency, and we use linear interpolation to convert the data 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 B.1 for variable definitions/descriptions. The advantage of the data in Table A7 is that it is based on the nationality principle, meaning 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, it should be noted that the coverage of the BIS reporting banks even in the 1990s amounted to about 90%.

[11]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 credit after the GFC. Figure A.11 in the Online Appendix documents that while identified 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.

risk shocks also manifests in a drop in RoW equity prices and an increase in the EMBI spread. And the third row presents the responses of 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 the global factors in 'portfolio' and 'other investment' flows from Miranda-Agrippino et al. (2020), respectively. In particular the common factors in global 'other investment' flows—which includes bank loans—and in risky asset prices drop markedly in response to a global risk shock.

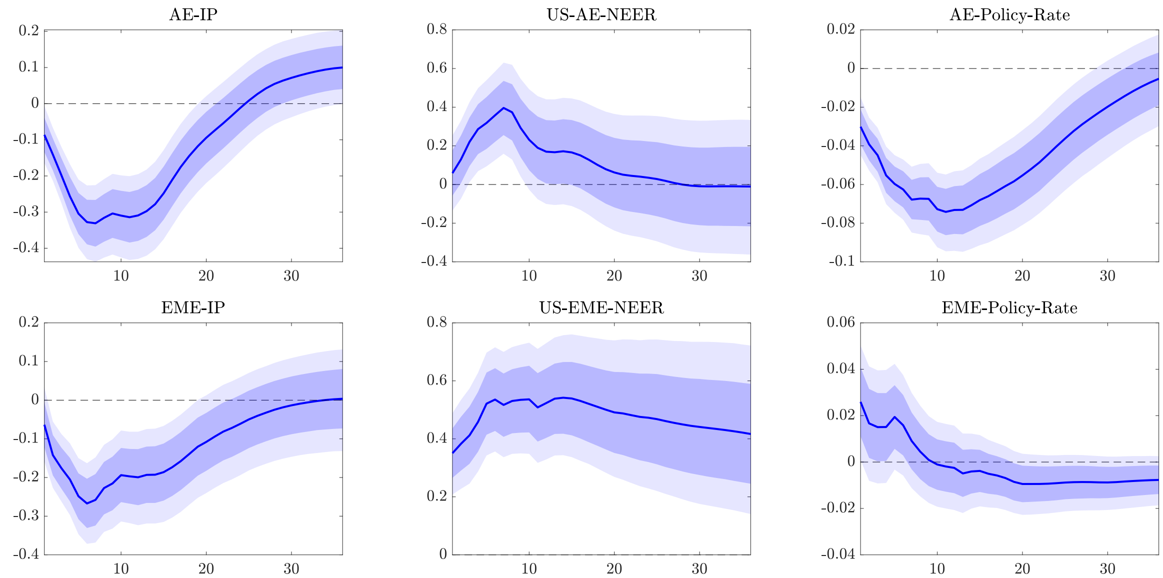
Note that our results also support the notion of an 'exorbitant duty' of the US (Gourinchas et al. 2012, 2017). In particular, the US is—roughly speaking—long in RoW-currency denominated portfolio equity and direct investment assets but short in dollar-denominated safe US portfolio debt—in particular US Treasury—liabilities. In this setting, when a global risk shock materializes that 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. Thus, the US provides insurance to the RoW in a global risk event. Our results are qualitatively consistent with this notion of an 'exorbitant duty': the dollar appreciates, RoW equity prices drop, and that yields of US Treasury securities decline—implying increases in their prices—in response to global risk shocks (see Figure 2). A more direct test for the empirical validity of the 'exorbitant duty' based on impulse responses of the US net foreign asset position is not possible, as even quarterly data are only available from 2006 onwards.

4.4 Effects in advanced and emerging market economies

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 5 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 on industrial production are very similar across AEs and EMEs (panels in the first column). At the same time, the dollar appreciates much more strongly against EME than against AE currencies (panels in the second column). And also monetary policy responses are starkly different: While interest rates fall in AEs, they actually rise in EMEs (panels in the third column). 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 (Corsetti et al. 2021). Interestingly, the responses of interest rates and exchange rates in Figure 5 are consistent with those in Kalemli-Özcan (2019): as EMEs try to prevent depreciation against the dollar by tightening monetary policy, this turns out to be self defeating as it induces currency risk premia to rise, so that depreciation is eventually larger.

Figure 5: Impulse responses for AEs and EMEs to a global risk shock



Note: The figure presents the baseline and counterfactual impulse responses to a one-standard deviation global risk shock for AEs and EMEs. 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.

4.5 Robustness

First, we verify whether what we identify as the effect of a global risk shock is not merely a global demand shock (Leduc & Liu 2016). To do so, we identify a global demand shock in addition to the global risk shock (and the US monetary policy shock) by imposing sign restrictions on the responses of several variables; importantly, we leave the response of the dollar unrestricted, as, for instance, in Enders et al. (2011).[12] Figure A.2 in the Online Appendix shows that a contractionary global demand shock entails impulse responses which are mostly qualitatively similar to those of a global risk shock. However, the response of the dollar exchange rate is different: it does not appreciate on impact and only weakly over time.[13] Moreover, Figure A.3 in the Online Appendix shows that while the gold price declines in response to a global demand shock, it increases in response to a global risk shock; note that this is not hard-wired in the estimation, because our identifying assumptions only impose that the gold price increases on the day—and not the entire month—a global risk shock occurs.

A second issue is that the gold price surprises might—even if we consider narratively selected

[12]Specifically, we impose the contemporaneous restrictions that a contractionary global demand shock reduces industrial production in the US and the RoW, US consumer prices in the as well as interest rates in the US and the RoW. At the same time, we restrict the global demand shock to raise the excess bond premium.

[13]We also note that our estimated responses to the global risk shock are qualitatively different from those that have 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.

events—be contaminated by the effects of other structural shocks. In order to address this concern, we relax the identifying assumption $E[p_t^{\epsilon,r}\boldsymbol{\epsilon}_t^o] = \mathbf{0}$ in Equation (10b). In particular, as in Ludvigson et al. (forthcoming) we allow the structural shocks in $\boldsymbol{\epsilon}_t^o$ to be correlated with the gold price surprises, and only impose that the correlation between the gold price surprises and the global risk shock is stronger than for any other structural shock: $|E[p_t^{\epsilon,r}\epsilon_t^r]| > |E[p_t^{\epsilon,r}\epsilon_t^\ell]|$ for $\ell \neq r$. Figure A.4 in the Online Appendix shows that the results for this alternative identification scheme are very similar to the baseline in Figure 2.

Third, 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. Of course, these are reduced-form correlations that are in general not indicative of the strength of structural relationships such as those we explore in this paper. Nevertheless, it may be worthwhile to explore possible differences in our findings before and after the GFC in 2007. Figure A.5 in the Online Appendix documents that our results regarding the effects of global risk shocks on the VXO and the dollar exchange rate 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, which could have made it more difficult to forecast the dollar exchange rate with the VIX than after the GFC.

Fourth, Figure A.6 in the Online Appendix documents that our results are very similar if instead of estimating our baseline BPSVAR model with flat priors we estimate a large BPSVAR model that includes simultaneously many additional variables. Due the implied dimensionality of the model, it needs to be estimated with informative Minnesota-type priors and optimal hyperpriors/prior tightness (Giannone et al. 2015).

Fifth, Figures A.7 and A.8 in the Online Appendix document that two alternative definitions of the gold price surprise proxy variable deliver very similar results. Figure A.7 shows results for the case in which we also consider the days with 'European' and 'other' rather than only 'global' and 'US' risk events. And in Figure A.8 we consider only those events that were associated with positive gold price surprises.

Sixth, recall that we think of global risk shocks as incidents that are associated with an increase in the demand for safe assets. In theory, this may be driven both by an increase in risk aversion ('price of risk') and uncertainty ('quantity of risk'). In fact, Piffer & Podstawski (2018) construct the gold price surprises to study what they label 'uncertainty' rather than 'risk' shocks. Figure A.9 in the Online Appendix displays the impulse responses of separate measures of risk aversion and uncertainty constructed by Bekaert et al. (forthcoming). Indeed, we find that both increase in response to a global risk shock.

Finally, Figure A.10 in the Online Appendix documents that our results are very similar if we do not impose a relevance threshold.

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

The results in Figure 4 suggest that dollar appreciation induced by a global risk shock may be transmitting both through an expansionary trade and a contractionary financial channel. Given that these channels work in opposite directions, we now provide an assessment of the net contribution of dollar appreciation to the effects of global risk shocks in the RoW. To do so, we construct a counterfactual in which a global risk shock does *not* cause dollar appreciation. We first provide some details on the approach we adopt to construct the counterfactual.

5.1 MRE 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 their 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 this idea to impulse responses based on the notion that these can be conceived of as conditional forecasts.

Assume for simplicity of exposition 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 corresponds to 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 are unknown and have to be estimated based on 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

$$f(\widetilde{\boldsymbol{y}}_{T+h}|\boldsymbol{y}_{1,T}, \mathcal{I}_a, \widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}) \propto p(\boldsymbol{\psi}) \times \ell(\boldsymbol{y}_{1,T}|\boldsymbol{\psi}, \mathcal{I}_a) \times \nu, \tag{11}$$

where $\mathcal{I}_a$ represents the identification assumptions in Equations (10a) and (10b), $p(\cdot)$ is the prior about the structural VAR parameters $\boldsymbol{\psi}$, and ν the volume element of the mapping from the structural VAR parameters to the impulse responses; the pointwise mean of f is shown as the blue solid lines in Figure 2.

