Has Globalization Changed the International Transmission of U.S. Monetary Policy?*

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Abstract

We estimate a time-varying parameter vector autoregression to examine the evolution of international spillovers of U.S. monetary policy in light of increasing globalization in real and financial markets. We find that the adverse international effects of a U.S. tightening have substantially increased over the past three decades, peaking during the Great Recession before stabilizing – a timing that aligns well with observed trends in globalization and slowbalization dynamics. We observe that a sharper contraction in real economic activity is closely associated with a more pronounced decline in global trade, while the impact on global financial conditions has remained relatively stable over time. Cross-country analysis and counterfactual simulations suggest that the amplification of economic activity spillovers over time is primarily driven by the surge in trade integration, while rising financial integration contributes only modestly.

Keywords: Monetary Policy; International Spillovers; TVP-VARs.

JEL Codes: C32, E32, E52, F42, F62.

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1. Introduction

Two well-known stylized facts in international economics – portrayed in Figure 1 – read as follows. First, the global macro-financial system is centered around the U.S. dollar, which represents the dominant currency in trade invoicing and issuance of financial assets (Gopinath, 2015; Rey, 2016). This implies that U.S. interest rate hikes affect global outcomes by depressing global trade and financial conditions.[1] Second, globalization led to a massive *increase* in global trade and financial integration, which could have substantially modified the global transmission of U.S. shocks. Motivated by the interaction of these two facts, this paper examines the implications of globalization for the international transmission of U.S. monetary policy. We estimate a proxy-VAR model that allows for time-varying parameters to capture the possible change in the impact of policy disturbances on the global financial and real cycles. We do so because – as shown once again in Figure 1 – globalization has materialized at a changing pace over time, with a change in direction after the Great Recession, a phenomenon known as "slowbalization".

We document that globalization has led to substantial time variation in the international ramifications of monetary policy shocks: U.S. policy hikes generate stronger recessionary effects over time, with a flattening out of effects after the Great Recession. In 1993, a one percentage point hike led to roughly a –0.4% contraction in global industrial production, while the same shock in 2008 resulted in a downturn of about –2.5%. While there is vast evidence showing that a monetary hike engineered by the Federal Reserve (Fed) generates global recessionary effects,[2] we are the first to show the relevance of the time variation.

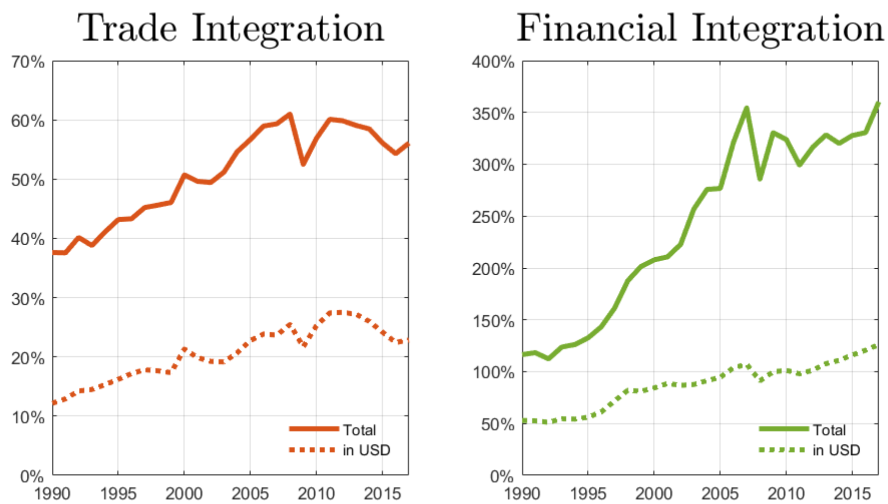
To study the effects of exogenous U.S. monetary policy shocks and their international spillovers over time, we use a time-varying parameter vector autoregression (TVP-VAR). We analyze the transmission of U.S. monetary policy mediated by the reaction of U.S. industrial production, U.S. prices, and global economic, trade, and financial indicators. To gauge the effects on global outcomes, we rely on global (rest of the world – RoW, excluding the U.S.) production and trade indices and international asset prices. We estimate our time-varying model using Bayesian techniques (Primiceri, 2005; Paul, 2020) and achieve shocks' identification using the high-frequency instrument of Miranda-Agrippino and Ricco (2021), which directly controls for the information channel of monetary policy.

The effects of U.S. monetary policy transmit globally through both trade and financial channels. The former refers to the traditional Mundell-Fleming effects, in which the current account adjusts

[1] The pivotal role of the dollar as an international currency brought many observers to view the Federal Reserve as a world banker, see for instance Gourinchas and Rey (2007), Gopinath (2015), Rey (2016), Gopinath and Stein (2018), Gourinchas et al. (2019) and Ilzetzki et al. (2019).

[2] See for instance Dedola et al. (2017), Iacoviello and Navarro (2019), Degasperi et al. (2020), Georgiadis and Schumann (2021), Breitenlechner et al. (2022), Kalemli-Özcan and Unsal (2023), Bräuning and Sheremirov (2023), and Camara et al. (2024).

Figure 1: Motivating evidence: global trade and financial integration as share of GDP.



Notes: Global (dollar) trade integration is defined as the sum of global exports and imports (invoiced in dollar) as a percentage of GDP (World Bank national accounts data and Boz et al. (2022)). Global (dollar) financial integration is defined as the sum of global external assets and liabilities (denominated in dollar) as a percentage of GDP (data from Lane and Milesi-Ferretti, 2018; Bénétrix et al., 2019).

because monetary policy affects both aggregate demand (of foreign goods) and the currency. A domestic policy tightening generates recessionary effects at home, which depress demand for imports. This demand effect should be partly counteracted by a revaluation of the currency, which leads to expenditure-switching from home goods to foreign goods. However, since U.S. imports are mainly invoiced in the dollar, a revaluation of the currency neither boosts nor dampens the competitiveness of foreign goods; thus, the expenditure-switching channel is of minor importance for the U.S. The U.S. dollar, however, is the dominant invoicing currency not only for U.S. trade but for global trade in general (Goldberg and Tille, 2008; Boz et al., 2022). Hence, expenditure-switching plays a substantial role in trade between non-U.S. countries that use the dollar as their invoicing currency. The prevalence of dollar invoicing in global trade leads to both inflation spillovers via a widespread surge in import prices as well as to negative output spillovers in response to the appreciation of the dollar (Gopinath and Neiman, 2014; Gopinath et al., 2020; Cook and Patel, 2023).[3] Employing time-series methods and small open economy models, Bräuning and Sheremirov (2023) and Camara et al. (2024) find that the decline in trade is the primary channel through which a U.S. monetary contraction impacts global economic conditions.[4] Monetary policy also operates globally through a financial channel: When the Fed raises rates, foreign currencies

[3] Georgiadis and Schumann (2021) argue that asymmetries in dollar invoicing shares between countries and imports/exports lead to higher output spillovers. In the case of a *full* dominant currency paradigm (all trade is invoiced in dollar), third-country expenditure-switching effects would nullify (in terms of output). Asymmetries then cause expenditure-switching between and within countries.

[4] Relatedly, Di Giovanni and Hale (2022) find strong international stock market spillovers to U.S. monetary policy via production networks.

depreciate against the dollar, financial constraints tighten in foreign economies, and global asset prices fall (Farhi and Werning, 2014; Bruno and Shin, 2015; Rey, 2016; Akinci and Queralto, 2024; Wu et al., 2024). This causes a strong case for an international risk channel of monetary policy. Particularly, the existence of the global financial cycle acts as a transmitter for the financial channel (Miranda-Agrippino and Rey, 2020). Again, the currency plays an important role because the dollar dominates the global financial system and is an important transmitter of global risk shocks (Georgiadis et al., 2024, 2023). Against the evidence presented, this paper is particularly interested in understanding whether and how the relevance of the *trade* and *financial* channel has changed over time due to globalization. As indicated by the motivating evidence in Figure 1, it is plausible that both channels have gained strength during globalization and disentangling them is ultimately an empirical matter.

We find strong evidence of growing global spillovers of U.S. monetary policy. The negative response of RoW industrial production has increased significantly over the last decades of rising globalization. The magnitude of the spillover only stabilizes with the onset of the Great Recession, which is consistent with the slowdown in trade and financial integration that we see in the data. The observed time variation is substantial and statistically significant. While a one percentage point hike in monetary policy leads to about a -0.4% contraction in RoW industrial production in 1993, a same-sized shock in 2008 results in a downturn of about -2.5%.[5] When we examine the transmission, we find that the monetary policy shocks generate a strong contraction in global trade and the financial cycle. However, we find a disconnect between the two variables when looking at the time variation. While the effects on RoW trade are significantly time-varying and track well the pattern we find for RoW industrial production, the ones on global financial conditions are relatively constant.

Given the observed divergence in the effects on global trade (which vary over time) and financial conditions (which remain relatively constant), our estimations point to trade integration as a possible driver of the variability in spillover effects. However, in the presence of a large and time-varying *global financial multiplier*, even minor shifts in the impact on financial conditions could explain a significant proportion of the variations in the reactions of RoW industrial production. We dig deeper into this aspect by performing a heterogeneity analysis and counterfactual simulations. Regarding the heterogeneity analysis, we generally find stronger spillover effects for emerging markets (EMEs) than advanced (AE) economies. Given that EMEs are more impacted by U.S. disturbances and their share of world economic activity has increased from approximately 20% to 40% over the past three decades (Lane, 2019), part of the observed time variation is explained *mechanically* by composition effects. We then construct a (balanced) panel of 26 countries and estimate the

[5] We note that spillovers could simply get stronger if domestic effects also get stronger over time. However, we find similar results when using a different normalization, which induces a 1% U.S. recession at each point in time (and holds the domestic effects constant).

response of country-specific industrial production to U.S. monetary policy shocks over time. We analyze the outcomes of this analysis in conjunction with country-specific financial and trade integration data. This enables us to explore the correlation between the growth of spillovers and the increasing economic integration and understand the nature of this relationship. Overall, we find that countries exhibiting greater historical levels of trade and financial integration tend to undergo more pronounced economic downturns in the aftermath of contractionary U.S. policy shocks. Both channels are active *on average*. However, when it comes to explaining the *time variation* in effects, our estimations again suggest that the trade dimension holds greater significance. While countries that have substantially enhanced their trade integration tend to experience exacerbated recessionary impacts as time progresses, no clear negative association emerges in relation to enhanced financial integration.

Finally, we also conduct counterfactual simulations to isolate trade and financial transmission channels within our VAR framework. To achieve this, we identify a *global trade shock* and a *global financial shock* that are contemporaneously orthogonal to business cycle variables and then simulate counterfactual scenarios in which U.S. monetary disturbances do not affect global trade or global financial conditions (Sims and Zha, 2006). By shutting down these two transmission channels of U.S. monetary policy shocks, we obtain the following results. The financial channel, while accounting for approximately one-third of the overall response of RoW industrial production *on average*, has only seen a modest increase in importance over time. In contrast, the trade channel exhibits a similar average contribution but has grown massively in relevance over time.

Why has the rise in spillover effects from increasing financial integration been modest? A series of facts can explain it. First, it is well documented that, since the 1990s, many countries – while increasing their overall financial integration – have shifted their net foreign currency positions from negative (more liabilities than assets) to more positive ones, progressively reducing their foreign currency exposure (Lane and Shambaugh, 2010; Allen and Juvenal, 2024).[6] This matters because when a country has a negative foreign position, a U.S. monetary tightening that appreciates the dollar leads to significant adverse balance sheet effects – reducing the local value of foreign assets relative to liabilities (Gourinchas and Rey, 2014). Thus, although countries became more financially integrated by conventional measures, a more positive foreign position makes them better able to absorb global shocks, as they face reduced valuation losses and can use currency depreciation to counter external shocks (Bénétrix et al., 2015). Second, the composition of external assets and liabilities has changed. On the liabilities side, dollar debt funding has declined (Kalemli-Özcan and Unsal, 2023), while a larger share is now in domestic currency and equity instruments (portfolio equity and foreign direct investments), which may promote international risk-sharing and

[6] Bénétrix et al. (2015) document that 65% of countries had a negative net position in 1996, declining to 36% in 2007. This shift is observed in both emerging and advanced countries (Allen and Juvenal, 2024).

stability (Bénétrix et al., 2015; Obstfeld, 2021). On the asset side, many countries have built up foreign reserves (Lane and Milesi-Ferretti, 2018), which are not marked to market and thus not subject to price valuation effects. As documented by Bussière et al. (2015), the increase in foreign reserves – that *de facto* contributes to the rise in measured financial integration – enabled domestic monetary authorities to act as lenders-of-last-resort for institutions with short-term foreign-currency liabilities, reducing the impact of adverse shocks like U.S. monetary contractions (see also Catão and Milesi-Ferretti, 2014 and Benigno et al., 2022). Finally, while conventional measures of financial integration – based on cross-border assets and liabilities – show significant increases over time, Lane and Milesi-Ferretti (2018) argue that multiple factors have contributed to the growth in cross-border holdings to an extent that overstates both the level and composition of underlying cross-border financial linkages (for instance, tax management practices and regulatory arbitrage give rise to round tripping arrangements by which foreign assets and foreign liabilities essentially offset each other, with no true cross-border financial linkage). Cross-border holdings have grown, simply "*reflecting the "passage" of investments through financial centers en route to their final destination*." In this sense, routinely used measures of financial integration may overstate the actual variation in genuine financial integration. Our evidence supports this interpretation.

Contribution to the literature. The paper contributes to the vast literature that uses linear models to document the negative effects of U.S. monetary policy hikes on the global and real financial cycles (Miranda-Agrippino and Rey, 2020, Dedola et al., 2017, Iacoviello and Navarro, 2019, Degasperi et al., 2020, Bräuning and Sheremirov, 2023, and Camara et al., 2024).[7] In contrast, we use a time-varying model and document the increasing international spillovers of U.S. policy shocks. Additionally, we contribute to the literature examining the time-varying dimension of international monetary policy. Our paper connects closely to Liu et al. (2022), who estimate a time-varying parameter model to jointly model monetary policy decisions in the U.S., U.K., and Euro area. While their study highlights time-varying network structures in central banks' decisions, we investigate the global spillovers of U.S. shocks and their underlying drivers. Crespo Cuaresma et al. (2019) estimate a time-varying global VAR using sign restrictions but do not find a clear time pattern in the effects. Ilzetzki and Jin (2021) compare the international transmission of U.S. monetary policy shocks prior to and after the 1990s. They find that, before the 1990s, world industrial production declines in response to a U.S. monetary tightening, while during the period 1990-2007, U.S. contractions are expansionary abroad. In contrast to their paper, we model the changes in international transmission channels using a time-varying model (vs. sample-splitting strategy) and focus on the dynamics *within* the post-1990s period. Furthermore, and contrary to

[7] See also Canova (2005), Maćkowiak (2007), Georgiadis (2016), Feldkircher and Huber (2016), and Ca'Zorzi et al. (2023) for analyses of the foreign responses of real activity to U.S. policy decisions. See Gerko and Rey (2017), Jordà et al. (2019), Habib and Venditti (2019), and Dées and Galesi (2021) for the transmission to international financial conditions.

their findings, we find evidence in favor of a growing (negative) role of U.S. policy tightenings for global economic activity.

Outline. The remainder of the paper proceeds as follows. Section 2 discusses the empirical strategy and specification. In section 3, we present the empirical results, including various extensions, the heterogeneity analysis, the counterfactual exercises, and a battery of sensitivity checks. Finally, section 4 concludes.