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:

$$\begin{gathered} Min_{\psi}\ \mathcal{D}(f^*||f) \quad \text{s.t.} \\ \int f^*(\widetilde{\boldsymbol{y}})\widetilde{y}^{\$}d\widetilde{\boldsymbol{y}} = E(\widetilde{y}^{\$}) = 0, \ \ \int f^*(\widetilde{\boldsymbol{y}})d\widetilde{\boldsymbol{y}} = 1, \ \ f^*(\widetilde{\boldsymbol{y}}) \geq 0, \end{gathered} \tag{12}$$

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 (12) for simplicity). In general, 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. The counterfactual VAR model with structural parameters $\boldsymbol{\psi}^*$ is then implied by the counterfactual impulse responses $\widetilde{\boldsymbol{y}}$ based on the mapping between impulse responses and structural VAR parameters (see Arias et al. 2018, Appendix B); note however that for the purposes of our paper we do not need to back out the counterfactual structural VAR model parameters $\boldsymbol{\psi}^*$ once we have determined the counterfactual impulse responses (see below). Intuitively and roughly speaking, MRE determines that counterfactual VAR model in which the dollar exchange rate 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.[14]

It turns out that the solution to the problem in (12) in terms of the counterfactual distribution f^* can be computed by updating the baseline posterior f given the 'information' that the dollar exchange rate is unresponsive to a global risk shock. Specifically, we have

$$\begin{aligned} &f^*\left(\widetilde{\boldsymbol{y}}_{T+h}|\boldsymbol{y}_{1,T},\mathcal{I}_a,\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h},\widetilde{y}^{\$}_{T+h}=0\right)\propto \\ &\qquad f(\widetilde{\boldsymbol{y}}_{T+h}|\boldsymbol{y}_{1,T},\mathcal{I}_a,\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h})\times\tau\left(\widetilde{y}^{\$}_{T+h}(\boldsymbol{\psi})\right), \end{aligned} \tag{13}$$

where τ is a 'tilt' function (Robertson et al. 2005). The tilt τ down-weights the actual posterior beliefs for those values of the VAR parameters that are associated with large deviations from the counterfactual constraint that the dollar exchange rate 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 Online Appendix C for details). Once the tilted weights are obtained, importance sampling techniques can be used to estimate the mean and percentiles of the counterfactual posterior distribution.

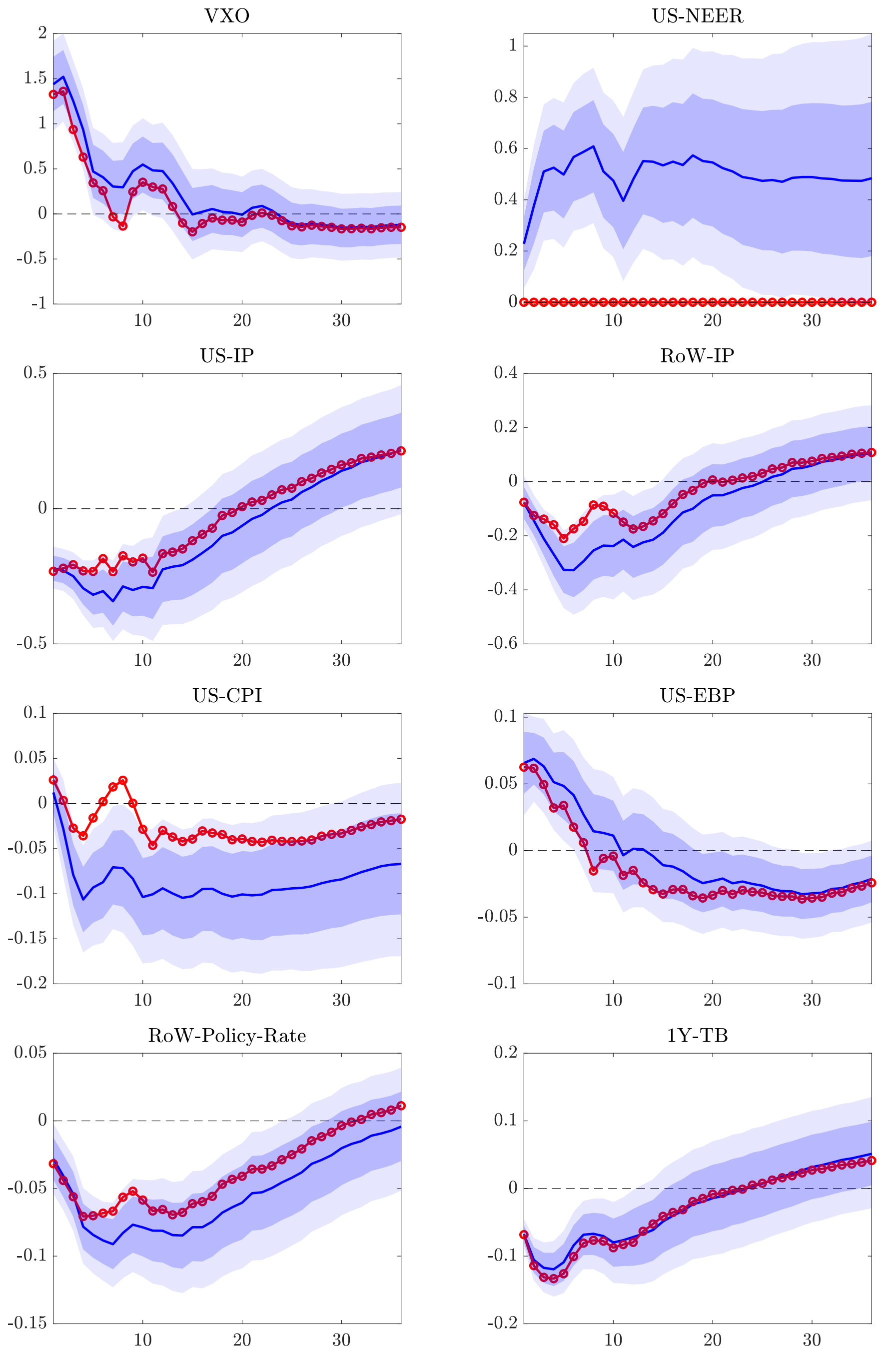
5.2 A 'no-appreciation' counterfactual benchmark

Figure 6 shows that in the counterfactual in which the dollar exchange rate 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 when dollar appreciation is absent implies that the contractionary financial channel dominates the expansionary trade channel.

One may ask whether the differences between the baseline and the counterfactual are 'statistically significant'. This question is not as straightforward as in a standard regression in which one tests

[14]Earlier studies have instead constructed conterfactuals 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; in general, the bias is 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.

Figure 6: 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.

the null of a coefficient estimate being equal to a non-random benchmark. Instead, in our context we suggest such a test for 'statistical significance' to assess how 'different' the baseline and the tilted posterior distributions are. To do so, we consider a test for first-order stochastic dominance.

To provide intuition, consider the effect of a global risk shock on RoW real activity. 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 in this sense meaningfully different from the baseline.

We next explore whether the finding that the financial channel dominates the trade channel also emerges from the responses of global financial conditions and trade. Indeed, Figure 7 shows that 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 through expenditure switching is not very powerful in the first place. This is 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 (Georgiadis et al. 2019).

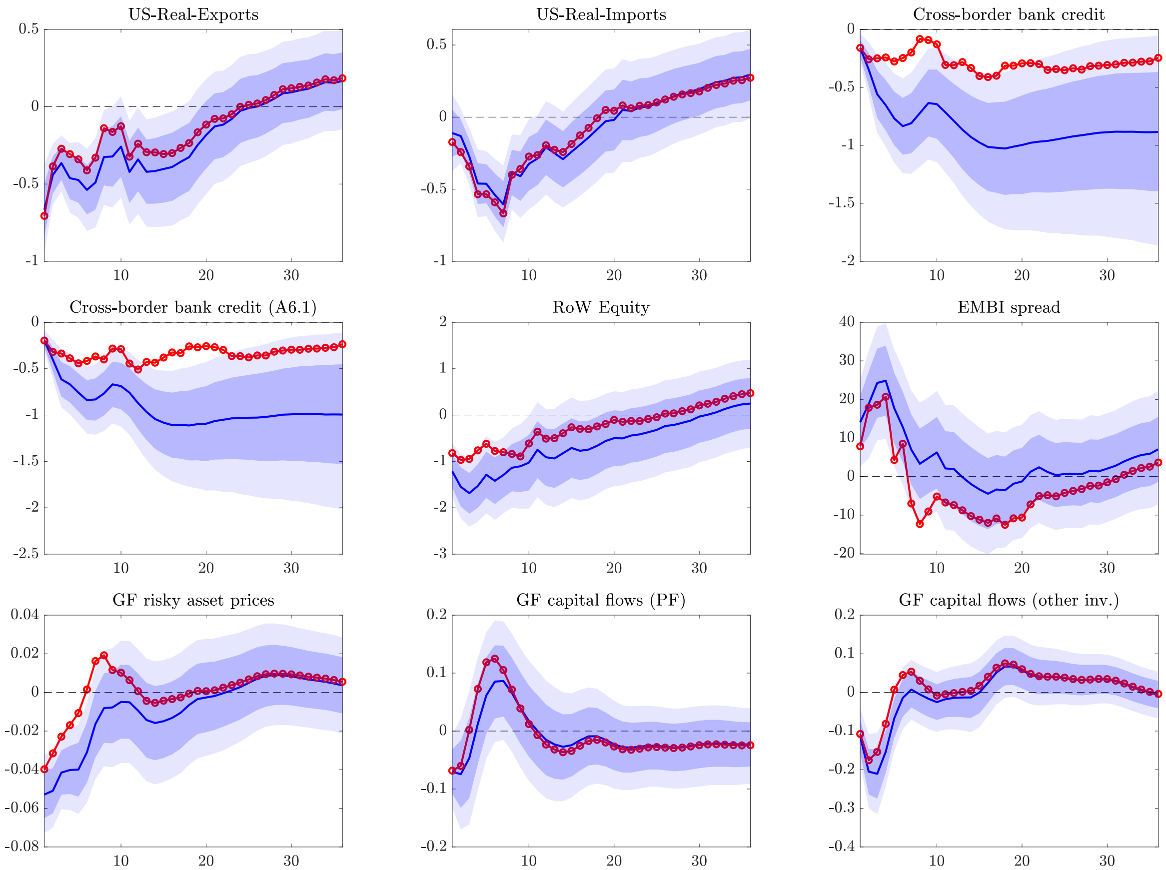
A much bigger difference between the baseline and the counterfactual is obtained for the variables reflecting global financial conditions. Figure 7 shows that cross-border bank credit and RoW equity prices contract by much less in the counterfactual.[15] The EMBI spread rises by less, although the difference is not as pronounced as for cross-border credit and equity prices. Finally, especially the global factors in risky asset prices and—although less pronounced—in 'other investment' flows fall by much less in the counterfactual.

In sum, the findings in Figure 7 are consistent with the implication of the counterfactual response of RoW industrial production in Figure 6 that dollar appreciation is on balance contractionary in the context of global risk shocks because it has larger effects through financial than through trade channels.[16] Interestingly, our finding in the context of global risk shocks is analogous to that of Degasperi et al. (2020) that financial channels dominate trade channels in the international transmission of US monetary policy shocks.

[15] 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. However, Figure A.12 in the Online Appendix 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 7 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.

[16] Liu et al. (2017) estimate factor-augmented VAR models and find that innovations to the dollar effective exchange rate ordered last are followed by a slowdown in real activity in South Korea, and argue this is due to negative US demand effects outweighing expenditure switching. While they do not consider the role of the financial channel as a competing explanation, they find that dollar appreciation hardly affects China and Japan, both of which are arguably not dependent on cross-border bank credit. Shousha (2019) estimates a panel VAR model for EMEs and finds that innovations to the dollar ordered last among US variables are followed by contractions in output abroad, which are deeper when a greater share of credit to the non-financial private sector is denominated in dollar; but it is not clear to what extent the shocks considered reflect global risk or other structural shocks. Hofmann et al. (2020) study whether the financial channel offsets the trade channel in reduced-form regressions.

Figure 7: 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 6.

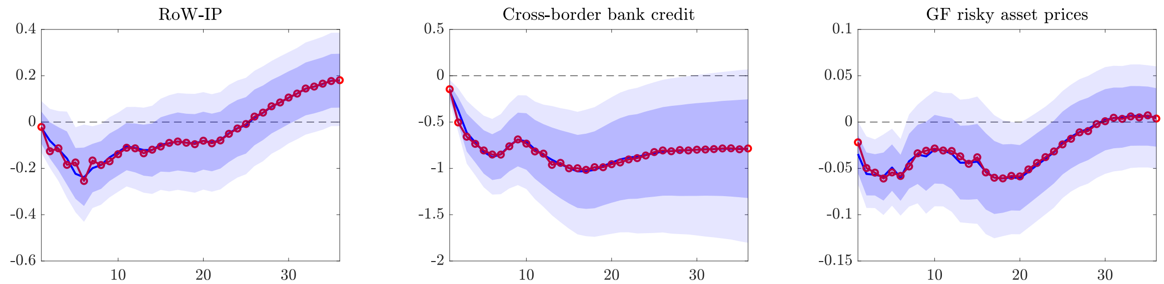
5.3 The special role of the dollar exchange rate

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 8 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 alter the transmission of global risk shocks to the RoW in the way the dollar does.

5.4 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

Figure 8: Baseline and MRE-based counterfactual responses when Japanese yen instead of US dollar is unresponsive to a global risk shock



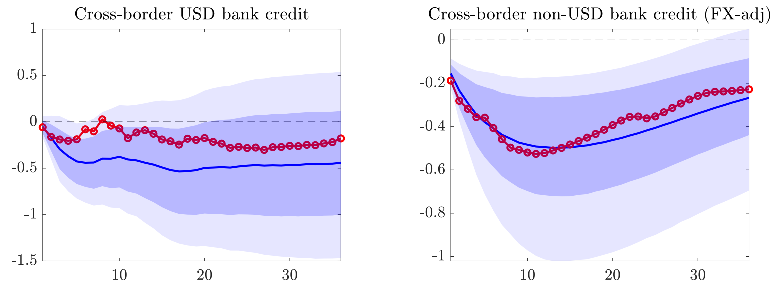
Note: See the notes to Figure 6. The red dotted lines depict the responses of rest-of-the-world industrial production (first panel), cross-border bank credit (middle panel) and the global factor in risky asset prices of Miranda-Agrippino & Rey (2020) in the counterfactual in which the Japanese yen is constrained to not respond to the global risk shock, while the response of the dollar is constrained to be identical to the baseline.

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 a 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 9 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 9: 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 6. 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.

6 What if US monetary policy stabilized the dollar?

During the COVID-19 pandemic the Federal Reserve provided unprecedented emergency liquidity to a number of countries through various facilities. It is widely believed that these measures were crucial to prevent a global financial crisis (see Cetorelli et al. 2020).[17] Theoretically, Federal Reserve swap lines can be conceived as increasing the supply of safe dollar 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 approximate this policy experiment by deviations from the past policy rule in terms of monetary policy shocks. We first lay out the technicalities and then present results.

We follow Antolin-Diaz et al. (2021; henceforth ADPRR) and consider a 'structural shock counterfactual' (SSC). In contrast to the MRE approach that determines counterfactual impulse responses by tilting the baseline posterior distribution, SSC determines a realization of shocks over the impulse response horizon whose effects offset those of the risk shock on the dollar exchange rate. More specifically, iterate forward the VAR model in Equation (1) to obtain

$$\boldsymbol{y}_{T+1,T+h} = \boldsymbol{b}_{T+1,T+h} + \boldsymbol{M}'\boldsymbol{\epsilon}_{T+1,T+h}, \tag{14}$$

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

[17]Interestingly, centralbanking.com honoured the Federal Reserve with its '2020 Central Bank of the Year Award' in particular because of the "overwhelming Fed interventions in March 2020 [that] forestalled a damaging global financial crisis".

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 (14) 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.

ADPRR 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

$$\overline{\boldsymbol{C}}\widetilde{\boldsymbol{y}}_{T+1,T+h} = \overline{\boldsymbol{C}}\boldsymbol{M}'\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\overline{\boldsymbol{f}}_{T+1,T+h}, \overline{\boldsymbol{\Omega}}_f), \tag{15}$$

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

$$\boldsymbol{\Xi}\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\boldsymbol{g}_{T+1,T+h}, \boldsymbol{\Omega}_g), \tag{16}$$

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.[18] ADPRR show how to obtain the SSA solution

$$\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\boldsymbol{\mu}_\epsilon, \boldsymbol{\Sigma}_\epsilon), \tag{17}$$

that satisfies the counterfactual constraint in Equation (15) and the constraint on the structural shocks in Equation (16). The SSC impulse response is then given by $\widetilde{\boldsymbol{y}}_{T+1,T+h} = \boldsymbol{M}'\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}$. Given our focus, we use US monetary policy shocks to offset the effect of the global risk shock on the dollar.

Figure 10 presents the results from the SSC in which US monetary policy responds to the global risk shock in contrast to past regularities in the data so as to prevent dollar appreciation.[19] In the counterfactual, US monetary policy is loosened much more than in the baseline. The slowdown in RoW real activity in the counterfactual is muted substantially relative to the baseline. At the same time, non-trivial pressures on US consumer prices and overshooting in US real activity result. Given the Federal Reserve's domestic mandate, this counterfactual outcome may represent an important practical obstacle to such a policy response. As we compare Figure 10 to the result for the MRE-based counterfactual shown in Figure 6, we note that the overall picture is quite similar. However, we note that while a counterfactual based on a more accommodating US monetary policy

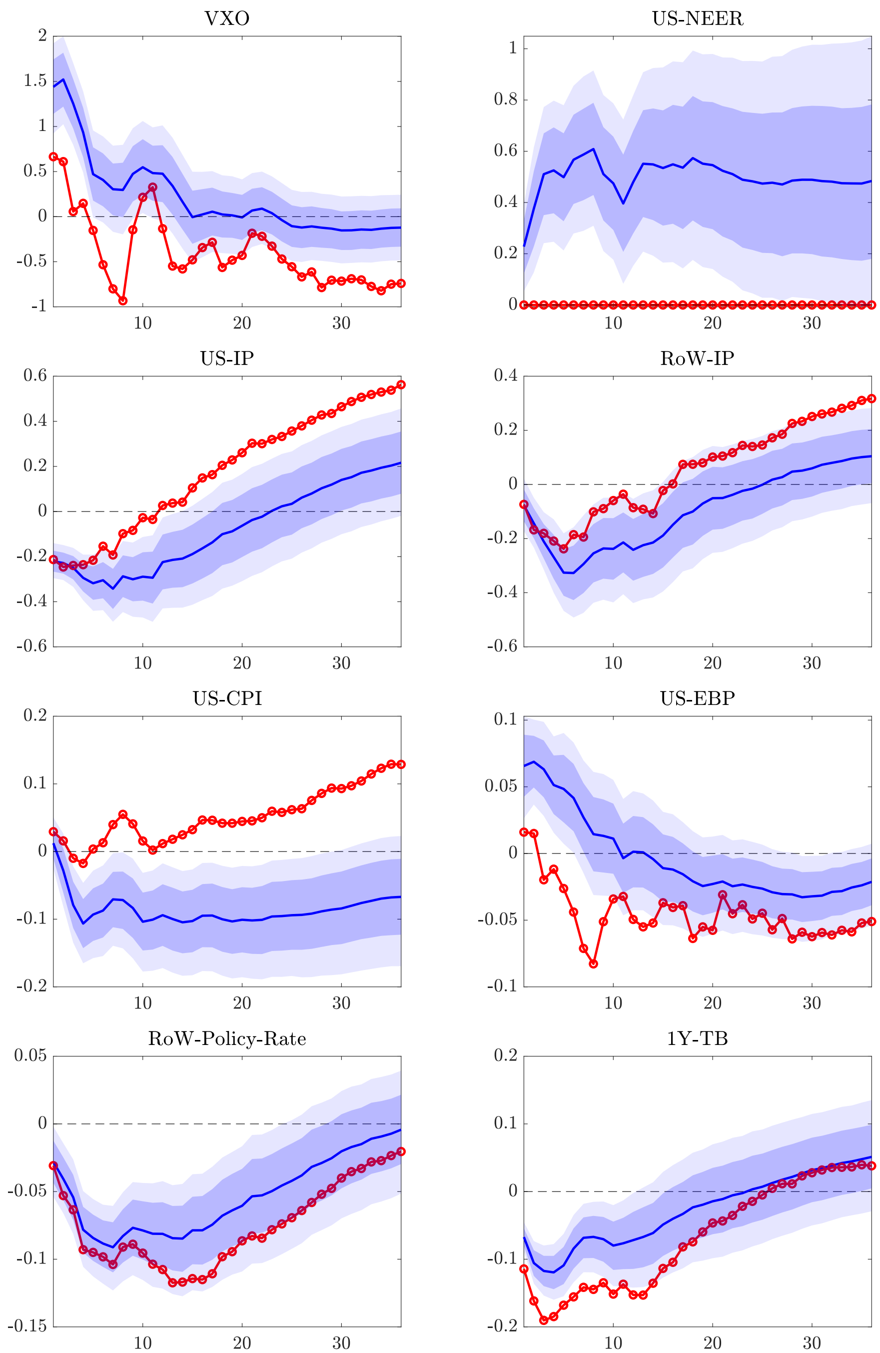
[18]For example, if the dollar exchange rate is ordered last in the vector of endogenous variables $\boldsymbol{y}_t$ in the VAR model and we impose in the counterfactual that it remains at its baseline along the impulse response horizon without any uncertainty, then in Equation (15) we have