2. Empirical methodology

We examine the international spillovers of U.S. monetary policy using a medium-scale TVP-VAR model that allows for time variation in the parameters. On the domestic level, we include the federal funds rate, U.S. consumer price index, and U.S. industrial production. This information set allows us to track the domestic transmission channels to U.S. monetary policy shocks. Given our interest in international spillovers, we further include indicators for global economic activity, trade, and the financial cycle. We proxy the global real and trade cycles using the rest of the world (RoW, i.e., excluding the U.S.) industrial production and export indices constructed by the Federal Reserve Bank of Dallas (Grossman et al., 2014).[8] The global financial cycle index is derived from a dynamic factor model constructed from a comprehensive panel of risky asset prices traded worldwide (Miranda-Agrippino and Rey, 2020), summarizing global financial conditions.[9] Consistent with the literature, we stationarize the variables before estimating the TVP-VAR. We use the indicators for U.S. prices, U.S. industrial production, RoW industrial production, and RoW exports in log-differences to compute growth rates and keep the remaining variables in levels. Our monthly dataset covers the time span from 1980M1 to 2017M12, which we split into two parts. The first part ranges from 1980M1 to 1992M12 and is used as a training sample for prior calibration. The second part concerns the estimation sample ranging from 1993M1 to 2017M12. Since our sample contains the zero lower bound period, we replace the federal funds rate with the shadow rate of Wu and Xia (2016) during the period 2008M12-2015M12.[10] Unless otherwise noted, we estimate all model specifications with $p = 3$ lags due to the curse of dimensionality inherent in TVP-VAR estimation (this is a standard choice, see for instance Paul, 2020). The exact transformations and data sources can be found in Appendix A.

[8] We proxy global trade with RoW exports since it completely excludes U.S. produced goods. However, we find very similar results when considering RoW imports. Notice that RoW exports are aggregated from individual countries merchandise exports to the world. Therefore, the measure also includes intra-RoW exports.

[9] Aldasoro et al. (2023) show that the global financial cycle index as a price-based global factor is remarkably similar to a quantity-based global factor based on cross-border capital flows.

[10] Potentially, this series captures the effects of conventional and unconventional monetary policies. Since we achieve identification with high-frequency instruments of conventional monetary policy around FOMC meetings, identification should not be confounded with the effects of unconventional monetary policy. In addition, in Appendix D we show that our results are similar when excluding the zero lower bound period from the estimation sample.

To identify U.S. monetary policy shocks, we rely on a high-frequency monetary policy instrument (Gürkaynak et al., 2005), which is inserted as an exogenous variable in the VAR following Paul (2020). We use the instrument of Miranda-Agrippino and Ricco (2021) to rule out the presence of the so-called *information effect* (Melosi, 2017; Nakamura and Steinsson, 2018).[11] Notably, Miranda-Agrippino and Ricco (2023) demonstrate that correct identification of U.S. monetary policy shocks using the instrument of Miranda-Agrippino and Rey (2020) (as we do) can be achieved also with small-scale VARs that include only a few lags – i.e., in a setting like ours.[12]

2.1 Econometric framework

We use a time-varying parameter vector autoregression (TVP-VAR) to measure the effects of U.S. monetary policy. The model specification strongly resembles the one in Paul (2020).

Let $\{\boldsymbol{y}_t\}_{t=1}^{T}$ denote an n-dimensional time series process that evolves according to

$$\boldsymbol{y}_t = \boldsymbol{c}_t + \sum_{j=1}^{p} \boldsymbol{A}_{jt}\boldsymbol{y}_{t-j} + \boldsymbol{B}_t z_t + \boldsymbol{u}_t, \quad \boldsymbol{u}_t \sim \mathcal{N}_n(\boldsymbol{0}, \boldsymbol{\Sigma}), \tag{2.1}$$

where $\boldsymbol{c}_t$ is a $n \times 1$ vector of time-varying intercepts and $\boldsymbol{A}_{jt}$ $(j = 1, \ldots, p)$ are $n \times n$ time-varying coefficient matrices of the lagged endogenous variables. Reduced-form innovations are given by the $n \times 1$ vector $\boldsymbol{u}_t$, which follow a multivariate Gaussian distribution with zero mean and constant covariance matrix $\boldsymbol{\Sigma}$. Additionally, the model includes an exogenous variable z_t (the monetary policy instrument, that is used for identification – see next section) with its respective $n \times 1$ time-varying coefficients $\boldsymbol{B}_t$. When suppressing the time subscripts on the coefficients, this model nests a linear specification, which will be used in the empirical analysis as a preliminary exercise.

One remark is necessary. Contrary to the seminal contribution of Primiceri (2005), we do not allow for stochastic volatility in our baseline estimations. There are two reasons behind this choice. First, our estimation sample starts in the Great Moderation (the 1960-70s are excluded). Second, the relatively short sample we have (due to data constraints) makes it difficult to estimate heavily parameterized models, and stochastic volatility adds another significant estimation burden (see Paul (2020) for the same choice). Furthermore, as we show below in detail, the model still yields time-varying effects on impact through the coefficient vector $\boldsymbol{B}_t$, which are usually governed

[11] The instrument of Miranda-Agrippino and Ricco (2021) is constructed exploiting Greenbook forecasts, which are released to the public with a lag of five years. The instrument is available up until 2017M12 (Degasperi and Ricco, 2021), which constrains our sample to this specific time frame.

[12] Miranda-Agrippino and Ricco (2023) show that even just two lags (vs. three lags we employ) are sufficient for adequate shock identification, and argue that this "suggests that misspecification due to omitted variables or lags is minor." In addition, as shown in Bauer and Swanson (2023a), the construction of this instrument is robust to the “Fed response to news” channel of Bauer and Swanson (2023b).

by the covariance matrix.[13] However, we show in the robustness checks that, when we allow for stochastic volatility (at the cost of reducing the number of lags and having informative priors), we find similar results.

Let $\boldsymbol{\theta}_t = \text{vec}\left(\boldsymbol{c}_t, \boldsymbol{A}_{1t}, \ldots, \boldsymbol{A}_{pt}, \boldsymbol{B}_t\right)$ be a vector that stacks all coefficients on the right-hand side of Eq. 2.1. The dynamic evolution of the model's time-varying parameter is specified as follows

$$\boldsymbol{\theta}_t = \boldsymbol{\theta}_{t-1} + \boldsymbol{\nu}_t, \quad \boldsymbol{\nu}_t \sim \mathcal{N}(\mathbf{0}, \boldsymbol{Q}). \tag{2.2}$$

Hence, the elements of the vector $\boldsymbol{\theta}$ are modelled as driftless random walks. Furthermore, we assume that the innovations of the observation equation (Eq. 2.1) and the state equation (Eq. 2.2) are orthogonal. We pursue a Bayesian approach to estimation, which follows the procedures of Primiceri (2005) and Del Negro and Primiceri (2015). We use a linear VAR estimated over a pre-sample (1980M1-1992M12) to calibrate the prior distributions (Normal distribution for $\boldsymbol{\theta}$ and inverse-Wishart for $\boldsymbol{\Sigma}$ and $\boldsymbol{Q}$). We observe the high-frequency surprises starting from 1991M1. Since the policy instrument is included directly into the specification as an exogenous variable, the sample period of the instrument in principle constrains the sample length of the VAR. Following Paul (2020), the missing values in the surprise series are censored to zero prior to 1991M1 (i.e., during the pre-sample period from 1980M1 to 1991M1, we always observe the monetary instrument during the estimation sample).[14] The hyperparameter governing the prior belief on the amount of time variation in $\boldsymbol{\theta}$ is set as in Paul (2020). All results are based on 15,000 draws from the full posterior density simulated with a Gibbs sampler. We discard the first 5,000 draws as burn-ins. All the estimation details are described in Appendix B.

2.2 Identification of monetary policy shocks

We now outline our identification procedure. Let $\boldsymbol{\varepsilon}_t$ be a vector of structural disturbances, which are related to the reduced-form innovations $\boldsymbol{u}_t$ via a linear mapping $\boldsymbol{u}_t = \boldsymbol{S_t}\boldsymbol{\varepsilon}_t$, where $\boldsymbol{S_t}$ collects the contemporaneous impulse matrix at t. We are interested in the effects of U.S. monetary policy and thus in identifying one particular column of the matrix $\boldsymbol{S_t}$, which we denote by $\boldsymbol{s_t}$. Without loss of generality, we assume that the impulse vector $\boldsymbol{s_t}$ corresponds to the structural monetary policy shock ε_{1t} in our empirical specification (with the policy rate ordered first). To achieve identification, we assume that the monetary policy instrument z_t is i) correlated to the unobserved monetary policy shock and ii) orthogonal to all the other structural shocks. z_t is further assumed to be linked with

[13] If identification is achieved via a decomposition of the covariance matrix (e.g., using short-run restrictions or an external instrument), the covariance matrix has to be time-varying to yield time-varying effects on impact. In our case, we circumvent this via the exogenous variable approach of Paul (2020).

[14] See Noh (2019) for a formal justification of this procedure. In the robustness checks, we show that our results are similar when employing an uninformative prior on $\boldsymbol{B}_t$.

the shock via

$$z_t = \varphi \varepsilon_{1t} + \zeta_t, \qquad \zeta_t \sim \mathcal{N}(0, \sigma_\zeta^2), \tag{2.3}$$

where ζ_t is orthogonal to all other variables. This assumption implies that the relation between the instrument and the monetary policy shock is not time-varying.[15]

The contemporaneous relative impulse response of a variable i in $\boldsymbol{y}_t$ at time t to a shock generating a unit increase in the policy rate is then given by

$$r_{t,i1} = \frac{s_{t,i}}{s_{t,1}} = \frac{B_{t,i}}{B_{t,1}}. \tag{2.4}$$

Paul (2020) shows that this identification strategy is essentially equivalent to the external instrument approach (Stock and Watson, 2012; Mertens and Ravn, 2013).[16] The posterior quantities of $\boldsymbol{A}_{jt}$ ($j = 1, \ldots, p$; $t = 1, \ldots, T$) are then used to trace out subsequent impulse responses. In our baseline estimations, we normalize the impact effect of a contractionary monetary policy shock on the federal fund rate to be a one percentage point hike in 1993M1. Such a shock implies a particular variation in z_t that can then be exploited to calculate the impulse responses for the remaining periods in a way to compare same-sized shocks (using Eq. 2.3; see Paul, 2020). In addition, we report an alternative normalization that does not depend on Eq. 2.3, presenting the responses to a monetary policy shock normalized to induce a 1% decline in U.S. industrial production at each t.

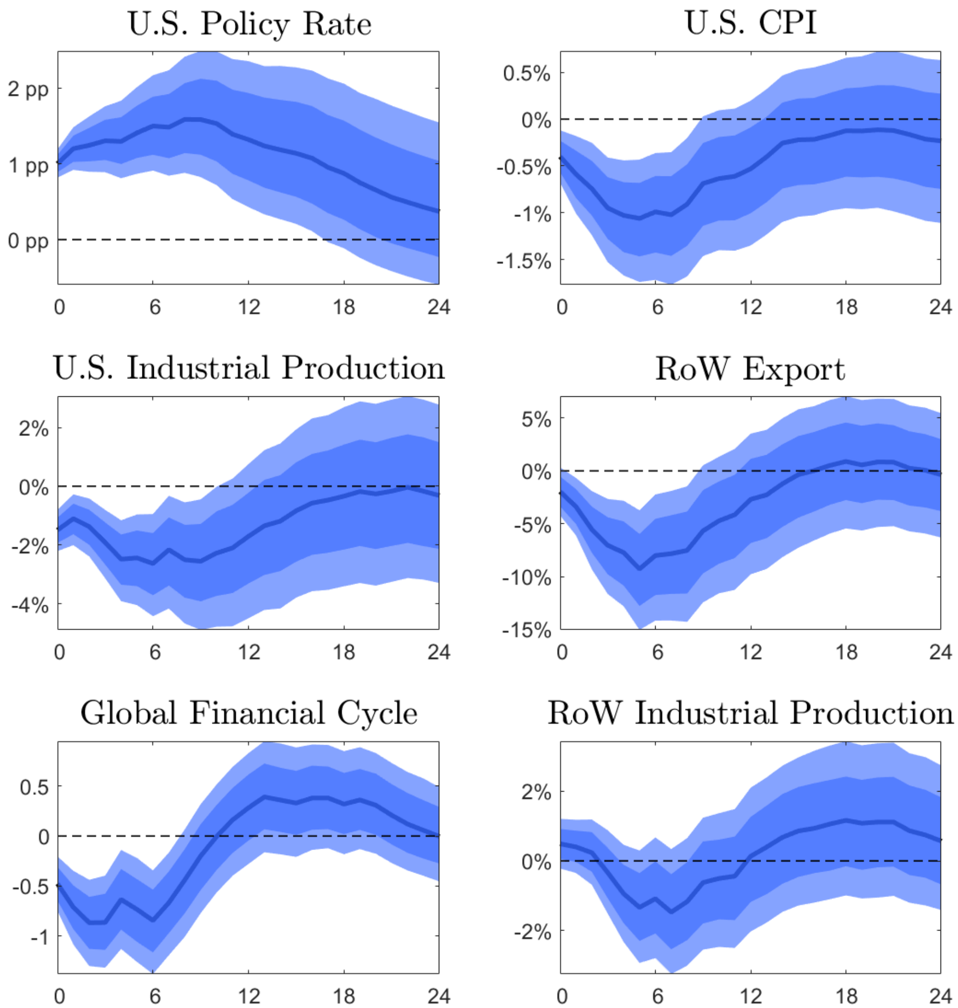
3. Empirical evidence

This section presents the empirical findings of the baseline specification and various extensions. In order to set the stage, we first show the dynamic effects of U.S. monetary policy shocks in a constant parameter VAR. Then, we examine the results obtained from the time-varying parameter model. By doing so, we address the main question of the paper: Have the global spillovers of U.S. monetary policy changed over time? Once this result is established, we investigate what causes this relationship to change over time. We investigate differences between AEs and EMEs and examine the potentially time-varying relevance of the *trade* and *financial* channel.

[15] This assumption has been recently tested by Amir-Ahmadi et al. (2023). They show that the relationship between the conventional monetary instruments and the unobserved shocks rises episodically. While we acknowledge this concern, we note our results point toward a *gradual* increase in the global spillovers of U.S. monetary policy, which cannot be driven by such an *episodic* relationship. Our assumption follows much of the previous literature on time-varying proxy VARs (Paul, 2020; Mumtaz and Petrova, 2023).

[16] Specifically, the contemporaneous relative impulse responses estimated by the two approaches are always the same. In addition, also the subsequent responses coincide if z_t is orthogonal to the VAR regressors. For this reason, in the empirical exercise we follow Paul (2020) and project the policy instrument on the lags of the observables and consider the residual of such projection as the exogenous variable. Intuitively, the external instrument approach (proxy VAR) boils down to i) estimating the reduced-form VAR and retrieving the residuals; ii) projecting them onto the instrument to identify the impact effects (up to a normalization). This is equivalent to what happens in Eq. 2.1.

Figure 2: Impulse responses functions of linear model.



Notes: Responses to a one percentage point (1 pp = 100 basis points) contractionary monetary policy shock. Median response and 68% and 90% confidence intervals are reported (wild bootstrap; 2,000 samples).

3.1 Evidence from the linear specification

We first estimate a linear VAR considering the variables described before.[17] We test the relevance of the policy instrument in the linear VAR, and we find a first-stage F-statistic of 22.8, which is safely above the standard threshold of 10 (see, for instance, Montiel Olea et al., 2021).

Figure 2 collects the dynamic responses to a one percentage point contractionary U.S. monetary policy shock. A monetary policy tightening is followed by strong domestic recessionary effects, consistent with standard macroeconomic theory: Domestic real activity deteriorates and prices

[17] Following the standard practice in the linear VAR literature, we include $p = 12$ lags and consider the variables in (log-)levels. We use frequentist procedures. Since a pre-sample period is not needed, we constrain the sample by the availability of the monetary policy instrument (1991M1-2017M12).

decline over the business-cycle horizon. These effects are statistically significant and similar to previous estimates from the VAR literature. Turning to the global effects, the policy shock generates substantial contractions in terms of global trade and the financial cycle, with these effects fading out only after almost one year (see e.g. Miranda-Agrippino and Rey (2020) and Degasperi et al. (2020) for similar results). The influence of U.S. monetary policy extends beyond national borders, playing a significant role not only in shaping domestic dynamics but also in influencing the broader global economic outcomes. The ramifications of reduced trade and heightened financial stress give rise to a worldwide economic downturn, quantified by a -1.6% decline in industrial production in the RoW. Our results align with the estimates by Breitenlechner et al. (2022), who find a global downturn of about -1.8% in response to a same-sized shock. Hence, even though the effects are only borderline significant in this linear framework, the magnitude of the global recession is economically substantial, particularly in comparison to the U.S. downturn.

All in all, our linear VAR seems to successfully identify U.S. monetary policy shocks, producing results that align closely with the existing literature. In order to analyze whether and how the global and local effects of US shocks changed over time, we now turn to the outcomes of the time-varying parameter model.

3.2 *Evidence from the time-varying specification*

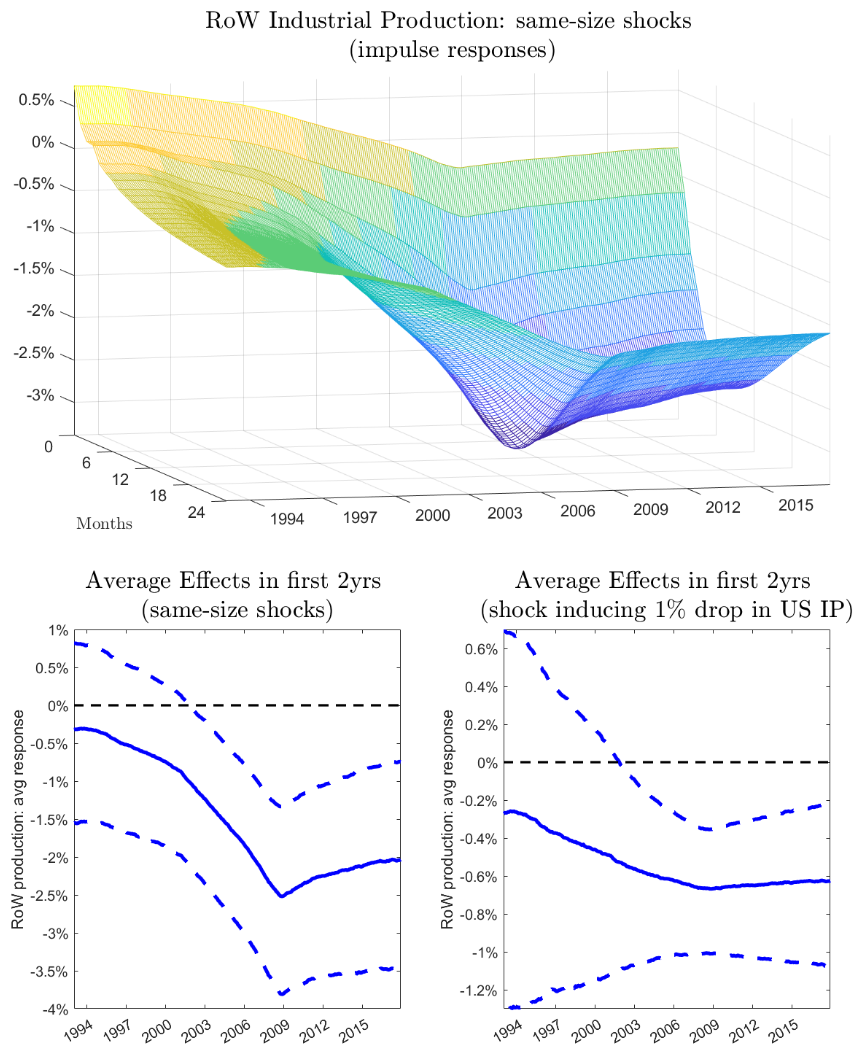
We report the time-varying impulse response functions from the TVP-VAR in Figure 3 (RoW industrial production) and Figure 4 (global trade and financial cycle). We report the domestic effects in Figure C1 in the Appendix. Since we stationarize the variables to estimate the TVP-VAR, we transform the variables in growth rates back to levels by taking the cumulative sum of the responses. The responses are thus scaled in percent for RoW exports and RoW industrial production. The responses of the global financial factor are to be interpreted in terms of standard deviations.

To facilitate interpretation, we present the results in two ways. First, we report the evolution of the median impulse-response functions over time. Second, we display the average response of the variables in the first two years following the shock at each t, along with their 68% credibility sets. We report the time-varying effects using two alternative normalizations. As a baseline, we follow Paul (2020) and determine the size of the shock required to induce a one-percentage-point increase in the U.S. policy rate on impact at 1993M1 (the first month of the estimation sample). Then, using Eq. 2.3, we trace out the effects of shocks of the same size in subsequent periods ("same-size shocks" normalization). This is our preferred normalization, as it assumes only a constant relationship between the monetary instrument and the unobserved shock, allowing the data to determine the remaining effects. Additionally, we present an alternative normalization, in which we report the responses of all variables conditional on a shock that, at each t, leads to an average contraction of U.S. production by 1% in the first two years following the shock. We consider this normalization a lower bound for our time variation estimates. Assuming constant domestic effects may underestimate time variation, as stronger global responses to U.S. monetary policy shocks could generate spillover-spillback loops, influencing domestic aggregates in a time-dependent manner (Obstfeld, 2020; Breitenlechner et al., 2022; Wu et al., 2024).

Figure 3 presents the main result of our paper.[18] Consistent with an amplified role of the international transmission channel of U.S. monetary policy shocks, we find an increase in the (recessionary) effects of U.S. shocks on RoW industrial production. While RoW production responds negatively throughout the period considered, the size of the effects strongly increases over time (lower left panel). According to our baseline normalization, a contractionary monetary policy shock, which increases the U.S. policy rate by 1 percent in 1993, generates a very mild (and not statistically significant) contraction at the beginning of the sample – a decline of about 0.4% in RoW production. Yet, a same-size shock becomes progressively more recessionary at the global level, peaking during the Great Recession, when it generates a 2.5% recession. Although the magnitude of the global effects has been growing since the beginning of the sample, the pace of this growth notably accelerates in the early 2000s. This downward trend stabilizes only with the Great Recession. Several explanations could account for this stabilization. It may be attributed to the slowdown in trade and financial integration resulting from the financial crisis (as shown in Figure 1), the effectiveness of international macro-prudential policies (which reduced banks' risk-taking propensity and their relevance in intermediation), or the presence of the zero lower bound period. Our empirical model appears to effectively capture this economic narrative. In this regard, the linear VAR seems to capture well the mean effect over time, masking though the time-specific heterogeneity. The time variation in global spillovers is also evident in our second normalization, which analyzes the evolving effects on RoW production in response to shocks that induce a constant U.S. recession (lower right panel). Early in the sample, a contractionary shock leading to a 1% U.S. recession results in a mild and statistically insignificant -0.26% global downturn, but this impact grows to a statistically significant -0.67% by the end of the period. To put it differently, while the global recession is about one-fourth of the U.S. downturn in the 1990s, by the end of the sample it exceeds two-thirds of it. Hence, our first stylized fact is that, no matter the normalization we use to compare shocks over time, we find evidence of substantial time variation in the international ramifications of policy shocks: U.S. policy hikes generate stronger recessionary effects over time, with a flattening out of effects after the Great Recession.

[18] Figure C1 in the Appendix confirms the results of the linear VAR, showing that monetary tightening leads to negative demand effects in the U.S. economy. Industrial production contracts alongside falling prices, with magnitudes consistent with the linear model. Two key observations emerge. First, maintaining same-size shocks requires a weaker impact on the policy rate over time, reflecting the long-run decline in U.S. interest rates, which bottoms out during the zero lower bound period (2009-2015). Second, average contractions in industrial production increase from -1.2% in 1993 to -3.7% in 2008, stabilizing only with the Great Recession, in line with the estimates of Paul (2020). A similar trend appears for prices, with contractions increasing from -0.3% to -1.0% by the Great Recession.

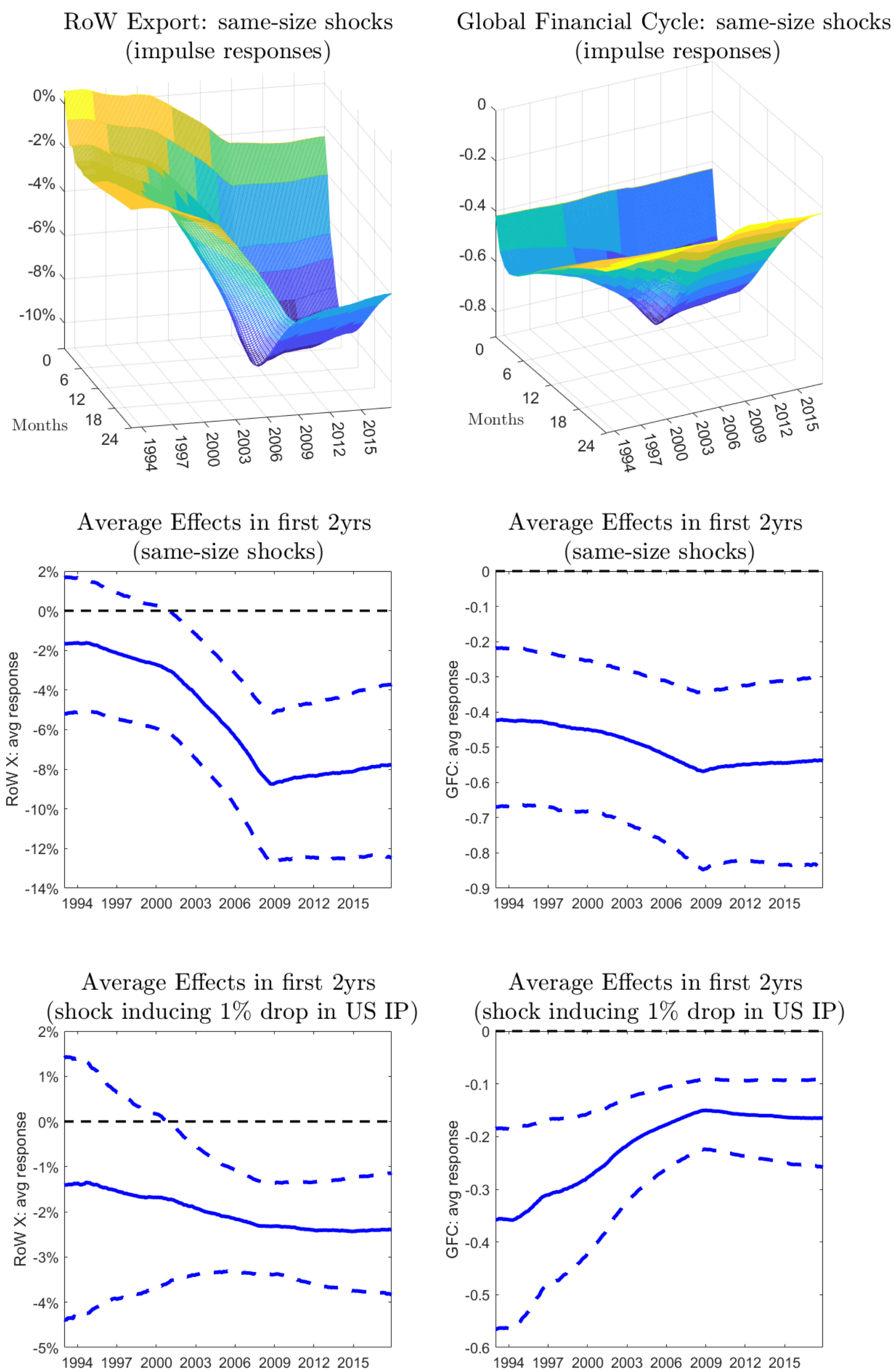
Figure 3: Time-varying impulse responses functions of RoW industrial production.



Notes: Responses of RoW industrial production to a contractionary U.S. monetary policy shock. Top panel: evolution of the median impulse-response functions over time ("same-size shock" normalization). Bottom panels: average effects in the first two years after the shock, using the "same-size shocks" normalization (left) and the normalization that induces a 1% average contraction in U.S. industrial production over the same period (right). Dashed lines: 68% posterior credible sets. Response of RoW industrial production is given in percent.

So far, we have documented economically meaningful time variation in the global spillovers of U.S. monetary policy on real activity. Since our analysis is motivated by the hypothesis that these effects could be driven by the increasing role of trade and financial integration, we now examine their impact on two related global cycles, proxied by RoW exports and the global financial cycle (Figure 4). Consistent with the linear VAR, we find sharp contractions in both RoW exports and the global financial cycle. However, once again, time variation appears to be a crucial factor. Under our baseline normalization, the response of RoW exports (as a proxy for trade) grows significantly

Figure 4: Time-varying impulse responses functions of global variables.



Notes: Responses of RoW export and Global Financial Cycle to a contractionary U.S. monetary policy shock. Top panels: Evolution of the median impulse-response functions over time ("same-size shock" normalization). Bottom panels: average effects in the first two years after the shock, using the "same-size shocks" normalization and the normalization that induces a 1% average contraction in U.S. industrial production over the same period. Dashed lines: 68% posterior credible sets. Response of RoW export is given in percent. Response of the global financial cycle is given in standard deviation.

in magnitude over time, increasing from an average of -2% to -8%.[19] The time variation in the impulse response of the global financial cycle is more moderate, increasing from -0.43 to -0.55. However, while the growing trade spillovers remain important even under the normalization that induces a constant U.S. recession, the contraction in the global financial cycle, if anything, exhibits a slight decline over time. This brings us to our second stylized fact: The increasingly recessionary global effects of U.S. monetary policy are accompanied by growing contractions in global trade, while financial spillovers remain relatively stable.

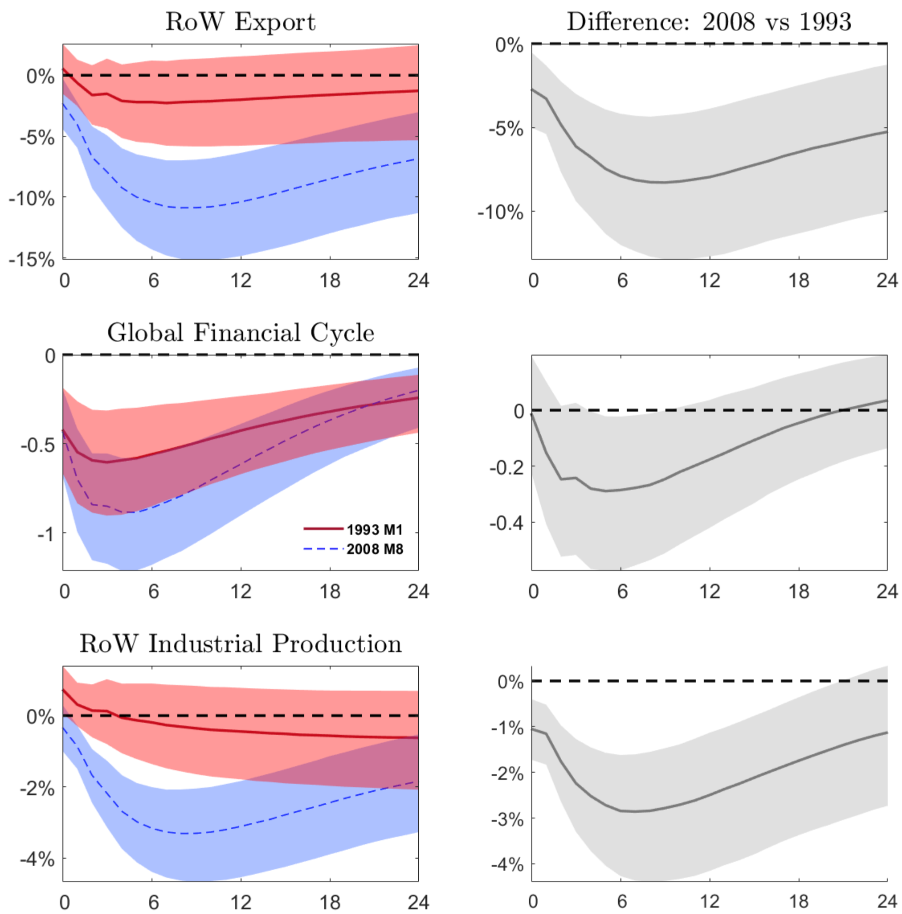
Finally, we investigate whether time variation is statistically significant by focusing on particular episodes in the sample, focusing our baseline normalization. We consider the periods 1993M1 and 2008M8. We report in Figure 5 the impulse response functions for the two time periods (left column) and the posterior distribution of the differences (right column). The primary distinction between the two examined periods lies in the effects on RoW industrial production. On the one hand, we find no evidence in support of a global downturn in real activity following monetary policy contractions in 1993. On the other hand, such negative effects are clearly present in 2008. A similar pattern arises for RoW export. In addition, as Figure 5 indicates, differences for RoW exports and RoW industrial production between these time periods are statistically significantly different from zero (right column). Time variation is a relevant pattern for global real spillovers. In contrast, the evidence for the global financial cycle is much weaker: While U.S. disturbances generate significant financial frictions in both periods, the difference in the impulse-response functions is only borderline statistically different from zero at a few horizons.