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

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

[19]The impulse responses to a US monetary policy shock are shown in Figure A.13 in the Online Appendix. They are consistent with existing evidence for the domestic effects in the US (Gertler & Karadi 2015; Miranda-Agrippino & Rey 2020) and for spillovers to the rest of the world (Georgiadis 2016; Dedola et al. 2017; Degasperi et al. 2020).

Figure 10: Baseline and SSC responses to a global risk shock based on US monetary policy shocks as offsetting shocks



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

stance has similar effects on the RoW—after all, it keeps the dollar from appreciation just like in the MRE-based counterfactual—it is relatively more expansionary in the US.

7 Conclusion

In this paper we document that global risk shocks cause a slowdown in world real activity, a tightening in global financial conditions, an increase in the US Treasury premium, foreign holdings of US Treasury securities, banks' dollar asset liquidity ratio and the share of dollar-denominated in total international debt securities. The dollar exchange rate appreciates in response to global risk shocks; also other safe-haven currencies such as the Japanese yen or the Swiss franc appreciate, while non-safe haven currencies such as the euro or the British pound depreciate. These results lend support to theoretical models that aim to account for the special role of the dollar exchange rate and dollar assets (Farhi & Gabaix 2016; Bianchi et al. 2021; Jiang et al. 2021a).

We also assess whether dollar appreciation mitigates or exacerbates the effects of global risk shocks in the RoW. In particular, we explore counterfactuals in which the dollar does not respond to a global risk shock. We find that the contractionary financial channel dominates the expansionary trade channel: without dollar appreciation, the slowdown in RoW real activity is less pronounced than in the baseline. Indeed, in the counterfactual the response of US net exports is hardly different compared to the baseline, while the tightening in global financial conditions is much weaker. This finding also implies that global risk shocks affect the US and the RoW economies fairly symmetrically only because they induce the dollar to appreciate. Absent the appreciation, the burden of the shock would fall disproportionately on the US. Our results thus illustrate how the dollar's special role critically shapes international adjustment mechanisms.

References

Akinci, O., & Queralto, A. (2019). Exchange Rate Dynamics and Monetary Spillovers with Imperfect Financial Markets. *Federal Reserve Bank of New York Staff Reports*, *849*.

Alessandri, P., Gazzani, A., & Vicondoa, A. (2020). Uncertainty Matters: Evidence from a High-frequency Identification Strategy. *Bank of Italy Temi di discussione*, *1284*.

Angelini, G., Bacchiocchi, E., Caggiano, G., & Fanelli, L. (2019). Uncertainty Across Volatility Regimes. *Journal of Applied Econometrics*, *34*(3), 437-455.

Antolin-Diaz, J., Petrella, I., & Rubio Ramírez, J. (2021). Structural Scenario Analysis with SVARs. *Journal of Monetary Economics*, *117*, 798-815.

Aoki, K., Benigno, G., & Kiyotaki, N. (2018). Monetary and Financial Policies in Emerging Markets. *mimeo*.

Arias, J., Rubio Ramírez, J., & Waggoner, D. (forthcoming). Inference in Bayesian Proxy-SVARs. *Journal of Econometrics*.

Arias, J., Rubio-Ramírez, J. F., & Waggoner, D. (2018). Inference Based on Structural Vector Autoregressions Identified with Sign and Zero Restrictions: Theory and Applications. *Econometrica*, *86*(2), 685–720.

Avdjiev, S., Bruno, V., Koch, C., & Shin, H.-S. (2019). The Dollar Exchange Rate as a Global Risk Factor: Evidence from Investment. *IMF Economic Review*, *67*(1), 151-173.

Avdjiev, S., Du, W., Koch, C., & Shin, H.-S. (2019). The Dollar, Bank Leverage, and Deviations from Covered Interest Parity. *American Economic Review: Insights*, *1*(2), 193-208.

Avdjiev, S., Gambacorta, L., Goldberg, L., & Schiaffi, S. (2020). The Shifting Drivers of Global Liquidity. *Journal of International Economics*, *125*(C).

Bachmann, R., & Sims, E. (2012). Confidence and the Transmission of Government Spending Shocks. *Journal of Monetary Economics*, *59*(3), 235-249.

Baker, S., Bloom, N., & Davis, S. (2016). Measuring Economic Policy Uncertainty. *Quarterly Journal of Economics*, *131*(4), 1593–1636.

Banerjee, R., Devereux, M., & Lombardo, G. (2016). Self-oriented Monetary Policy, Global Financial Markets and Excess Volatility of International Capital Flows. *Journal of International Money and Finance*, *68*, 275–297.

Banerjee, R., Hofmann, B., & Mehrotra, A. (2020). Corporate Investment and the Exchange Rate: The Financial Channel. *BIS Working Paper*, *839*.

Basu, S., & Bundick, B. (2017). Uncertainty Shocks in a Model of Effective Demand. *Econometrica*, *85*(3), 937–958.

Baumeister, C., & Hamilton, J. (2015). Sign Rrestrictions, Structural Vector Autoregressions, and Useful Prior Information. *Econometrica*, *83*(5), 1963-1999.

Baur, D., & McDermott, T. (2010). Is Gold a Safe Haven? International Evidence. *Journal of Banking & Finance*, *34*(8), 1886-1898.

Bekaert, G., Engstrom, E., & Xu, N. R. (forthcoming). The Time Variation in Risk Appetite and Uncertainty. *Management Science*.

Bekaert, G., Hoerova, M., & Lo Duca, M. (2013). Risk, Uncertainty and Monetary Policy. *Journal of Monetary Economics*, *60*(7), 771-788.

Bhattarai, S., Chatterjee, A., & Park, W. (2020). Global Spillover Effects of US Uncertainty. *Journal of Monetary Economics*, *114*(C), 71-89.

Bianchi, J., Bigio, S., & Engel, C. (2021). Scrambling for Dollars: International Liquidity, Banks and Exchange Rates. *mimeo*.

BIS. (2020). Global coverage of BIS Locational Banking Statistics.

Bloom, N. (2009). The Impact of Uncertainty Shocks. *Econometrica*, *77*(3), 623–685.

Bobasu, A., Geis, A., Quaglietti, L., & Ricci, M. (2021). Tracking Global Economic Uncertainty: Implications for the Euro Area. *ECB Working Paper*, *2541*.

Breitenlechner, M., Georgiadis, G., & Schumann, B. (2021). What Goes Around Comes Around: How Large are Spillbacks from US Monetary Policy? *GRU Working Paper*, *003*.

Bruno, V., & Shin, H.-S. (2015). Cross-Border Banking and Global Liquidity. *Review of Economic Studies*, *82*(2), 535-564.

Bruno, V., & Shin, H.-S. (2021). Dollar and Exports. *CEPR Discussion Paper*, *16311*.

Caballero, R., Farhi, E., & Gourinchas, P.-O. (2008). An Equilibrium Model of "Global Imbalances" and Low Interest Rates. *American Economic Review*, *98*(1), 358-393.

Caballero, R., Farhi, E., & Gourinchas, P.-O. (2020). Global Imbalances and Policy Wars at the Zero Lower Bound. *mimeo*.

Caldara, D., & Herbst, E. (2019). Monetary Policy, Real Activity, and Credit Spreads: Evidence from Bayesian Proxy SVARs. *American Economic Journal: Macroeconomics*, *11*(1).

Calvo, G., & Reinhart, C. (2002). Fear of Floating. *Quarterly Journal of Economics*, *117*(2), 379-408.

Caramichael, J., Gopinath, G., & Liao, G. (2021). U.S. Dollar Currency Premium in Corporate Bonds. *IMF Working Paper*, *2021/185*.

Carriere-Swallow, Y., & Cespedes, L. (2013). The Impact of Uncertainty Shocks in Emerging Economies. *Journal of International Economics*, *90*(2), 316-325.

Carriero, A., Clark, T., & Marcellino, M. (forthcoming). The Identifying Information in Vector Autoregressions with Time-varying Volatilities: An Application to Endogenous Uncertainty. *Journal of Econometrics*.

Cesa-Bianchi, A., Ferrero, A., & Rebucci, A. (2018). International Credit Supply Shocks. *Journal of International Economics*, *112*(C), 219-237.

Cetorelli, N., Goldberg, L., & Ravazzo, F. (2020, May). Have the Fed Swap Lines Reduced Dollar Funding Strains during the COVID-19 Outbreak? *New York Fed Liberty Street Economics Blog*.

Coeurdacier, N., Guibaud, S., & Jin, K. (2015). Credit Constraints and Growth in a Global Economy. *American Economic Review*, *105*(9), 2838-2881.

Cogley, T., Morozov, S., & Sargent, T. J. (2005). Bayesian Fan Charts for UK Inflation: Forecasting and Sources of Uncertainty in an Evolving Monetary System. *Journal of Economic Dynamics and Control*, *29*(11), 1893–1925.

Corsetti, G., Kuester, K., Müller, G., & Schmidt, S. (2021). The Exchange Rate Insulation Puzzle. *CEPR Discussion Paper*, *15689*.

De Bock, R., & de Carvalho Filho, I. (2015). The Behavior of Currencies During Risk-off Episodes. *Journal of International Money and Finance*, *53*(C), 218-234.

Dedola, L., Rivolta, G., & Stracca, L. (2017). When the Fed Sneezes, Who Gets a Cold? *Journal of International Economics*, *S1*(108), S23-S41.

Degasperi, R., Hong, S., & Ricco, G. (2020). The Global Transmission of US Monetary Policy. *Warwick Economics Research Paper Series*, *1257*.

Du, W., Im, J., & Schreger, J. (2018). The U.S. Treasury Premium. *Journal of International Economics*, *112*(C), 167-181.

Enders, Z., Müller, G., & Scholl, A. (2011). How do Fiscal and Technology Shocks Affect Real Exchange Rates?: New Evidence for the United States. *Journal of International Economics*, *83*(1), 53-69.

Engel, C., & Wu, S. P. Y. (2018). Liquidity and Exchange Rates: An Empirical Investigation. *NBER Working Paper*, *25397*.

Epstein, B., Finkelstein Shapiro, A., & Gonzalez Gomez, A. (2019). Global Financial Risk, Aggregate Fluctuations, and Unemployment Dynamics. *Journal of International Economics*, *118*(C), 351-418.

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

Favara, G., Gilchrist, S., Lewis, K., & Zakrajsek, E. (2016). Updating the Recession Risk and the Excess Bond Premium. *FEDS Notes*, *October*.

Fernández-Villaverde, J., Guerrón-Quintana, P., Kuester, K., & Rubio-Ramírez, J. (2015). Fiscal Volatility Shocks and Economic Activity. *American Economic Review*, *105*(11), 3352-84.

Georgiadis, G. (2016). Determinants of Global Spillovers from US Monetary Policy. *Journal of International Money and Finance*, *67*, 41-61.

Georgiadis, G. (2017). To Bi, or Not to Bi? Differences in Spillover Estimates from Bilateral and Multilateral Multi-country Models. *Journal of International Economics*, *107*(C), 1-18.

Georgiadis, G., Gräb, J., & Khalil, M. (2019). Global Value Chain Participation and Exchange Rate Pass-through. *ECB Working Paper*, *2327*.

Gertler, M., & Karadi, P. (2015). Monetary Policy Surprises, Credit Costs, and Economic Activity. *American Economic Journal: Macroeconomics*, *7*(1), 44-76.

Giacomini, R., Kitagawa, T., & Read, M. (forthcoming). Robust Bayesian Inference in Proxy SVARs. *Journal of Econometrics*.

Giacomini, R., & Ragusa, G. (2014). Theory-coherent Forecasting. *Journal of Econometrics*, *182*(1), 145-155.

Giannone, D., Lenza, M., & Primiceri, G. (2015). Prior Selection for Vector Autoregressions. *Review of Economics and Statistics*, *97*(2), 436-451.

Gilchrist, S., & Zakrajsek, E. (2012). Credit Spreads and Business Cycle Fluctuations. *American Economic Review*, *102*(4), 1692-1720.

Gleser, L. (1992). The Importance of Assessing Measurement Reliability in Multivariate Regression. *Journal of the American Statistical Association*, *87*(419), 696-707.

Gopinath, G., Boz, E., Casas, C., Diez, F., Gourinchas, P.-O., & Plagborg-Moller, M. (2020). Dominant Currency Paradigm. *American Economic Review*, *110*(3), 677-719.

Gourinchas, P.-O., Rey, H., & Govillot, N. (2017). Exorbitant Privilege and Exorbitant Duty. *Bank of Japan IMES Discussion Paper*, *10-E-20*.

Gourinchas, P.-O., Rey, H., & Truempler, K. (2012). The Financial Crisis and the Geography of Wealth Transfers. *Journal of International Economics*, *88*(2), 266-283.

Habib, M., & Stracca, L. (2015). Is There a Global Safe Haven? *International Finance*, *18*(3), 281-298.

Haddad, V., Moreira, A., & Muir, T. (forthcoming). When Selling Becomes Viral: Disruptions in Debt Markets in the COVID-19 Crisis and the Fed's Response. *Review of Financial Studies*.

Hassan, T., Schreger, J., Schwedeler, M., & Tahoun, A. (2021). Country Risk. *Institute for New Economic Thinking Working Paper*, *157*.

Hofmann, B., Shim, I., & Shin, H.-S. (2020). Bond Risk Premia and the Exchange Rate. *Journal of Money, Credit and Banking*, *52*(S2), 497-520.

Ivashina, V., Scharfstein, D., & Stein, J. (2015). Dollar Funding and the Lending Behavior of Global Banks. *Quarterly Journal of Economics*, *130*(3), 1241-1281.

Jarociński, M., & Karadi, P. (2020). Deconstructing Monetary Policy Surprises: The Role of Information Shocks. *American Economic Journal: Macroeconomics*, *12*(2), 1–43.

Jentsch, C., & Lunsford, K. (2019). The Dynamic Effects of Personal and Corporate Income Tax Changes in the United States: Comment. *American Economic Review*, *109*(7), 2655-78.

Jiang, Z., Krishnamurthy, A., & Lustig, H. (2021a). Dollar Safety and the Global Financial Cycle. *mimeo*.

Jiang, Z., Krishnamurthy, A., & Lustig, H. (2021b). Foreign Safe Asset Demand and the Dollar Exchange Rate. *Journal of Finance*, *76*(3), 1049-1089.

Jurado, K., Ludvigson, S., & Ng, S. (2015). Measuring Uncertainty. *American Economic Review*, *105*(3), 1177-1216.

Kalemli-Özcan, S. (2019). U.S. Monetary Policy and International Risk Spillovers. *Proceedings of the Jackson Hole Symposium*.

Kekre, R., & Lenel, M. (2021). The Flight to Safety and International Risk Sharing. *mimeo*.

Krishnamurthy, A., & Lustig, H. (2019). Mind the Gap in Sovereign Debt Markets: The U.S. Treasury Basis and the Dollar Risk Factor. *Proceedings of the Jackson Hole Symposium*.

Krishnamurthy, A., & Vissing-Jorgensen, A. (2012). The Aggregate Demand for Treasury Debt. *Journal of Political Economy*, *120*(2), 233-267.

Lakdawala, A. (2019). Decomposing the Effects of Monetary Policy Using an External Instruments SVAR. *Journal of Applied Econometrics*, *34*(6), 934-950.

Leduc, S., & Liu, Z. (2016). Uncertainty Shocks are Aggregate Demand Shocks. *Journal of Monetary Economics*, *82*, 20-35.

Liao, G. (2020). Credit Migration and Covered Interest Rate Parity. *Journal of Financial Economics*, *138*(2), 504-525.

Lilley, A., Maggiori, M., Neiman, B., & Schreger, J. (forthcoming). Exchange Rate Reconnect. *Review of Economics and Statistics*.

Liu, Z., Spiegel, M., & Tai, A. (2017). Measuring the Effects of Dollar Appreciation on Asia: A FAVAR Approach. *Journal of International Money and Finance*, *74*(C), 353-370.

Londono, J., & Wilson, B. (2018). Understanding Global Volatility. *FEDS Notes*, *January*.

Ludvigson, S., Ma, S., & Ng, S. (forthcoming). Uncertainty and Business Cycles: Exogenous Impulse or Endogenous Response? *American Economic Journal: Macroeconomics*.

Lustig, H., Roussanov, N., & Verdelhan, A. (2014). Countercyclical Currency Risk Premia. *Journal of Financial Economics*, *111*(3), 527-553.

Maggiori, M. (2017). Financial Intermediation, International Risk Sharing, and Reserve Currencies. *American Economic Review*, *107*(10), 3038-3071.

Martínez-García, E., Grossman, V., & Mack, A. (2015). A Contribution to the Chronology of Turning Points in Global Economic Activity (1980–2012). *Journal of Macroeconomics*, *46*, 170-185.

Meisenzahl, R., Niepmann, F., & Schmidt-Eisenlohr, T. (2019). The Dollar and Corporate Borrowing Costs. *CEPR Discussion Paper*, *14892*.

Mendoza, E., Quadrini, V., & Rios-Rull, J.-V. (2009). Financial Integration, Financial Development, and Global Imbalances. *Journal of Political Economy*, *117*(3), 371-416.

Mertens, K., & Ravn, M. (2013). The Dynamic Effects of Personal and Corporate Income Tax Changes in the United States. *American Economic Review*, *103*(4), 1212–47.

Miranda-Agrippino, S., Nenova, T., & Rey, H. (2020). Global Footprints of Monetary Policy. *CEPR Discussion Paper*, *13853*.

Miranda-Agrippino, S., & Rey, H. (2020). U.S. Monetary Policy and the Global Financial Cycle. *Review of Economic Studies*, *87*(6), 2754-2776.

Miranda-Agrippino, S., & Rey, H. (2021). The Global Financial Cycle. *CEPR Discussion Paper*, *16580*.

Miranda-Agrippino, S., & Ricco, G. (2021). The Transmission of Monetary Policy Shocks. *American Economic Journal: Macroeconomics*, *13*(3), 74-107.

Montiel Olea, J., Stock, J., & Watson, M. (forthcoming). Inference in Structural Vector Autoregressions Identified with an External Instrument. *Journal of Econometrics*.

Moon, H., & Schorfheide, F. (2012). Bayesian and Frequentist Inference in Partially Identified Models. *Econometrica*, *80*(2), 755–782.

Niepmann, F., & Schmidt-Eisenlohr, T. (2017). Foreign Currency Loans and Credit Risk: Evidence from U.S. Banks. *CESifo Working Paper*, *6700*.