Sensitivity checks of the baseline results. In Figure D1 (Appendix D), we discuss several robustness and sensitivity checks of the baseline results. We show that our results are robust to: i) considering alternative instruments for the identification of U.S. monetary policy shocks (Gertler and Karadi, 2015; Jarociński and Karadi, 2020; Bauer and Swanson, 2023a; Hoesch et al., 2023); ii) restricting the sample such that it excludes the zero lower bound or controlling for forward guidance and quantitative easing surprises (Swanson, 2021); iii) considering more lags in the VAR; iv) different prior selections for time variation and the instrument; v) adding stochastic volatility to the model. We refer to Appendix D for a detailed explanation of these exercises.

Alternative financial indicators and wider propagation channels. In the Appendix E, we show that the negative but relatively stable effects on global financial conditions remain robust when using alternative global indicators instead of the global financial cycle of Miranda-Agrippino and Rey (2020). Specifically, we obtain similar results when considering the RoW MSCI index (a stock market index excluding U.S. stocks), global private liquidity, global credit, and global cross-border flows. Additionally, we report the effects on a range of relevant macroeconomic variables, finding

[19] In appendix Appendix E, we obtain very similar results when considering RoW imports. This aligns with a symmetric contraction of both RoW exports and imports in response to U.S. policy shocks (Gopinath et al., 2020; Degasperi et al., 2020).

Figure 5: Differences in impulse responses: 1993M1 vs. 2008M8



Notes: Left column: median impulse-response functions and 68% posterior credible sets for the variables considered at 1993M1 vs. 2008M8. Vertical axis: percentage change; horizontal axis: impulse response horizon in months. Right column: difference in impulse responses in such periods (median and 68% posterior credibility intervals are reported).

results consistent with the existing literature and discussing their time variation. Notably, we find that contractionary U.S. monetary policy shocks generate a rather constant appreciation of the U.S. dollar effective exchange rate, together with a strong and highly time-varying fall of U.S. imports, that tracks well the dynamics observed for RoW trade in our benchmark model.

3.3 Heterogeneity analysis

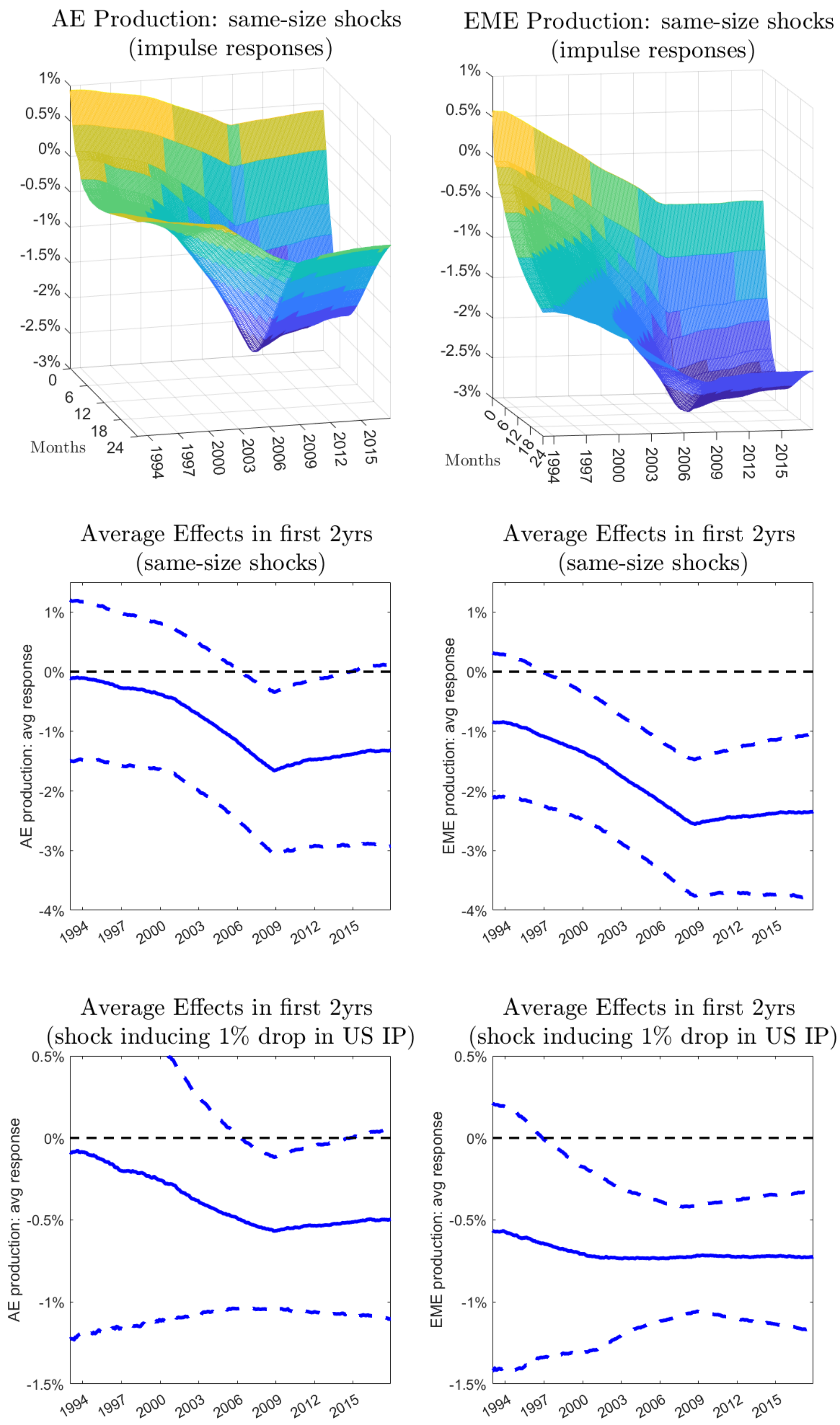
We have documented a significant time variation in the international spillovers of U.S. monetary policy. In this and the following section, we conduct additional exercises to explain the findings and link them to different channels. As highlighted by Kalemli-Özcan (2019), De Leo et al. (2022), and Ca'Zorzi et al. (2023), monetary policy spillovers are quite asymmetrical between advanced and emerging economies (AEs and EMEs). Given that EMEs have increased their share of world economic activity from approximately 20% to 40% over the past three decades, part of the observed time variation could be explained by *mechanical* composition effects. Hence, we look into differences in spillovers to AEs and EMEs. To start disentangling the channels at play, we are interested in finding cross-sectional variation. To do so, we estimate country-specific spillovers to industrial production. In the next section, we will finally combine these estimates with country-specific trade and financial data to evaluate the evolution of the transmission channels.

To look into the difference between AEs and EMEs, we adapt the baseline specification by replacing the RoW industrial production indicator and RoW export indicator with the respective indicator for AEs or EMEs.[20] The results, shown in Figure 6, reveal more pronounced recessionary effects in emerging economies – consistent with existing studies. The average effect of same-size shocks on emerging (advanced) countries' industrial production is -0.9% (-0.1%) in 1993 and -2.6% (-1.7%) in 2008 (we find evidence of time variation in our alternative normalization, see the middle panels of Figure 6). In those years, the response of RoW industrial production, which combines both emerging and advanced economies, is -0.4% and -2.5%, respectively. While at the beginning of the sample period, the response of RoW industrial production is strongly tilted towards one of advanced economies (emerging economies have little relevance in the overall index), the response in 2008 aligns more with the effect in EMEs. Since the composition of RoW industrial production between 1993 and 2008 has strongly changed (in favor of EMEs), composition effects can account for part of the rising spillovers that we observe in the data. However, the time variation in the responses of both emerging and advanced economies' industrial production – which is also present in our more conservative normalization, see the bottom panels of Figure 6 – signals that mechanical composition effects cannot fully explain the rising spillovers, which must depend on other factors.

In order to make progress on this issue, we are interested in retrieving more cross-sectional heterogeneity with respect to international spillovers. Therefore, we break down the response of RoW industrial production into its country-specific components. We re-estimate our benchmark VAR, replacing the aggregate RoW industrial production with country-level indices. This allows us to examine the response at a more granular level. We consider a total of 26 countries in our analysis. The selection of the countries is dictated by the availability of the industrial production series from 1980 onwards. Our sample includes: Austria, Australia, Belgium, Brazil, Canada, Chile, Colombia,

[20] Data is again taken from the Database of Global Economic Indicators of the Federal Reserve Bank of Dallas. The industrial production series for EMEs is available from 1987M1. To estimate the model starting from 1980M1, we assume that, from 1980M1 to 1986M12, the growth rate in industrial production of EMEs is equal to the one in the comprehensive RoW series (we always observe the actual EMEs series in the estimation sample). See the exact transformations and list in Appendix A.

Figure 6: Comparison of advanced and emerging market economies.



Notes: Responses to a contractionary U.S. monetary policy shock that induces a one percentage point (pp) increase in the federal fund rate in 1993M1. Upper panel reports the evolution over time of the median impulse-response functions. The lower panel reports the peak effects over time for each variable with 68% posterior credible sets.

Finland, France, Germany, Greece, India, Italy, Norway, South Korea, Japan, Malaysia, Mexico, Netherlands, Peru, Portugal, South Africa, Spain, Sweden, Switzerland, United Kingdom.[21]

To save space, we report each country's median time-varying impulse response functions in Appendix F (Figure F1-Figure F5). We also report the evolution of country-specific mean effects over time in Figure F6. The findings in Figure F6 align with our benchmark estimations and reveal a consistent pattern of increasing spillover effects. A U.S. policy tightening generates recessionary effects in most countries considered, especially after the early 2000s. These effects tend to intensify over time until the Great Recession, after which they stabilize. The magnitudes of the economic downturns are in the ballpark of our estimates for the aggregate RoW production. Additionally, two sources of cross-sectional heterogeneity among countries arise. Firstly, the *average* relevance of recessionary effects varies across countries: Certain countries, such as Canada and Mexico, historically exhibit a greater susceptibility to U.S. shocks, whereas others like the UK are less affected. This dimension can be summarized by taking the average over time of the effect in country i. Secondly, the *increase* in recessionary effects can be more or less substantial, and this dimension can be summarized by comparing the recessionary effects in country i at the end of the sample with those at the beginning. Time variation is pervasive in Japan, Germany, and Spain, while it is more limited in the Netherlands and UK.

3.4 Evaluation of the channels

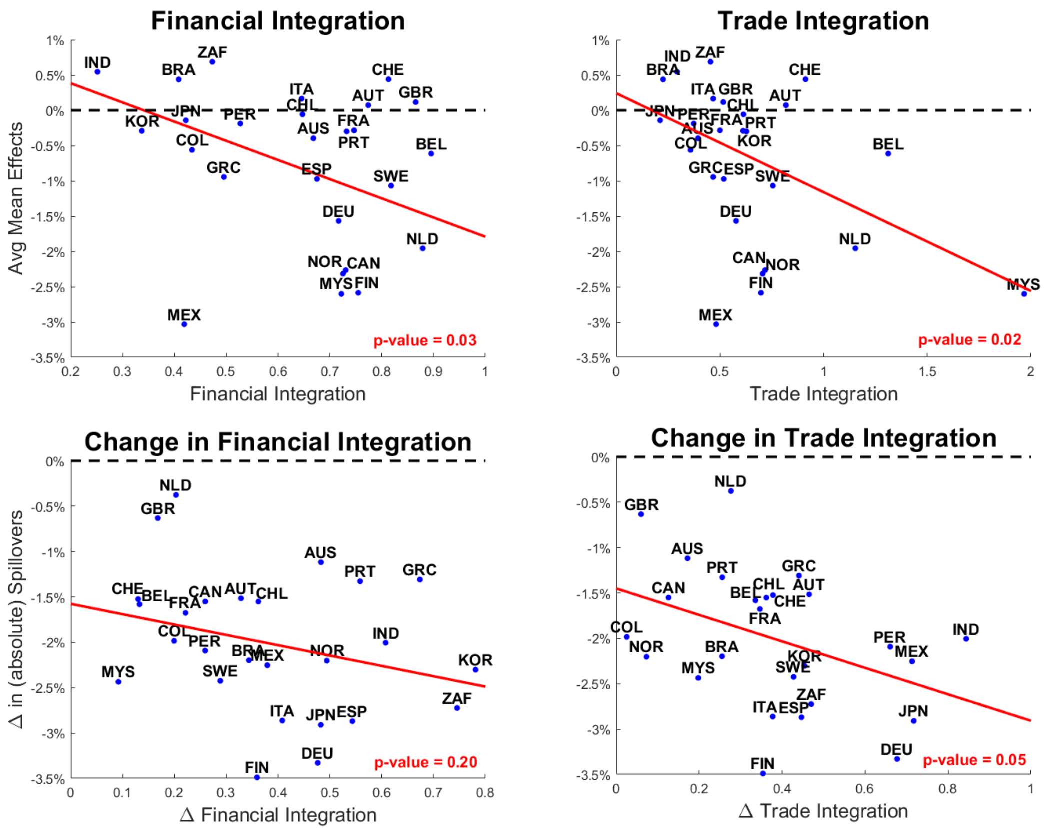
Our empirical analysis again points towards significant time patterns in global spillovers. But what are the main drivers of these dynamics? We shed light on this aspect by establishing a connection between country-specific effects and country-specific information on trade and financial integration. Our baseline measure of financial integration is the (de facto) financial globalization index from the KOF Swiss Economic Institute (Dreher, 2006). Following Lane and Milesi-Ferretti (2007), this index incorporates the assets and liabilities of foreign direct investments, the assets and liabilities of international equity portfolio investments, the inward and outward stocks of international portfolio debt securities and bank loans and deposits, as well as international reserves excluding gold.[22] Trade integration is measured as the ratio of total trade to GDP.[23] Financial and trade integration are then computed as the historical average of yearly measures for each country. Given that our aim

[21] We thank the authors of Grossman et al. (2014) for kindly sharing their data with us. Data sources are described here: https://www.dallasfed.org/research/international/dgei. Additional series are retrieved from OECD and FRED.

[22] Data is available here: https://kof.ethz.ch/en/forecasts-and-indicators/indicators/kof-globalisation-index.html.

[23] Trade integration data is sourced from the World Bank (World Development Indicators). Given that country-specific shares of dollar-invoiced trade are rarely available for the 1990s and early 2000s (see Boz et al., 2022), we use the overall trade-to-GDP ratio (irrespective of currency) as a proxy for overall trade integration. This measure also has the great advantage of capturing potential indirect effects through network structures (see Bräuning and Sheremirov, 2023).

Figure 7: Correlations of spillovers and financial and trade integration.



Notes: Upper panels: Relationship between the average mean response of individual countries during the period 1993-2008 and their average financial and trade integration during the same period. Lower panels: Relationship between the absolute change in country-specific mean response (the difference between the mean response in 2008 and 1993) and the corresponding change in financial and trade integration (variation between 2007 to 1993; here we consider 2007 rather than 2008 to exclude the massive movements in trade and financial integration occurred in 2008). Red lines: univariate regressions interpolating the points. Given our small sample, we reduce the impact of outliers by employing an iteratively reweighted least-squares algorithm.

is to explain time variation in the spillovers, we also account for the changes in trade and financial integration over time. Specifically, we calculate the (log) difference between integration right before the Great Recession and integration in 1993 (first year in our sample), which encompasses the time period where time variation manifests. Hence, we also consider average and time variation in the effects over the same period.

In Figure 7, we examine the correlation between the two dimensions of spillover heterogeneity and country-specific trade and financial integration measures (both in terms of *average* and *time variation*).[24] In the upper panel, we report the correlation of average effects with average integration, while in the lower panels we show the correlation of variation in effects with variation in integration.

[24] Since the UK and Switzerland exhibit outlier behavior in their average financial integration measures, we include dummy variables in the corresponding regression to prevent them from distorting the inference (upper left panel of Figure 7).

Looking at the upper panels, we first observe that, on average, higher levels of financial and trade integration are associated with higher international spillover effects – the slope coefficients of both regressions are statistically negative at a 5% level (p-values are 0.03 and 0.02 respectively). Countries that exhibit greater integration in these dimensions tend to experience more significant economic downturns following contractionary U.S. policy shocks (consistent with the existing literature). Both channels are at play *on average*. This evidence suggests that trade plays a major role as a transmission channel for the global spillovers of U.S. monetary policy shocks. In fact, the role of trade appears to be at least comparable to, if not larger than, that of financial linkages. These results align with the recent findings of Bräuning and Sheremirov (2023) and Camara et al. (2024), who show that trade is the primary mechanism through which a U.S. monetary contraction impacts global real activity, using both time-series methods and estimated open-economy DSGE models. Second, when it comes to explaining the *time variation* in the effects, our estimations suggest that the trade dimension plays a more significant role (lower panels of Figure 7). Countries that undergo a substantial increase in trade integration tend to experience a worsening of recessionary effects generated by U.S. monetary policy over time (this is reflected in negative values in the absolute change in spillovers). The slope coefficient for a regression interpolating the observations is negative, and it proves statistically significant at conventional confidence levels when conducting hypothesis tests (p-value = 0.05). Conversely, there is no strong evidence of a statistically negative relationship between changes in financial integration and changes in recessionary effects (p-value = 0.20). Patterns remain similar when estimating bivariate regressions that jointly model change in trade and financial integration. Hence, while we acknowledge the inherent reduced-form nature of this analysis, we believe that this exercise provides intriguing correlational evidence indicating that the increase in trade integration is associated with the exacerbation in economic spillovers of U.S. monetary policy.

Robustness checks. In Figure F7 (Appendix F), we present evidence supporting the robustness of our results when using alternative measures of financial integration. Beyond the KOF index, which provides a comprehensive assessment, we consider additional metrics, including: (i) the ratio of a country's dollar-denominated external assets and liabilities to GDP (from Bénétrix et al., 2019); (ii) external assets alone; (iii) external liabilities alone; and (iv) the Chinn-Ito Index of capital account openness (Chinn and Ito, 2006). If anything, relative to our baseline analysis in Figure 7, these additional checks indicate an even weaker connection between the rise in observed financial integration and the corresponding increase in spillovers. In addition, Table F1 documents that our main cross-sectional results are confirmed when looking at peak effects or various conventional horizons (rather than average effects in the first two years after the shock).

3.5 Counterfactual analysis

Taken at face value, the evidence just presented aligns with the estimates from the baseline specification. Examining Figure 4, we observe a stark contrast in the degree of time variation between the effects on global trade—which exhibit significant fluctuations over time—and financial conditions, which remain relatively stable. Based on this evidence, we argue that while changes in the international financial landscape may play a role, they cannot fully explain the observed time variation in real spillovers. Instead, this variation appears to be more strongly correlated with shifts in trade integration. Although financial transmission is likely one of the key channels for international spillovers *on average* (Rey, 2016; Miranda-Agrippino and Rey, 2020), it does not seem to be the primary driver of the observed time dynamics. This reasoning relies on the assumption of a constant *financial multiplier*, which captures how global financial frictions amplify the real effects of monetary shocks. To further support this argument and highlight the dominant role of trade, we provide additional evidence through counterfactual simulations that isolate the trade and financial transmission channels within our VAR framework.

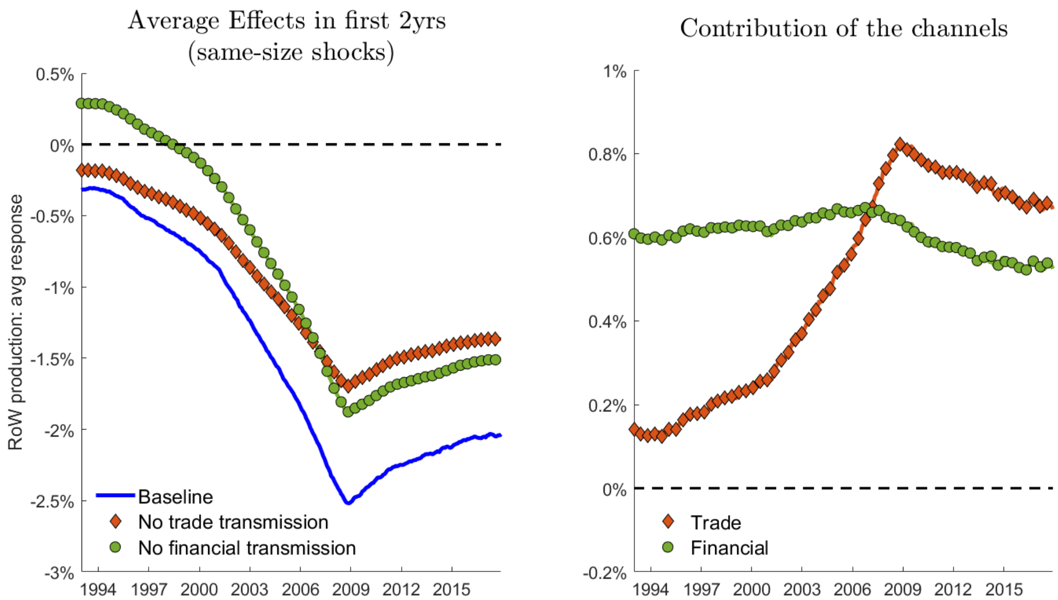
To achieve this, we identify both a global trade shock and a global financial shock. Following the seminal work of Gilchrist and Zakrajšek (2012), we identify the financial shock by estimating a recursive TVP-VAR, in which the financial indicator (in this case, the global financial cycle) is ordered after real activity and price indicators – considered slow-moving variables that respond only with a lag – but before the U.S. policy rate, which is fast-moving and allowed to react contemporaneously to the shock.[25] For symmetry, and to identify a global *trade-specific* shock, we apply the same methodology to trade, using RoW exports as the indicator.[26] Since we assume that neither shock affects business-cycle variables on impact, our identification is intentionally conservative and symmetric across shocks. Therefore, our estimates should be viewed as lower-bound estimates.

With the identified shocks at hand, we conduct counterfactual simulations to isolate the trade and financial transmission channels of U.S. monetary policy, shutting them down one at a time. In the left panel of Figure 8, the blue line represents the baseline average response of RoW industrial production (as estimated in Figure 3). The orange-diamond (green-circled) line shows the counterfactual responses in hypothetical scenarios where U.S. monetary disturbances no longer propagate through a deterioration of RoW trade (global financial conditions). To construct these

[25] Abbate et al. (2016) use a similar approach and find that global financial shocks have a substantial global impact, with relatively stable negative effects on economic activity over time.

[26] All in all, we estimate two recursive VAR models – using our baseline specification, dataset, and priors – where financial and trade indicators are placed (one at a time) at the bottom of the recursive ordering, followed only by the U.S. policy rate. Hence, our shocks may be interpreted as global trade and financial disturbances that are contemporaneously orthogonal to business cycle variables. The impulse responses to these shocks are reported in Figure G2 and Figure G3.

Figure 8: Counterfactual analysis.



Left panel: Peak responses of world industrial production to U.S. monetary shocks (blue line: baseline estimations) versus counterfactual scenarios obtained by zeroing out the response of RoW exports (orange diamond lines: "no trade transmission") and the global financial cycle's response to U.S. monetary shocks via global financial shocks (green circled lines: "no financial transmission"). Right panel: Absolute contribution of the two channels over time.

counterfactuals, we generate a sequence of global trade (financial) shocks that fully offset the impact of U.S. policy shocks on RoW export (global financial cycle) – using the method put forward in Sims and Zha (2006). This allows us to document how RoW production would have evolved in the absence of trade and financial transmission.[27]

Examining Figure 8, we find that both the trade and financial channels are quantitatively relevant. For instance, in 2008, RoW industrial production contracts by -2.5% in response to the monetary policy tightening we consider. Our estimates suggest that in the absence of the trade channel, the contraction would have been more limited at -1.7%, and without financial transmission, it would have been -1.9%. However, the importance of these channels evolves differently over time (right panel of Figure 8). The financial channel plays a consistently significant role throughout the sample

[27] The approach of Sims and Zha (2006) follows three main steps: i) At each t, estimate the effect of a trade (financial) shock using the method described earlier. ii) Identify the sequence of trade (financial) shocks that, when added to the "baseline system," would neutralize the response of RoW exports (global financial cycle). Since RoW exports (global financial cycle) contract in the baseline system, these sequences correspond to adverse trade (financial) shocks. iii) Compute the responses of system variables when subjected to these additional shocks. This procedure generates a simulated scenario in which U.S. policy shocks no longer affect the transmission variable of interest. McKay and Wolf (2023) propose a refinement that applies financial shocks only at horizon $h = 0$ to better approximate the counterfactual. However, implementing this refinement in our case would require identifying multiple global financial and trade shocks, which proves challenging.

period, gradually increasing from 0.6 to 0.7 percentage points until the Great Recession. In contrast, the trade channel is initially limited but rises sharply, peaking in 2008 as its contribution grows from 0.15 to 0.85 percentage points. Once again, time variation appears to be closely tied to trade dynamics. Finally, both channels' contributions decline in the aftermath of the Great Recession, aligning with the diminishing spillover effects observed in the data.

Robustness checks. Our identification is intentionally conservative, assuming that trade and financial shocks have no immediate impact on business-cycle variables. In Figure G1, we demonstrate that the trade channel plays an even larger role in explaining the time variation in spillovers when we use an alternative specification where trade shocks are assumed to have no immediate effect only on RoW industrial production (this restriction is necessary to distinguish trade shocks from conventional non-trade global disturbances; all other variables can be contemporaneously affected by trade disturbances). In this framework, the trade channel accounts for nearly half of the overall transmission of U.S. monetary policy shocks in 2008. Instead, when we perform the same exercise for the global financial shock, its overall contribution remains unchanged.

3.6 Taking stock

We hence conjecture that most of the variation is attributable to macroeconomic and trade-related factors. A battery of different results supports this. First, we observe the stylized fact that sharper contractions in real economic activity are closely linked to more pronounced declines in trade, whereas the impact on global financial conditions has remained relatively stable. Second, we show that emerging economies' rapid growth, combined with their higher sensitivity to adverse U.S. monetary policy shocks, mechanically increases overall spillovers via composition effects. Third, cross-sectional correlation analysis indicates rising trade integration as a likely driver of rising spillovers. Fourth, counterfactual simulations that shut down the role of trade and financial transmission of U.S. policy disturbances suggest that trade linkages play a major role, while financial linkages have a more limited impact as a potential driver of the observed time-varying dynamics in spillovers. As noted in the Introduction, several factors may explain why, despite the overall importance of the financial channel, rising financial integration (as conventionally measured) has not significantly amplified real effects over time. First, many countries shifted their net foreign currency positions from negative to positive as integration increased, mitigating adverse balance sheet effects in response to a U.S. monetary policy tightening that appreciates the dollar (Lane and Shambaugh, 2010). Second, falling dollar debt funding, progressively replaced by domestic-currency debt and equity instruments (portfolio equity and foreign direct investments), likewise promoted international risk sharing and acted as a stabilizing factor (Bénétrix et al., 2015; Obstfeld, 2021; Kalemli-Özcan and Unsal, 2023). Third, accumulated foreign reserves have enabled domestic authorities to act as lenders-of-last-resort for short-term liabilities, cushioning the impact of U.S. shocks (Bussière et al., 2015; Benigno et al., 2022). Fourth, the measured rise in financial integration may overstate the genuine increase in underlying cross-border financial linkages (Lane and Milesi-Ferretti, 2018).

Our results through the lens of a stylized model. In Appendix H, we show that a stylized open economy model with trade and financial frictions is consistent with our main documented facts. Specifically, we show that an increase in trade integration within a two-country framework leads to larger output spillovers in the RoW, more pronounced trade contractions in the RoW, and relatively stable financial transmission. The key takeaway from the model is that, as long as some degree of financial integration and frictions are present, rising trade integration generates qualitatively similar patterns of evolving effects over time as observed in the data.

4. Concluding remarks

We study whether and how the international effects of U.S. monetary policy shocks have changed over the last decades, providing evidence in support of growing global spillovers. To do so, we estimate a time-varying parameter VAR which features U.S. and global indicators. Identification is achieved using state-of-the-art methods which exploit high-frequency external instrument techniques. This enables us to account for time-varying responses of domestic and global aggregates to U.S. policy shocks. The need for a time-varying model is motivated by the substantial changes brought about by globalization over recent decades. The increased interconnectedness of global real and financial markets and the dominant role of the U.S. dollar make a strong case that international spillovers of U.S. monetary policy shocks have amplified over time. Our findings provide strong support for this hypothesis.

Our results reveal that the impact of a U.S. tightening on global industrial production has substantially increased over the last decades. The magnitude of the spillovers stopped growing only after the Great Recession. After this turning point, effects stabilized coherently with the observed slowdown in trade and financial integration after the crisis. Notably, while the pattern in the effects on global trade closely mimics that of global economic activity, we find that time variation in the response of global financial conditions is significantly smaller.

When evaluating the transmission mechanisms, we find that both the trade and financial channels are active on average. However, when we dig deeper into the channels explaining time variation, our estimations reveal that the increasing spillovers can primarily be attributed to the surge in trade integration. This conclusion is based upon a granular analysis of the country-specific time-varying effects and on counterfactual simulations aimed at shutting down the role of global financial and trade transmission of U.S. policy disturbances. Our findings suggest that global trade linkages and networks may be even more important than previously thought for the international transmission of U.S. monetary policy decisions.

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Online Appendix

Has Globalization Changed the International Transmission of U.S. Monetary Policy?

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A. Data

In Table A1, we define the exact transformations of the variables used in the benchmark estimation of the TVP-VAR.

Table A1: Variable definitions.

Variable	Transformation in TVP-VAR	Details
ffr_t	$\mathtt{FFR}_t$	$\mathtt{FFR}_t$ is the federal funds rate (Fred), replaced with the shadow rate of Wu and Xia (2016) during the ZLB period (2008M12-2015M12).
cpi_t^{US}	$100 \times \left[\ln\left(\mathtt{CPI}_t^{US}\right) - \ln\left(\mathtt{CPI}_{t-1}^{US}\right)\right]$	$\mathtt{CPI}_t^{US}$ is U.S. consumer price index (Fred).
ip_t^{US}	$100 \times \left[\ln\left(\mathtt{IP}_t^{US}\right) - \ln\left(\mathtt{IP}_{t-1}^{US}\right)\right]$	$\mathtt{IP}_t^{US}$ is U. S. industrial production (Fred).
gfc_t	$\mathtt{GFC}_t$	$\mathtt{GFC}_t$ is the global financial cycle indicator by Miranda-Agrippino and Rey (2020).
ip_t^{RoW}	$100 \times \left[\ln\left(\mathtt{IP}_t^{RoW}\right) - \ln\left(\mathtt{IP}_{t-1}^{RoW}\right)\right]$	$\mathtt{IP}_t^{RoW}$ is the RoW industrial production (excluding the U.S.) by Grossman et al. (2014).
ex_t^{RoW}	$100 \times \left[\ln\left(\mathtt{EX}_t^{RoW}\right) - \ln\left(\mathtt{EX}_{t-1}^{RoW}\right)\right]$	$\mathtt{EX}_t^{RoW}$ is RoW exports (excluding the U.S.) by Grossman et al. (2014).
ip_t^{AE}	$100 \times \left[\ln\left(\mathtt{IP}_t^{AE}\right) - \ln\left(\mathtt{IP}_{t-1}^{AE}\right)\right]$	$\mathtt{IP}_t^{AE}$ is the advanced economies' industrial production (excluding the U.S.) by Grossman et al. (2014).
ex_t^{AE}	$100 \times \left[\ln\left(\mathtt{EX}_t^{AE}\right) - \ln\left(\mathtt{EX}_{t-1}^{AE}\right)\right]$	$\mathtt{EX}_t^{AE}$ is AE exports (excluding the U.S.) by Grossman et al. (2014).
ip_t^{EME}	$100 \times \left[\ln\left(\mathtt{IP}_t^{EME}\right) - \ln\left(\mathtt{IP}_{t-1}^{EME}\right)\right]$	$\mathtt{IP}_t^{EME}$ is the emerging market economies' industrial production by Grossman et al. (2014).
ex_t^{EME}	$100 \times \left[\ln\left(\mathtt{EX}_t^{EME}\right) - \ln\left(\mathtt{EX}_{t-1}^{EME}\right)\right]$	$\mathtt{EX}_t^{EME}$ is the emerging market economies exports by Grossman et al. (2014).