Noeth, B., & Sengupta, R. (2010). Flight to Safety and U.S. Treasury Securities. *The Regional Economist*, *July*, 18-19.

Obstfeld, M., & Rogoff, K. (1996). *Foundations of International Macroeconomics.* The MIT Press.

Piffer, M., & Podstawski, M. (2018). Identifying Uncertainty Shocks Using the Price of Gold. *Economic Journal*, *128*(616), 3266-3284.

Poirier, D. (1998). Revising Beliefs In Nonidentified Models. *Econometric Theory*, *14*(4), 483-509.

Ranaldo, A., & Söderlind, P. (2010). Safe Haven Currencies. *Review of Finance*, *14*(3), 385-407.

Redl, C. (2020). Uncertainty Matters: Evidence from Close Elections. *Journal of International Economics*, *124*(C).

Rey, H. (2016). International Channels of Transmission of Monetary Policy and the Mundellian Trilemma. *IMF Economic Review*, *64*(1), 6-35.

Robertson, J., Tallman, E., & Whiteman, C. (2005). Forecasting Using Relative Entropy. *Journal of Money, Credit and Banking*, *37*(3), 383-401.

Rubio-Ramirez, J., Waggoner, D., & Zha, T. (2010). Structural Vector Autoregressions: Theory of Identification and Algorithms for Inference. *Review of Economic Studies*, *77*(2), 665–696.

Shim, I., Kalemli-Özcan, S., & Liu, X. (2021). Exchange Rate Fluctuations and Firm Leverage. *IMF Economic Review*, *69*, 90–121.

Shousha, S. (2019). The Dollar and Emerging Market Economies: Financial Vulnerabilities Meet the International Trade System. *International Finance Discussion Papers*, *1258*.

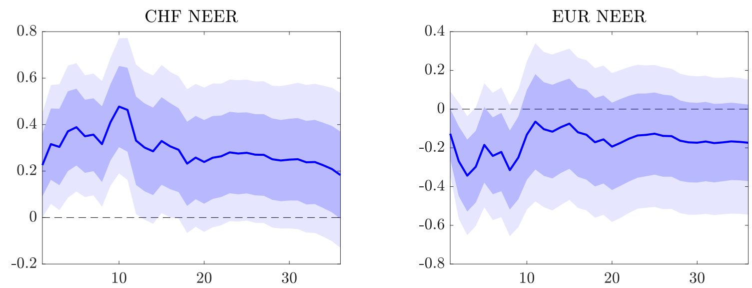
Stock, J. H., & Watson, M. (2012). Disentangling the Channels of the 2007-09 Recession. *Brookings Papers on Economic Activity*, *43*(1), 81-156.

Verdelhan, A. (2018). The Share of Systematic Variation in Bilateral Exchange Rates. *Journal of Finance*, *73*(1), 375-418.

Vicondoa, A. (2019). Monetary News in the United States and Business Cycles in Emerging Economies. *Journal of International Economics*, *117*(C), 79-90.

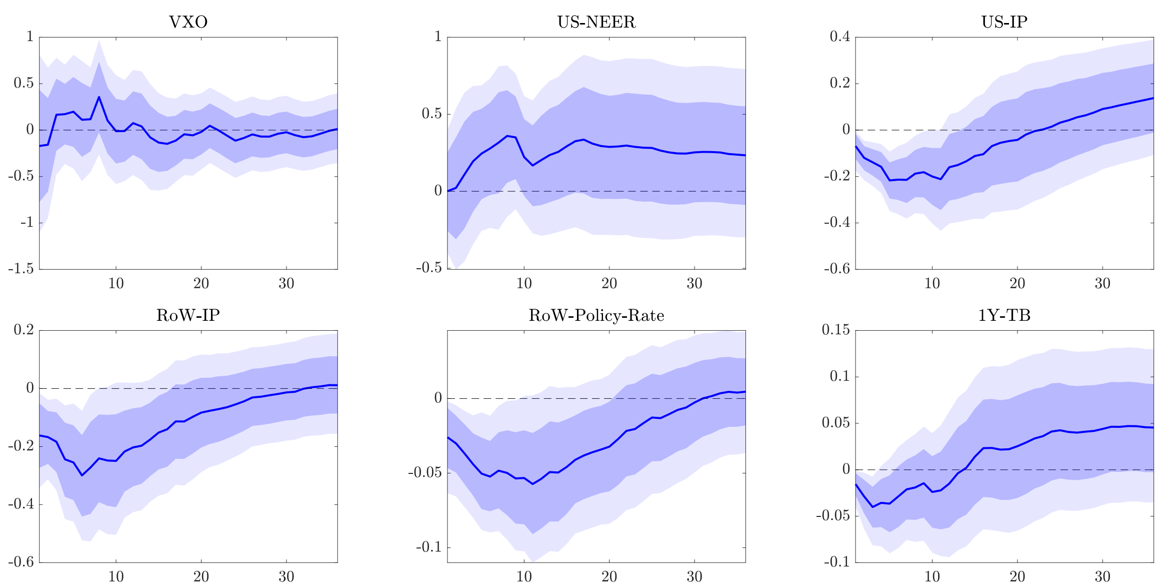
A Online appendix - Additional figures

Figure A.1: Impulse responses of Swiss franc and euro exchange rates to global risk shock



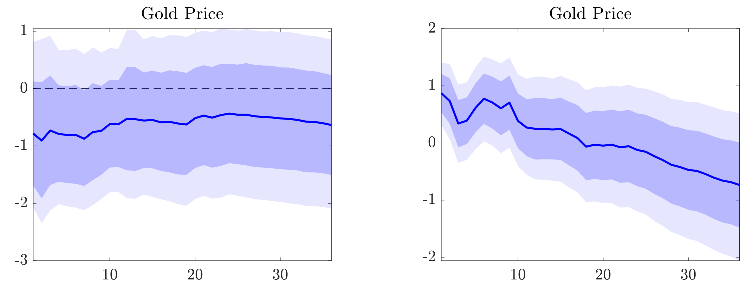
Note: See notes to Figure 2. Responses are obtained from re-estimating the baseline BPSVAR model with the vector $\boldsymbol{y}_t$ augmented with one additional variable at a time.

Figure A.2: Impulse responses to a global demand shock



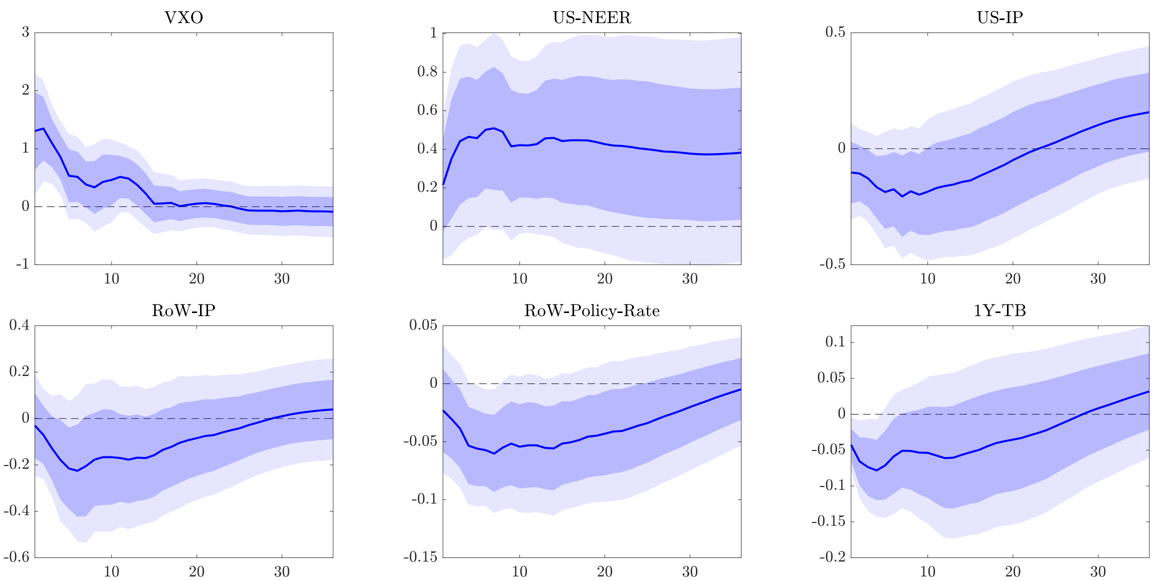
Note: The figure presents the impulse responses to a one-standard deviation global demand shock identified based on sign restrictions. See also the notes to Figure 2. Impulse responses of US CPI and the EBP are omitted to save space.

Figure A.3: Impulse responses of the gold price to global demand (left panel) and global risk (right panel) shocks



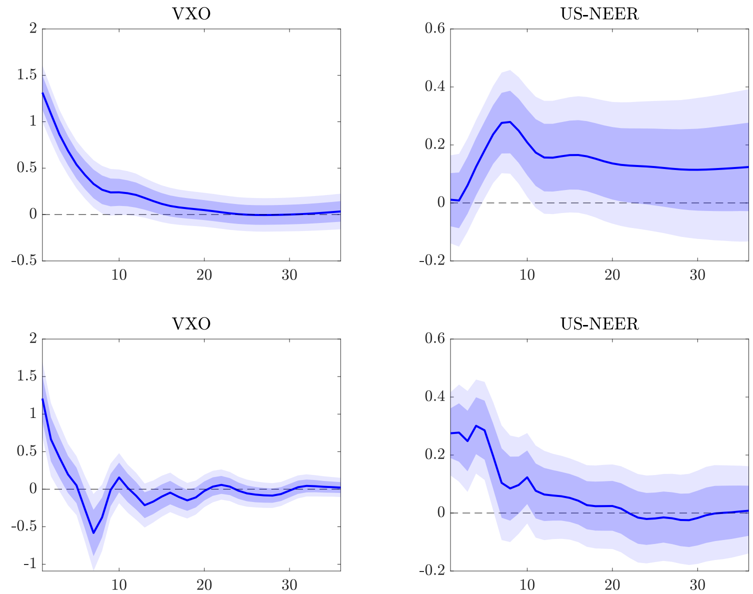
Note: See the notes to Figure A.2.

Figure A.4: Impulse responses to a global risk shock when allowing the gold price surprises to be correlated with all structural shocks



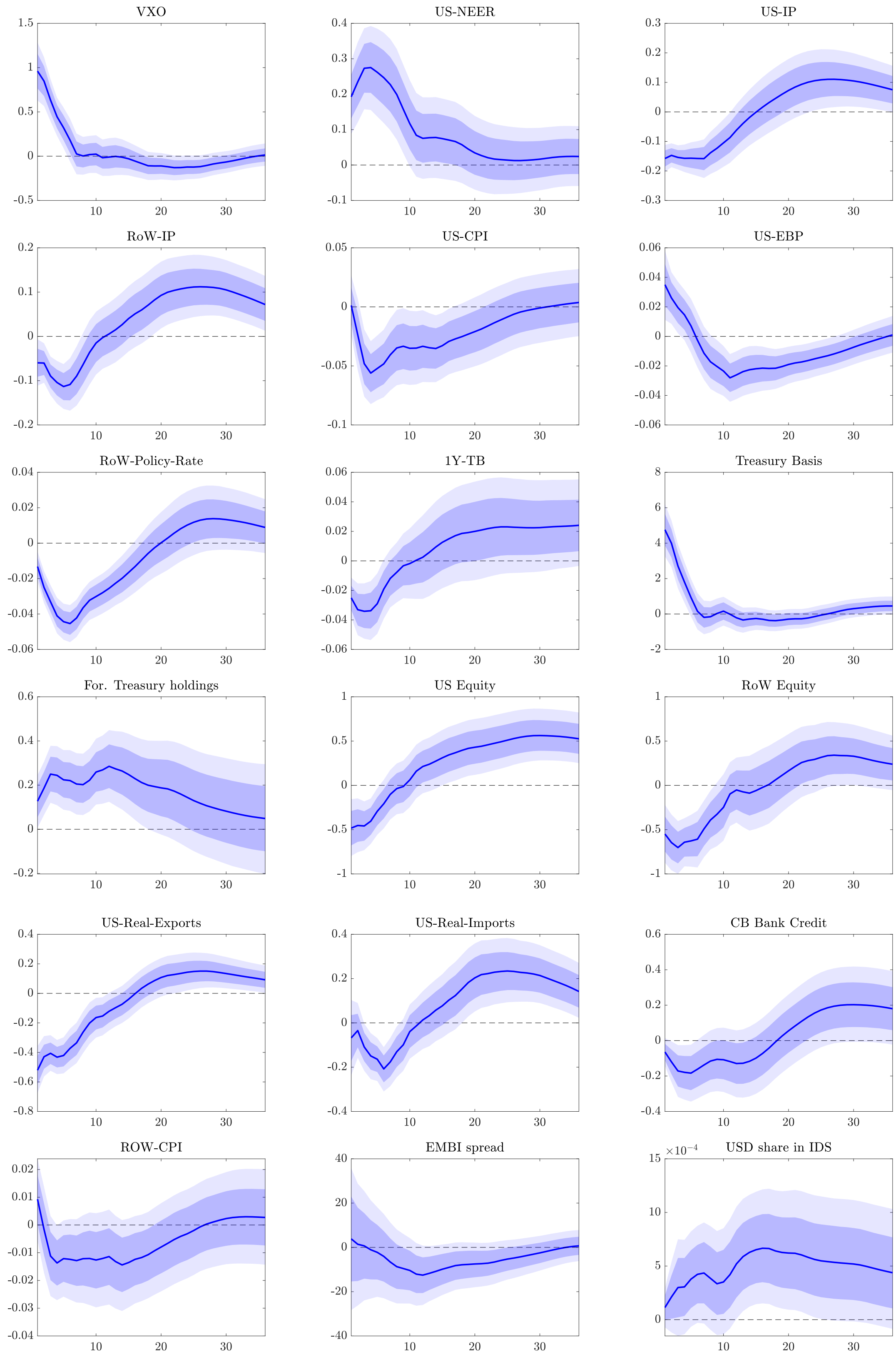
Note: The figure presents the impulse responses to a one-standard deviation global risk shock based on an alternative identification scheme in which the gold price surprises are allowed to be correlated with all structural shocks, imposing only that the correlation is strongest with the global risk shock.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

Figure A.5: Impulse responses to a global risk shock for the sample periods 1990-2006 (top row) and 2007-2019 (bottom row)



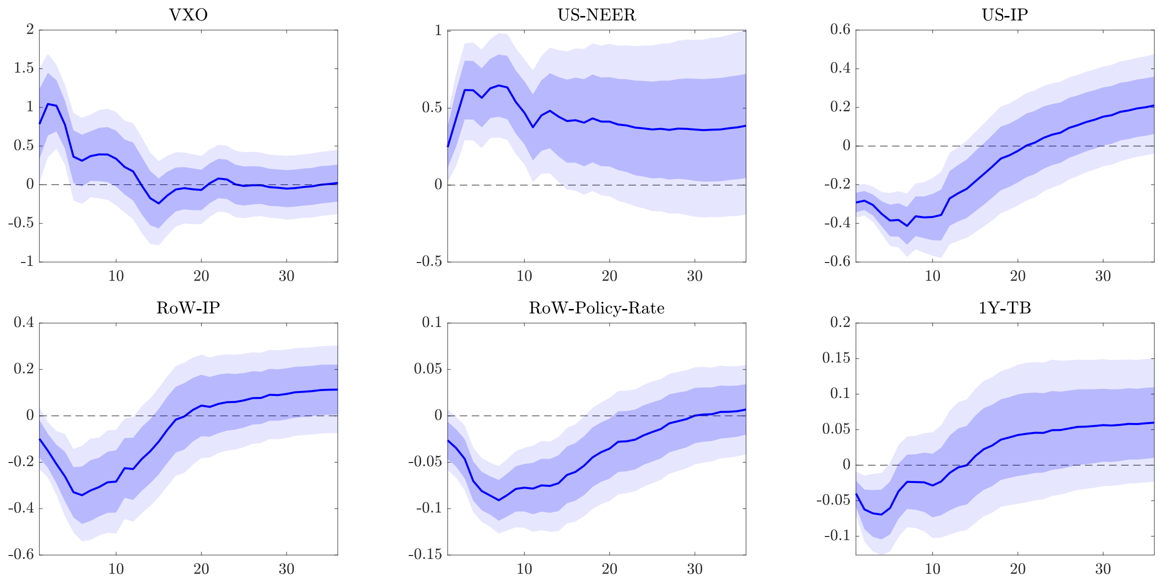
Note: The figure presents the impulse responses to a one-standard deviation global risk shock for the sample periods from 1990-2007 (top row) and 2007 to 2019 (bottom row). Due to the short sample period estimation uses informative Minnesota-type priors and optimal hyperpriors/prior tightness as suggested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

Figure A.6: Impulse responses to a global risk shock from a large BPSVAR model



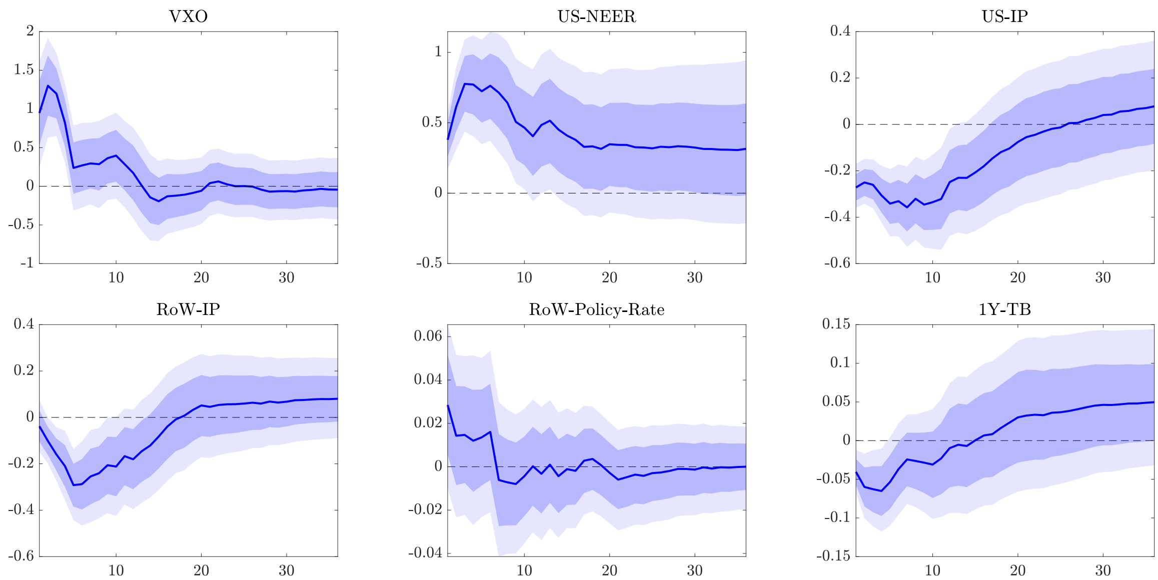
Note: See notes to Figure 6. The model is estimated with informative Minnesota-type priors and optimal hyperpriors/prior tightness as in Giannone et al. (2015). We do not include the liquidity ratio in the VAR model because it is only available for a substantially shorter sample period (see Table B.1).

Figure A.7: Impulse responses to a global risk shock when considering also 'European' and 'other' risk events in the construction of the time series of gold price surprises



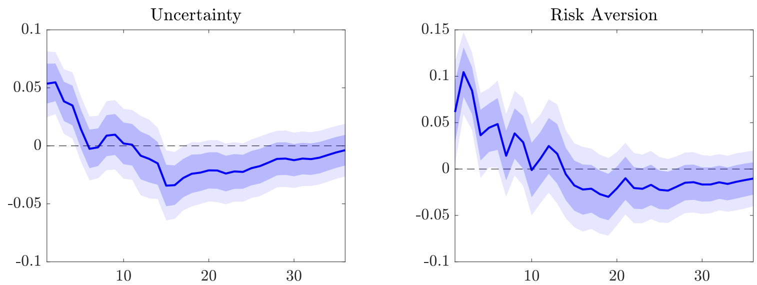
Note: See notes to Figure 6. Results are obtained from a BPSVAR model with gold price surprises also on risk events labelled as 'European' and 'other' risk events by Piffer & Podstawski (2018). Impulse responses of US CPI and the EBP are omitted to save space.