Notes: RoW is short for rest-of-world. All variables are available over the pre-sample and estimation sample period, ranging from 1980M5 to 2017M12. Exceptions is: $\mathtt{EME\ IP}_t$ 1987M1 to 2017M12.

B. TVP-VAR: prior settings and Bayesian estimation

We describe the details on the prior density choice and hyperparameter calibration. We have to set prior densities for the initial values of $\boldsymbol{\theta}$, which we denote with $\boldsymbol{\theta}_0$. For all these initial values, we specify Gaussian distributions. We also have to specify a prior density for the covariance matrices $\boldsymbol{Q}$, where we use an inverse-Wishart prior density. Last, we also need a prior distribution for the covariance matrix of the VAR, $\boldsymbol{\Sigma}$, which also follows an inverse-Wishart distribution.

To calibrate the prior distributions, we use the first 13 years as a training sample. This results in a training sample ranging from 1980M1 to 1992M12 of length $\tau = 156$. We obtain estimates using Ordinary Least Squares (OLS) from this training sample and assume the following prior densities for the coefficients in the TVP-VAR

$$\begin{aligned} \boldsymbol{\theta}_0 &\sim \mathcal{N}\left(\hat{\boldsymbol{\theta}}_{OLS}, 4 * V(\hat{\boldsymbol{\theta}}_{OLS})\right), \\ \boldsymbol{Q} &\sim iW\left(\kappa_Q^2 * \tau * V(\hat{\boldsymbol{\theta}}_{OLS}, \tau)\right), \end{aligned} \tag{B.1}$$

where the subscript OLS refers to the OLS estimator of the respective coefficient. We denote by $iW(S, d)$ an inverse-Wishart distribution with degrees of freedom d and scale matrix S. For the initial values, we use the OLS point estimates and four times its variance. For the covariance matrix $\boldsymbol{Q}$, the scaling matrix is chosen to be a fraction of the corresponding OLS estimates (multiplied with the corresponding degrees of freedom). Finally, we have to choose a value for the hyperparameter κ_Q^2, governing the time variation in the state equation. In particular, we assume $\kappa_Q^2 = 0.015$ (following Paul, 2020). We use a rather conservative value for this hyperparameter such that the time variation is not inflated by our prior.

Last, we discuss the prior density on the covariance matrix $\boldsymbol{\Sigma}$, which is defined as follows

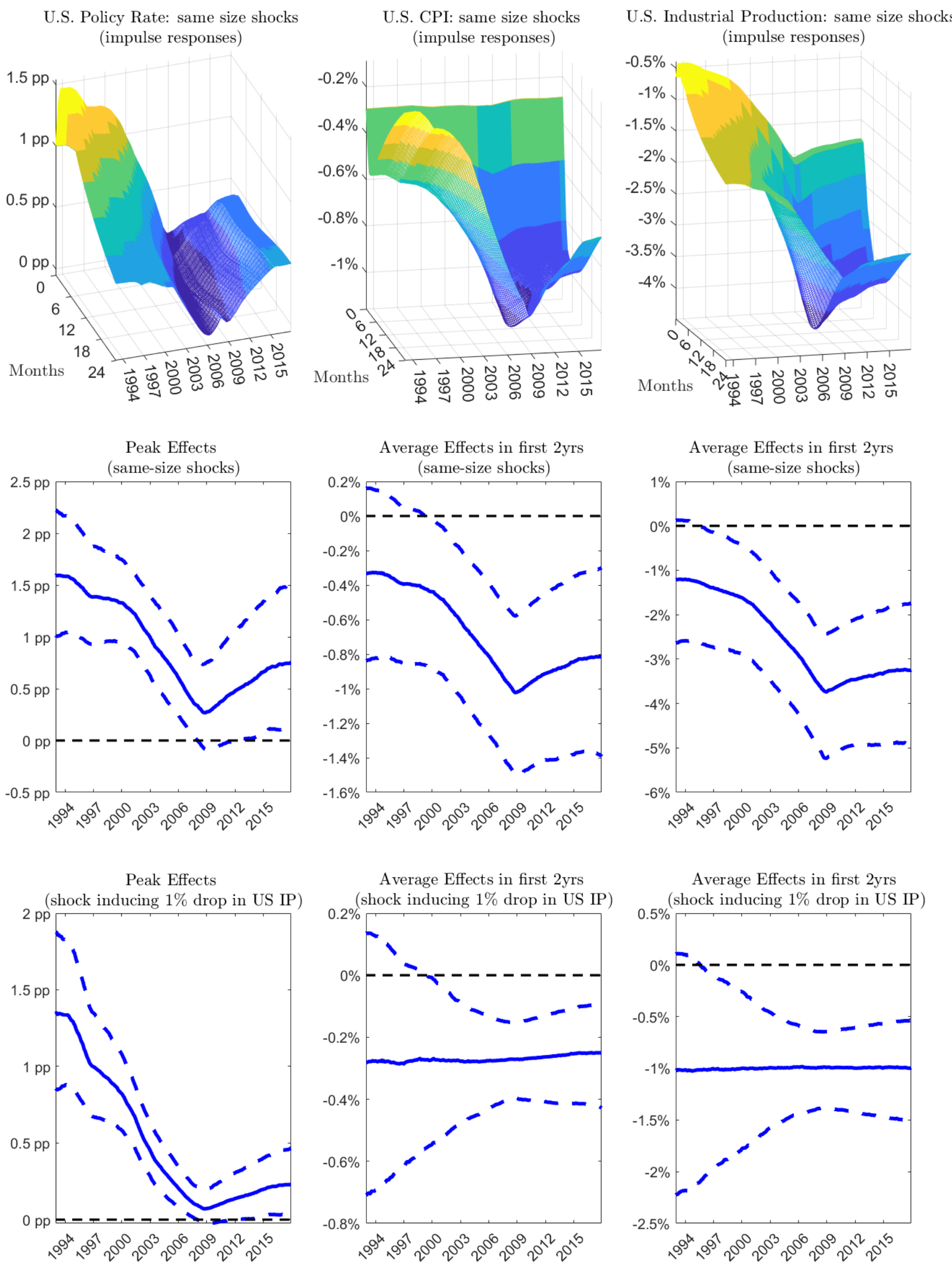
$$\boldsymbol{\Sigma} \sim iW\left(\boldsymbol{I}_M, M + 1\right),$$

where the scaling matrix is set to an identity matrix and the degrees of freedom are set to $M + 1$, as recommended by Karlsson (2013).

Regarding the estimation procedure, we set up the Gibbs sampler along the lines of Del Negro and Primiceri (2015) to obtain posterior distributions. In particular, we use Kalman filtering techniques to obtain the unobservable states in $\boldsymbol{\theta}^T = (\boldsymbol{\theta}_1', \ldots, \boldsymbol{\theta}_T')'$. This results in a Gaussian state space model, where standard Bayesian methods for the Kalman filter can be applied (Carter and Kohn, 1994; Frühwirth-Schnatter, 1994). The remaining posterior quantities are rather standard, and inference is conducted via an MCMC algorithm.

C. TVP-VAR: domestic effects

Figure C1: Time-varying impulse responses functions of U.S. variables.



Notes: Responses of U.S. variables to a contractionary U.S. monetary policy shock. Top panels: evolution of the median impulse-response functions over time ("same size shock" normalization). Mid panels: average effects in the first two years after the shock, using the "same size shocks" normalization. Bottom panels: average effects in the first two years after the shock, using normalization that induces a 1% average contraction in U.S. industrial production over the same period. Given that monetary shocks generate only temporary significant effects on the policy rate, we report peak effects for such a variable. Dashed lines: 68% posterior credible sets. As shocks tend to generate only temporarily significant effects on the U.S. policy rate, we report peak effects of this indicator.

D. Additional results of the sensitivity analysis

Here, we discuss several robustness and sensitivity checks of the baseline results. To limit the number of figures, we report only the results for the average effects in the first two years after the shock based on our baseline "same-size shocks" normalization. Results are reported in Figure D1.

Alternative monetary policy instruments. There exists a range of high-frequency instruments for identifying monetary policy. We employ the monetary policy instrument of Miranda-Agrippino and Ricco (2021), which accounts for the so-called information effect (Melosi, 2017; Nakamura and Steinsson, 2018) by projecting market-based monetary surprises onto the Fed's information set, proxied by Greenbook forecasts. For robustness, we re-estimate the model using several alternative instruments: the original measure by Gertler and Karadi (2015); the "poor man's" refinement by Jarociński and Karadi (2020), which removes the information effect by exploiting high-frequency co-movements between federal funds futures surprises and stock prices; and the instrument by Bauer and Swanson (2023a), which controls for the “Fed response to news” channel by regressing monetary surprises on pre-announcement economic and financial variables. Finally, to account for possible time variation in the informational content of the Greenbook forecasts, we adopt the instrument proposed by Hoesch et al. (2023) (available up until 2015), which extends Miranda-Agrippino and Rey (2020) by allowing the projection of market-based monetary surprises on the Fed's information set to vary over time. Across all specifications, we find highly robust results, as reported in the first row of Figure D1.

Zero lower bound. The presence of the zero lower bound in our sample may affect our estimates. We partly took care of this issue by considering the shadow rate of Wu and Xia (2016). To address this concern further, we re-estimate our model using a more limited sample that excludes the period of the zero lower bound (ending in 2008M11). In addition, to make sure our results are not driven by specific high-frequency non-conventional monetary surprises that occurred during the zero lower bound period (e.g. quantitative easing and/or forward guidance interventions), we run an additional check in which we orthogonalize the instrument of Miranda-Agrippino and Ricco (2021) (that is aimed at picking up conventional monetary policy decisions) with respect to the forward guidance and large-scale asset purchases factors identified by Swanson (2021). The results, reported in the second row of Figure D1, align well with our baseline estimations.

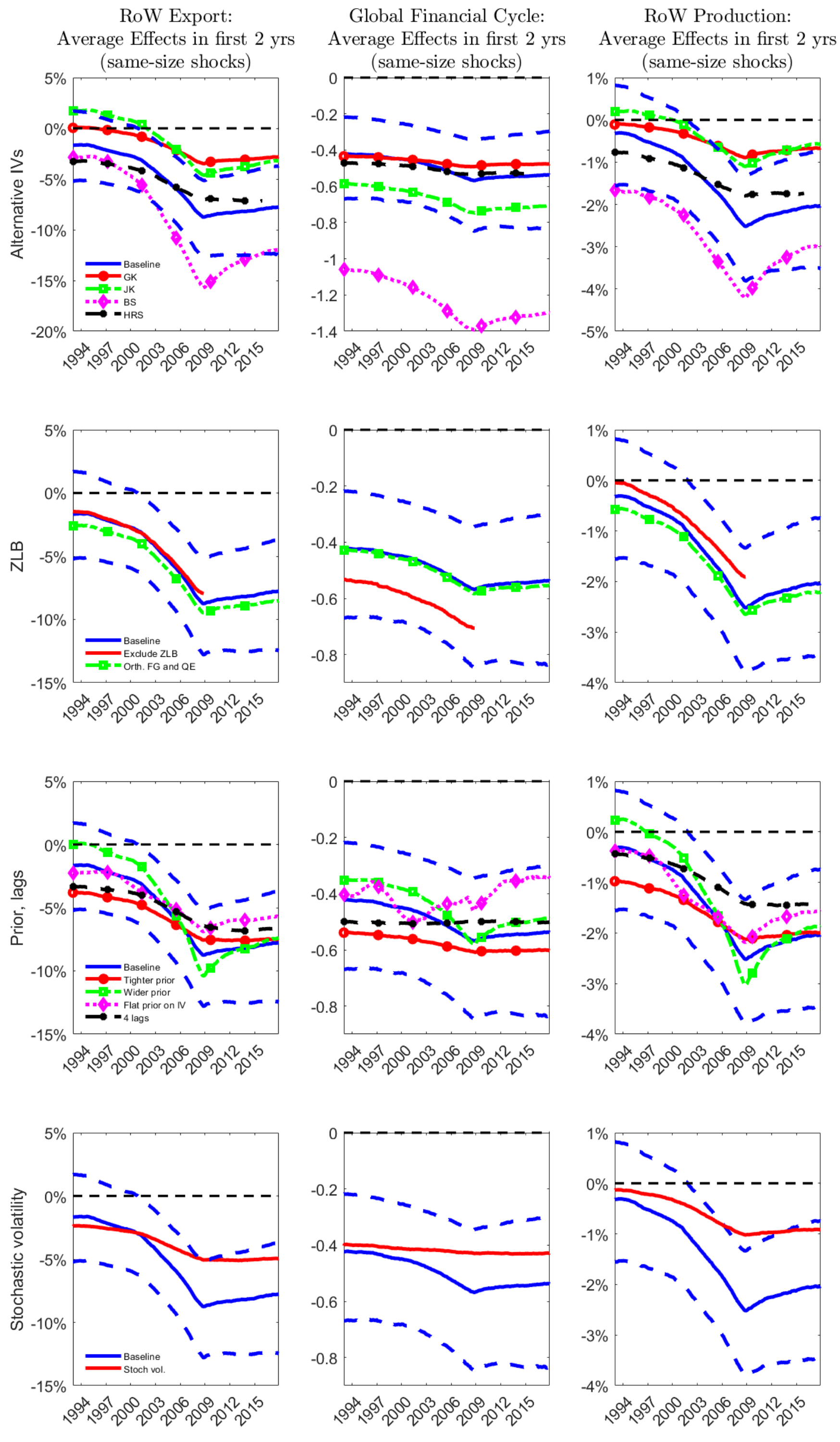
Prior selection and lags. We investigate the sensitivity of our results to different calibrations of κ_Q^2, which regulates the time-variation in the reduced-form VAR coefficients. In the baseline, we use $\kappa_Q^2 = 0.015$, which corresponds to the value used in Paul (2020). The fourth row of Figure D1 shows that the responses are qualitatively similar when using both tighter ($\kappa_Q^2 = 0.01$) and wider choices ($\kappa_Q^2 = 0.02$). In addition, we also obtain very similar results when using a flat prior on the coefficients of the monetary proxy $\boldsymbol{B}_t$ (rather than a relatively informative prior based on the pre-sample linear VAR, where the proxy contained many zeros).

In addition, we show that our results remain similar when incorporating four lags in the VAR (as opposed to our baseline choice of three). Using our estimation setup – which is the most widely used estimation procedure for time-varying VARs (Primiceri, 2005; Paul, 2020) – the Kalman filter breaks down and produces no estimates when more lags are included. This issue is a well-known problem in the estimation of time-varying VARs. Hence, to allow more lags, we estimate three four-variable VARs, each including: the U.S. policy rate, U.S. CPI, U.S. industrial production (these three variables are needed to generate the agents' information set for shocks' transmission), and – one at a time – RoW production, the global financial cycle, and RoW export. The reduced number of variables allows us to increase the number of lags to four (which would be otherwise unfeasible), and the results remain robust.

Stochastic volatility. We also extend the model with stochastic volatility to capture heteroskedasticity. As in Primiceri (2005), we use the following factorization of the now time-varying covariance matrix: $\boldsymbol{\Sigma}_t = \boldsymbol{A}_t^{-1}\text{diag}\left(\exp(h_{1t}), \ldots, \exp(h_{nt})\right) \boldsymbol{A}_t^{-1'}$. The matrix $\boldsymbol{A}_t^{-1}$ is lower uni-triangular and the free elements follow a random walk. Similarly, the log-volatilities in h_{it} ($i = 1, \ldots, n$) also follow a random walk. We reduce the number of lags to two to reduce the system dimension, as the model is already heavily parameterized, and allowing for stochastic volatility increases the issue even more.[1] The results of this exercise are shown in the last row of Figure D1. When comparing the global responses to the baseline estimates, the qualitative pattern remains unchanged. Time variation in real spillovers is present—albeit to a lesser extent. The recessionary effects on RoW production increase from -0.1% in 1993 to -1% in 2008, and a similar pattern emerges for global trade. In stark contrast, consistent with our benchmark estimates, the responses of the global financial cycle, while always negative, remain essentially flat over time.

[1] As in our benchmark analysis and in Primiceri (2005), we use OLS estimates on a pre-sample (1980M1-1992M12) to calibrate the prior distributions. In addition, a prior belief on the extent of time-variation in $\boldsymbol{A}_t^{-1}$ and h_{it} ($i = 1, \ldots, n$) must be specified. Using the notation of Primiceri (2005), this boils down to a selection choice on three parameters: κ_Q^2 (governing time-variation of autoregressive coefficients), κ_W^2 (variance of the residuals), κ_S^2 (covariance of the residuals). We set $\kappa_Q^2 = 0.015$ (as before), $\kappa_W = 0.001$, and $\kappa_S = 0.001$. The value of κ_W^2 is among the ones considered by Primiceri (2005), while our κ_S is relatively tighter (to avoid ill behaviors in our relatively shorter sample). We estimate the stochastic volatility model as in Kim et al. (1998) but with the refinement of Omori et al. (2007).

Figure D1: Robustness Checks.



Notes: Robustness checks. Responses of the global cycles to a contractionary U.S. monetary policy shock. Average effects in the first two years after the shock, using the "same size shocks" normalization. Blue lines are the effects in the baseline specification (with 68% posterior credible sets). First row: alternative monetary policy instruments. GK = Gertler and Karadi (2015); JK = Jarociński and Karadi (2020); BS = Bauer and Swanson (2023b); HRS = Hoesch et al. (2023). Second row: treatment of the ZLB. Third row: different priors on time variation and instrument; different lag choice. Fourth row: VAR with stochastic volatility.

E. TVP-VAR: Alternative financial variables and transmission channels

Alternative global financial variables. Figure E1 presents the time-varying response of additional global financial indicators to U.S. monetary policy shocks. To limit space, we report only the results based on our baseline "same-size shocks" normalization. Specifically, we examine the responses of an alternative measure of global stock prices—the MSCI index excluding the U.S.—alongside various proxies for global liquidity and capital flows. Except for the MSCI index, these additional data are taken from Miranda-Agrippino and Rey (2020) and are available only until 2012, which we set as the endpoint for these estimates (all variables are transformed to stationarity by considering log-differences). Consistent with our benchmark results, we find that global financial conditions generally contract following an exogenous tightening of U.S. monetary policy. However, in line with the response of the global financial cycle discussed in the main text, we find little evidence of time variation in these effects.

Wider propagation channels. Figure E2 presents the time-varying responses of a range of relevant macroeconomic and financial variables to get a better understanding of the transmission of U.S. monetary policy shocks. To compute the impulse responses, we augment the baseline VAR by one variable at a time, which results in specifications with a total of seven variables. Since the state-space would become too large to estimate sensible results, we reduce the number of lags to two. Estimation and prior specification are kept unchanged (trending variables are transformed to stationarity by considering log-differences).

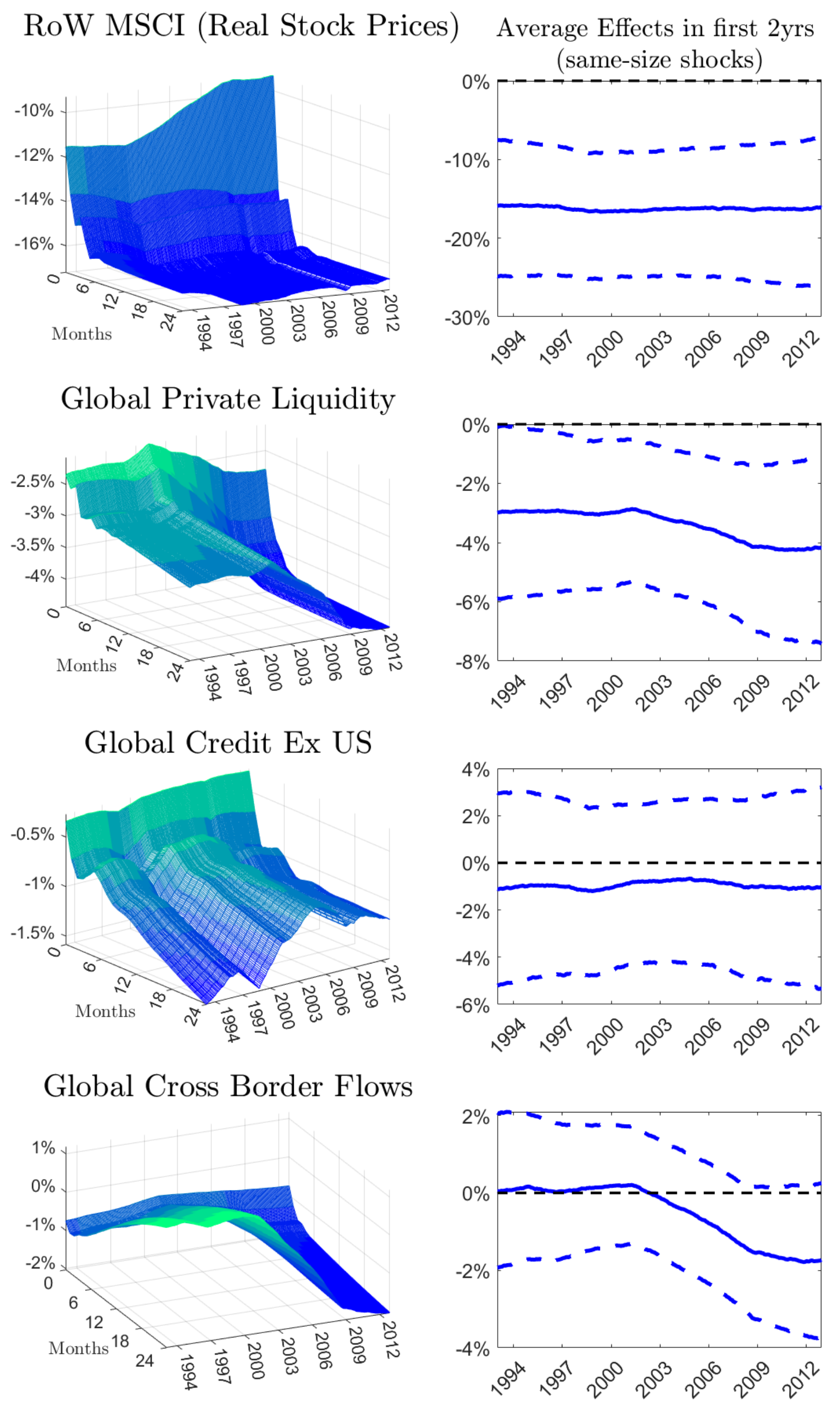
On the domestic level, we find that U.S. monetary tightenings are followed i) by abrupt increases in U.S. corporate credit spreads - proxied by the excess bond premium (EBP) of Gilchrist and Zakrajšek (2012); ii) an appreciation of the U.S. dollar effective exchange (retrieved from BIS); and iii) a fall in U.S. import (proxied with the index in the Database of Global Economic Indicators of the Federal Reserve Bank of Dallas). The signs of these effects are as expected. First, Caldara and Herbst (2019) highlight the role of financial conditions in transmitting monetary policy shocks. Similar to their findings, an increase in the EBP is associated with a tightening of financial conditions as expected through the (domestic) risk-taking channel of monetary policy. Second, the appreciation of the dollar is expected by the uncovered interest rate parity. Third, the negative response of the U.S. import implies no discernible expenditure-switching channel, as expected in the dominant currency paradigm (Gopinath et al., 2020). The dollar appreciation leads in principle to an increase in the competitiveness of foreign goods and to a boost in imports. However, if most of these imports are already priced in dollar, this counteracting force vanishes (see e.g. Degasperi et al., 2020 and Camara et al., 2024 for similar results). Hence, U.S. import contracts as domestic production is falling. Interestingly, while the responses of the exchange rate and EBP are relatively constant over time, we find that U.S. import contracts more and more over time, mirroring the effects we find on global trade (baseline specification).[2]

We now turn to the global responses in Figure E2. In line with the dominant currency paradigm, we find that U.S. tightenings result in some inflationary pressures in the RoW (via the revaluation of the dollar and a widespread surge in import prices; RoW CPI is retrieved from the Federal Reserve Bank of Dallas). Particularly interesting is the decline in inflation spillovers over time (from about +2.1% to +0.1% at peak), which can be attributed to the rise of global value chain participation (Georgiadis et al., 2019). The inflationary pressures in the RoW are tackled with an endogenous increase in interest rates from the major central banks, proxied by the policy rate indicator for RoW economies.[3] The gradual decrease in the policy rate response that we find in the data is consistent with i) the increasing effects on RoW production and ii) the diminishing inflation spillovers.

[2] A growing U.S. import leakage over time in response to U.S. monetary policy shocks aligns with the findings of Hofmann and Peersman (2024).

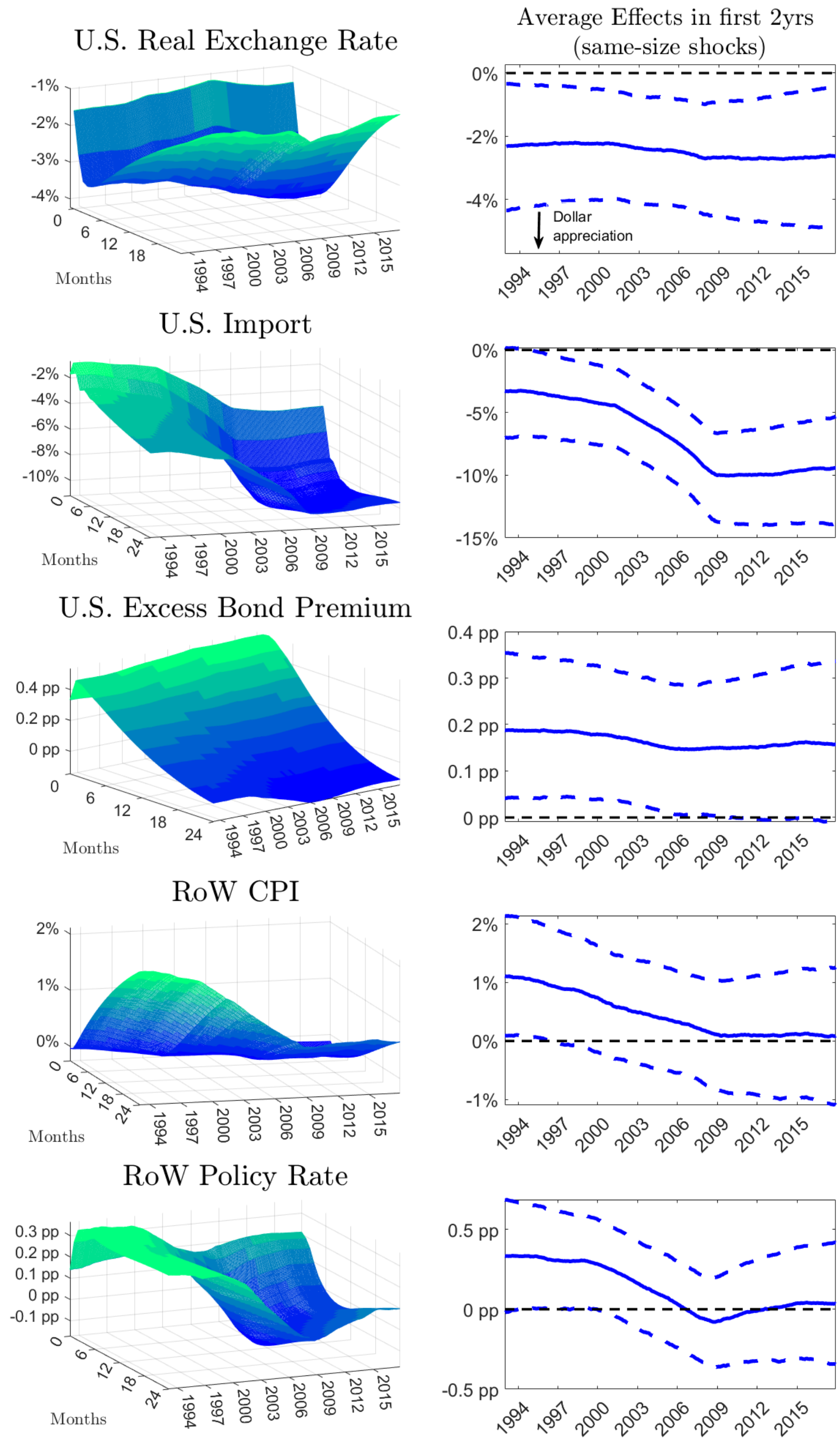
[3] The comprehensive RoW policy rate indicator published by the Federal Reserve of Dallas displays explosive patterns in the 1980s and 1990s, driven by the merging economies' data. This makes the estimation infeasible. Hence, we consider the index for RoW advanced economies as a proxy for RoW policy response.

Figure E1: Global financial spillovers: alternative indicators.



Notes: Responses of additional variables to a contractionary U.S. monetary policy shock. Left column: evolution of the median impulse-response functions over time ("same size shock" normalization). Right column: average effects in the first two years after the shock, using the "same size shocks" normalization. Dashed lines: 68% posterior credible sets.

Figure E2: Wider propagation channels.



Notes: Responses of additional global financial variables to a contractionary U.S. monetary policy shock. Left column: evolution of the median impulse-response functions over time ("same size shock" normalization). Right column: average effects in the first two years after the shock, using the "same size shocks" normalization. Dashed lines: 68% posterior credible sets.

F. Additional results of country-specific effects

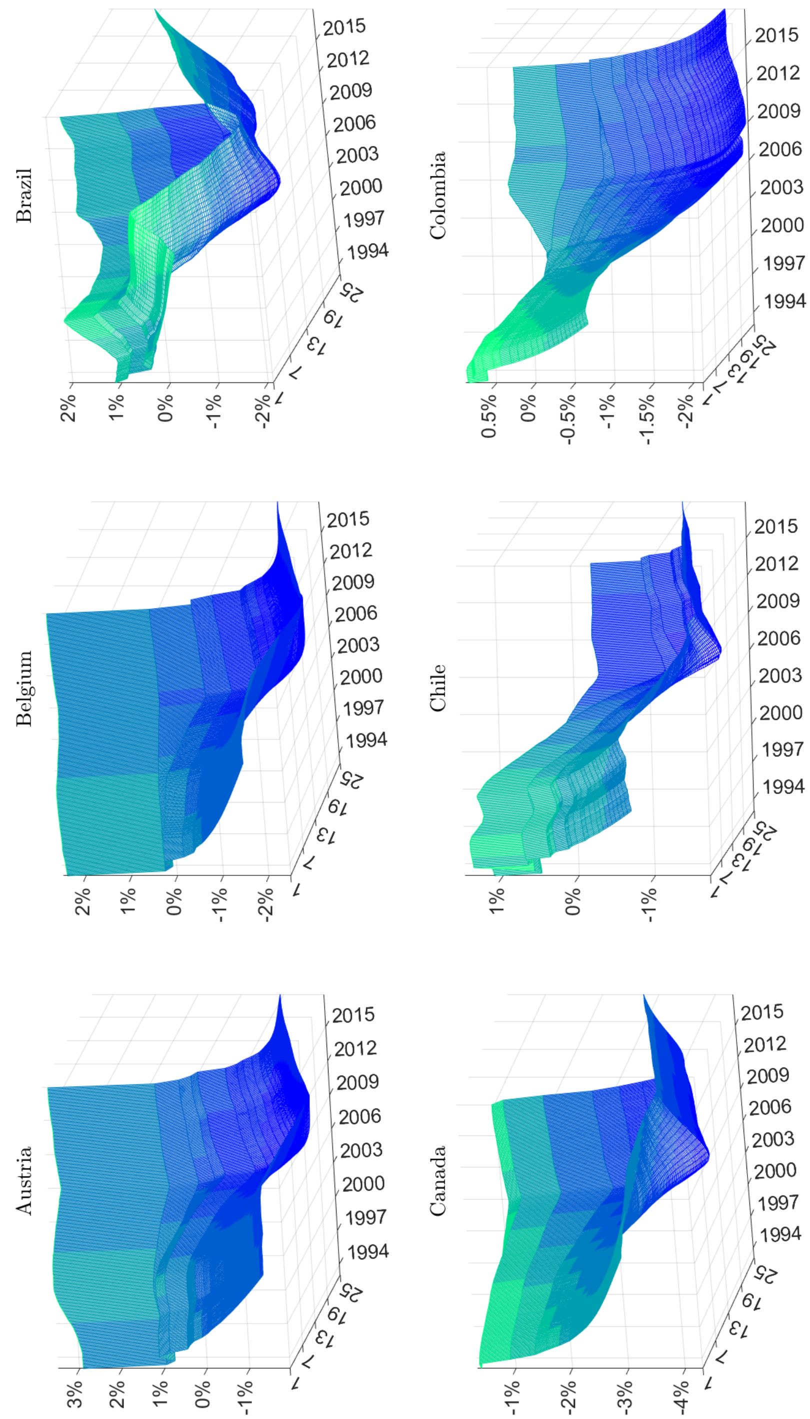
In this Appendix, we report additional results of country-specific effects. Figure F1 - Figure F5 report the country-specific impulse responses of industrial production for each country separately, while in Figure F6 we show the evolution of the average responses of country-specific industrial production over time.