Figure A.8: Impulse responses to a global risk shock when considering only risk events associated with positive gold price surprises



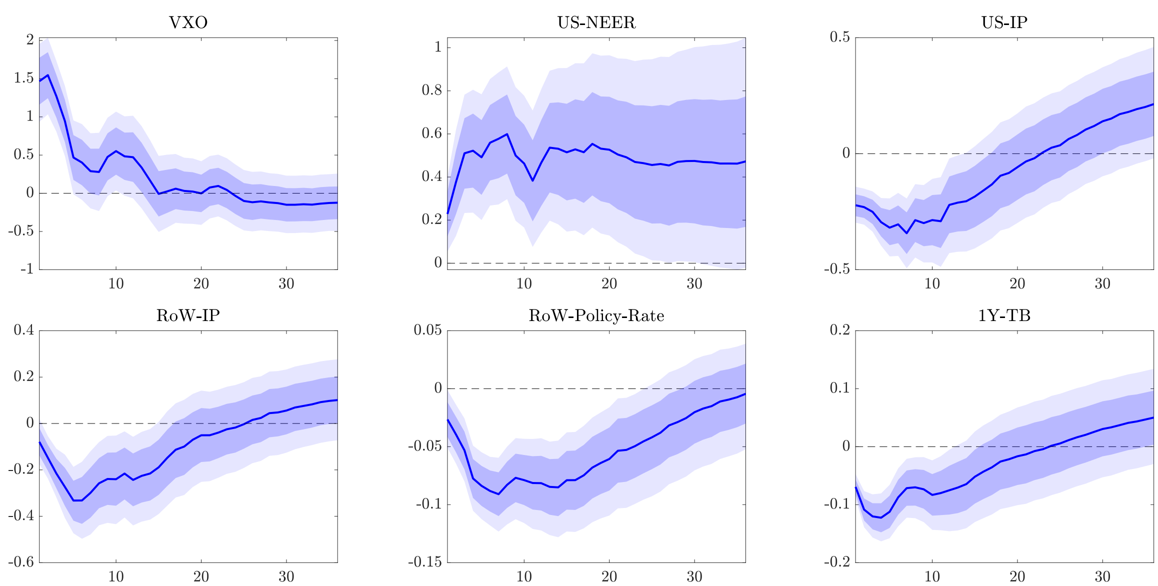
Note: See notes to Figure 6. The results are obtained from a BPSVAR model with only positive gold price surprises. Impulse responses of US CPI and the EBP are omitted to save space.

Figure A.9: Impulse responses of risk aversion and uncertainty a global risk shock



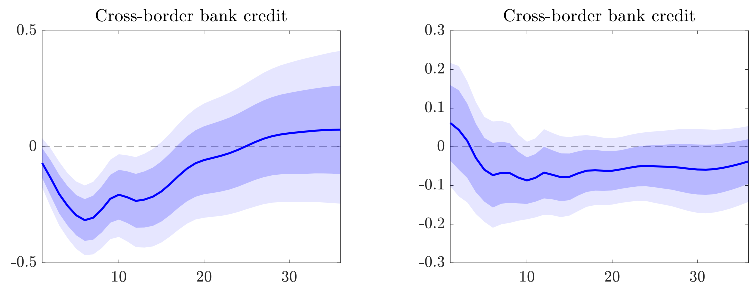
Note: See notes to Figure 2. Responses are obtained from re-estimating the baseline BPSVAR model with the vector $\boldsymbol{y}_t$ *augmented with one additional variable at a time.*

Figure A.10: Impulse responses to global risk shock when no relevance threshold is imposed



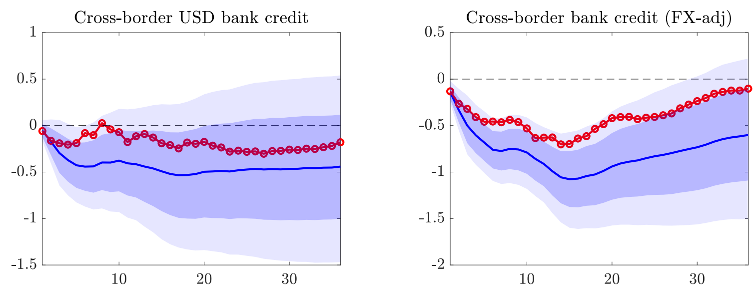
Note: The figure presents the impulse responses to a one-standard deviation global risk shock based on an alternative identification scheme in which we do not impose any relevance threshold. See also the notes to Figure 2. Impulse responses of US CPI and the EBP are omitted to save space.

Figure A.11: Impulse responses of cross-border bank credit and international debt securities to a global risk shock before 2007 (left) and from 2009 (right)



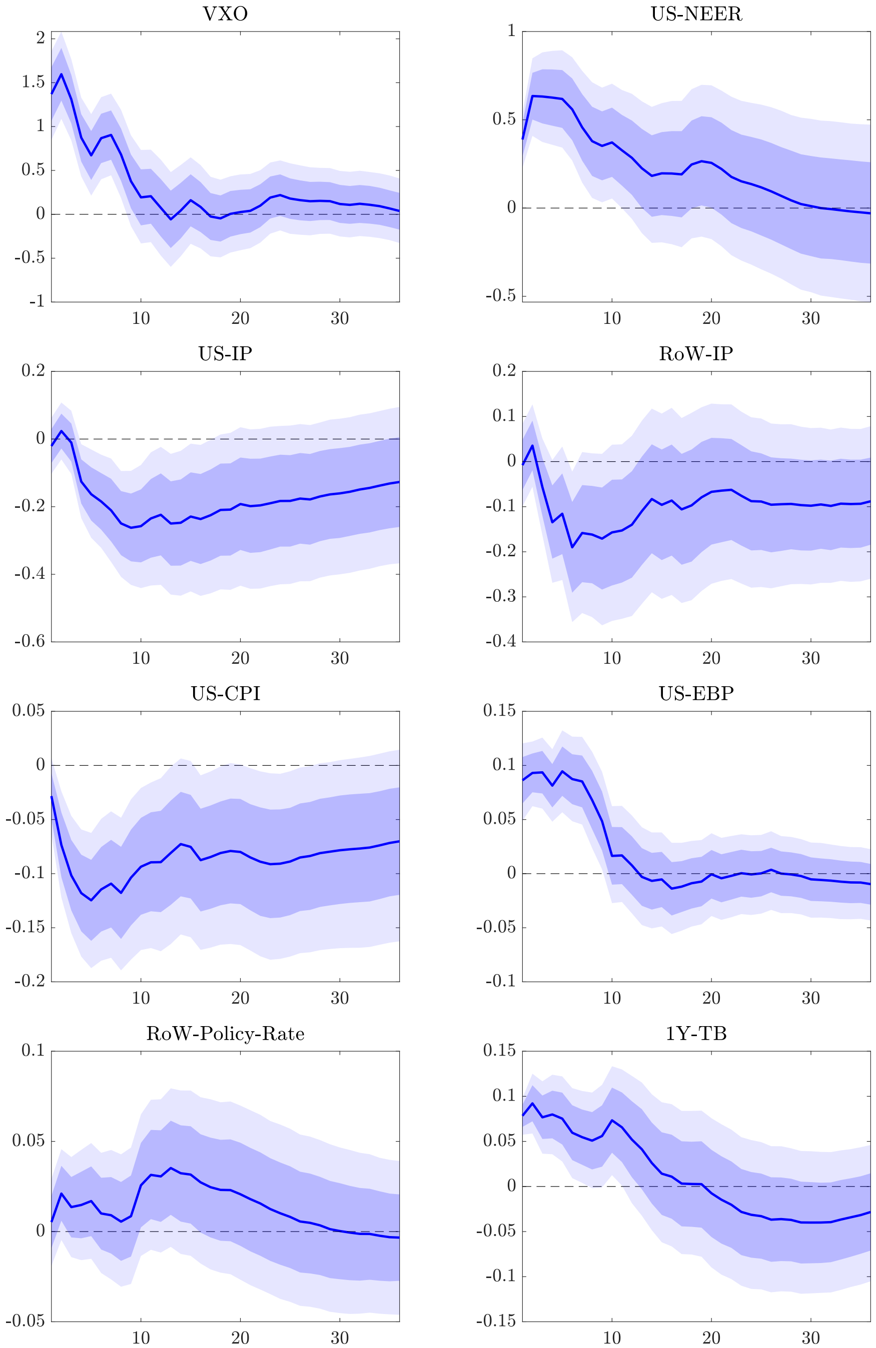
Note: The figure presents the impulse responses of cross-border bank credit to non-US borrowers for the pre and the post-GFC samples. See the notes to Figure 2.

Figure A.12: Baseline and MRE-based counterfactual responses of alternative cross-border bank credit variables



Note: See the notes to Figure 6. 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 A.13: 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.

B Online appendix - Additional tables

Table B.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 external liabilities 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 less external 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.

C 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, forthcoming), 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

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

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

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

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

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

where $\boldsymbol{y}_{T+1,T+h}^{(i)}$ are the impulse responses to a global risk shock shock as defined in Section 5. Equations (C.2) and (C.3) reflect that the weights are probabilities, and Equation (C.4) that the counterfactual posterior distribution shall satisfy some constraint.

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

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

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

$$w_{i,h}^{*} = \frac{w_{i,h} \exp\left[\lambda_h g(y_{\$,T+h}^{(i)})\right]}{\sum_{i=1}^{N} w_{i,h} \exp\left[\lambda_h g(y_{\$,T+h}^{(i)})\right]}, \quad i = 1, 2, \dots, N \tag{C.6}$$

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

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

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