Figure F7 replicates the analysis from Figure 7 using alternative measures of financial integration. In addition to the KOF index (our baseline measure), which offers a comprehensive assessment, we incorporate additional metrics, including: (i) the ratio of a country's dollar-denominated external assets and liabilities to GDP (data from Bénétrix et al., 2019); (ii) external assets alone; (iii) external liabilities alone; and (iv) the Chinn-Ito Index of capital account openness (Chinn and Ito, 2006). The results, presented in Figure F7, reveal patterns that closely align with those observed in our baseline estimates.[4]

In Figure 7, we report the graphical illustration of two sets of regressions: i) average mean effects in country i as a function of average financial and trade integration in country i; ii) change in the effects in country i as a function of change in financial and trade integration in country i. These findings are detailed in Table F1. Furthermore, our results remain consistent when we shift focus from mean effects in the first two years to peak effects and effects at selected horizons (0, 6, 12, 18, 24 months) – see Table F1.

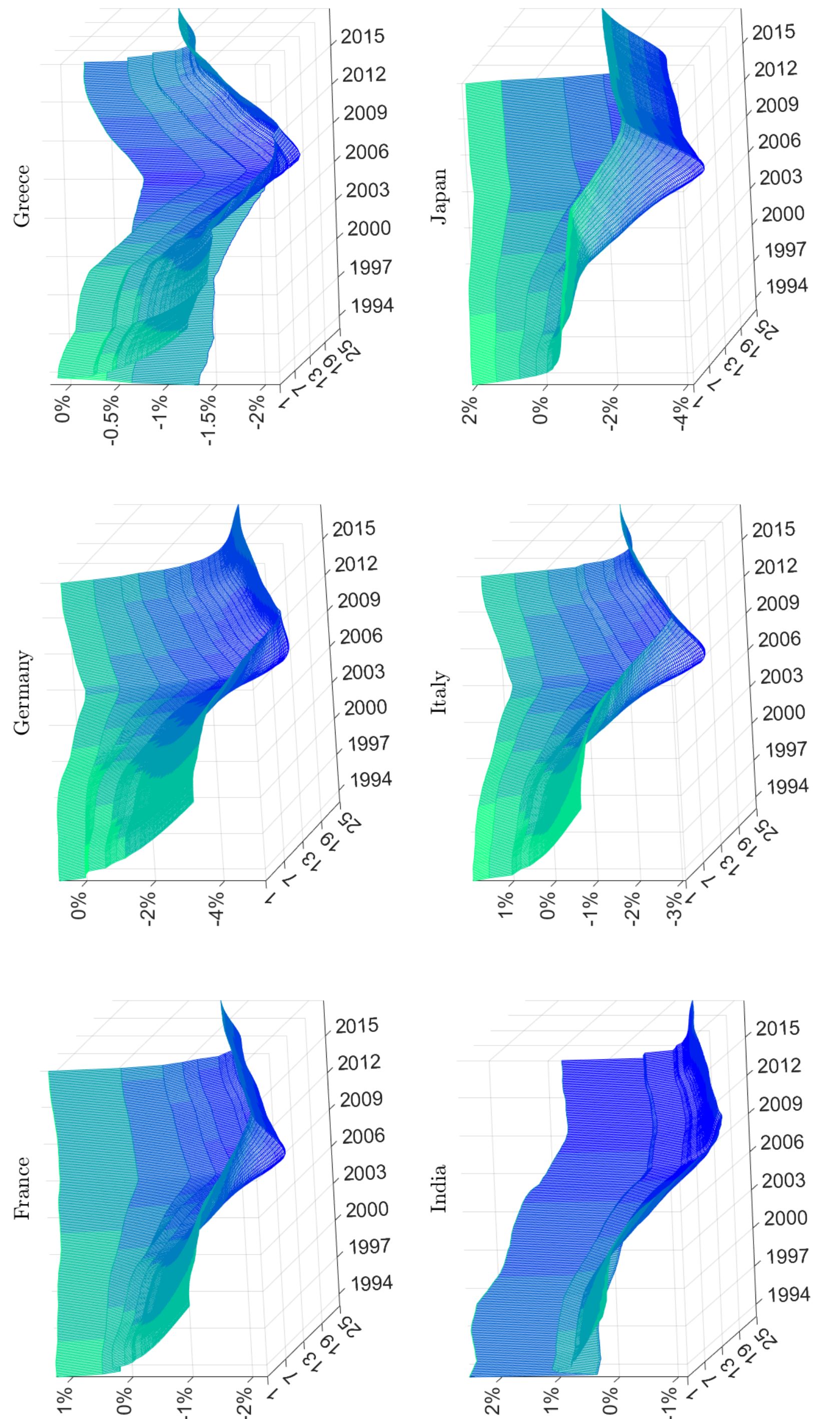
[4] The measures of (dollar-denominated) external assets and liabilities are subject to well-documented measurement issues, such as the financial center bias—particularly relevant when examining changes over time for the UK and the Netherlands. To mitigate potential biases, we add dummies for these two countries in related plots in the bottom panels of Figure F7. For Brazil, Chile, and Switzerland, the Chinn-Ito index of capital account openness is missing in 1993, so we substitute the missing values with the earliest available observation. We apply a similar approach for Colombia, where the index is recorded as zero in 1993.

Figure F1: Country-specific impulse responses of industrial production: Austria, Belgium, Brazil, Canada, Chile, and Colombia.



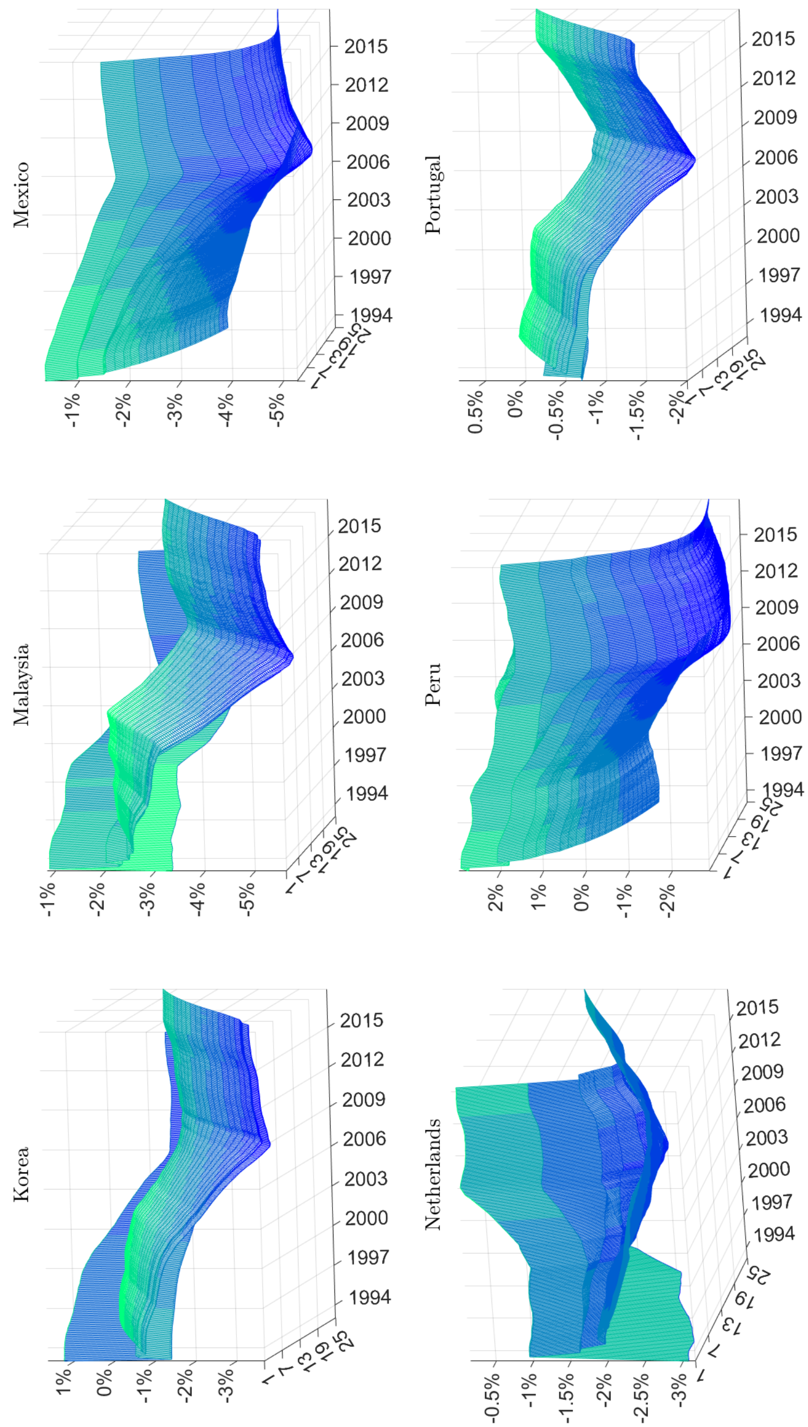
Notes: Responses to a contractionary U.S. monetary policy shock that induces a one percentage point (pp) increase in the federal fund rate in 1993M1.

Figure F2: Country-Specific impulse responses of industrial production: France, Germany, Greece, India, Italy, and Japan.



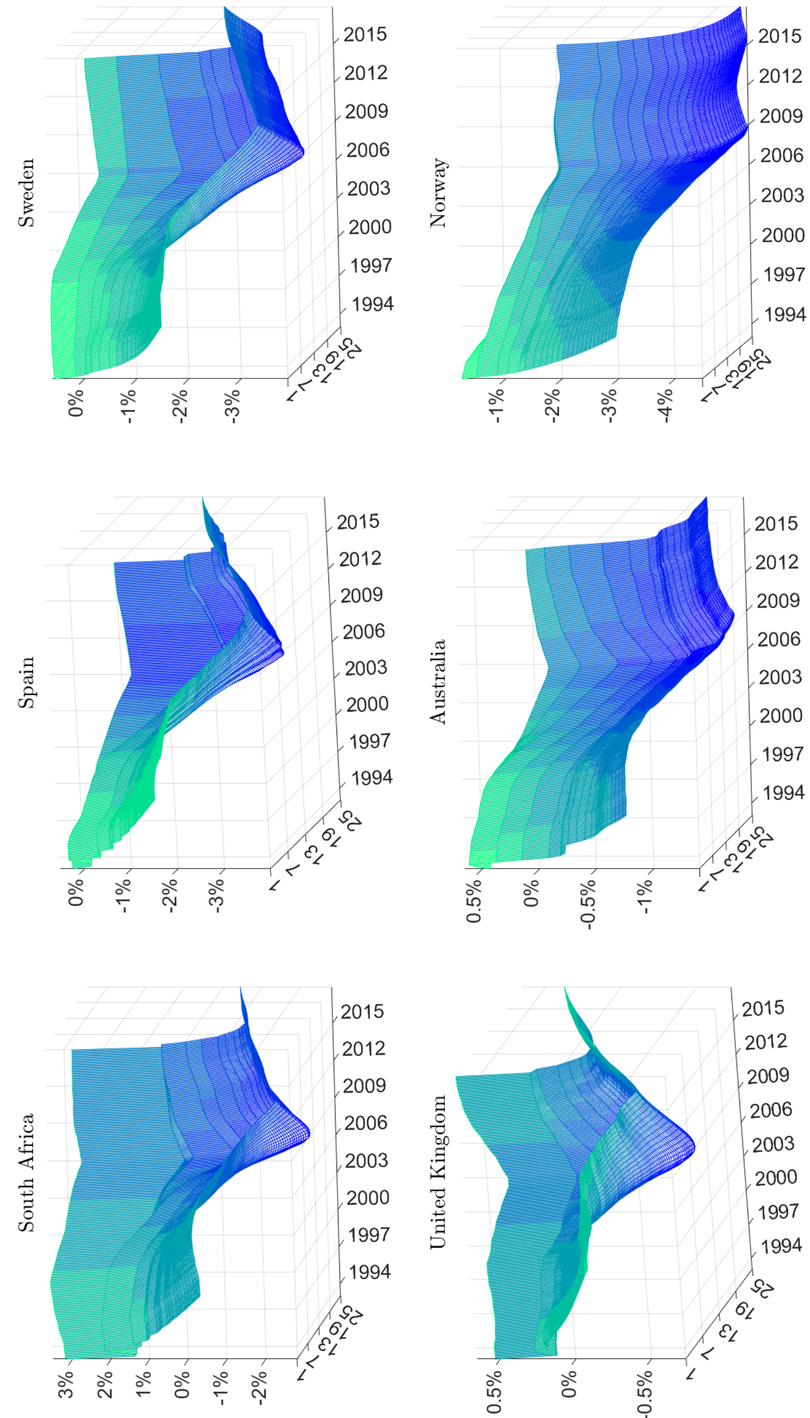
Notes: Responses to a contractionary U.S. monetary policy shock that induces a one percentage point increase in the federal fund rate in 1993M1.

Figure F3: country-specific impulse responses of industrial production: Korea, Malaysia, Mexico, Netherlands, Peru, and Portugal.



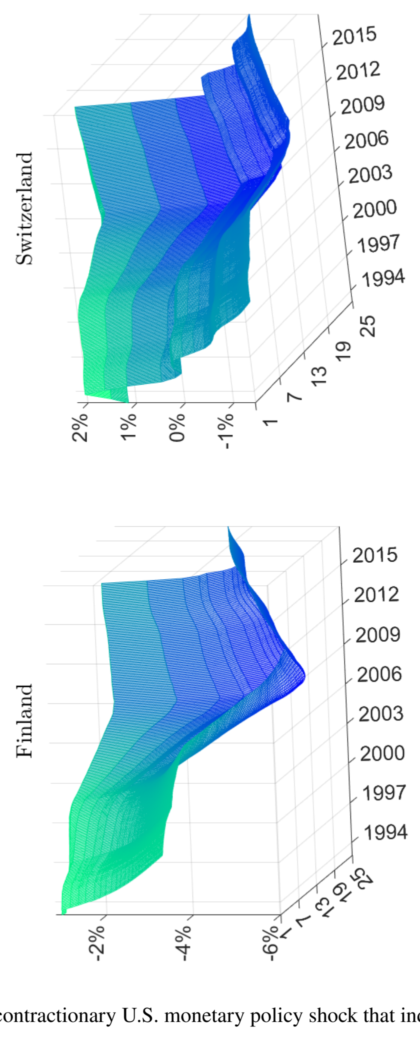
Notes: Responses to a contractionary U.S. monetary policy shock that induces a one percentage point increase in the federal fund rate in 1993M1.

Figure F4: country-specific impulse responses of industrial production: South Africa, Spain, Sweden, United Kingdom, Australia, Norway.



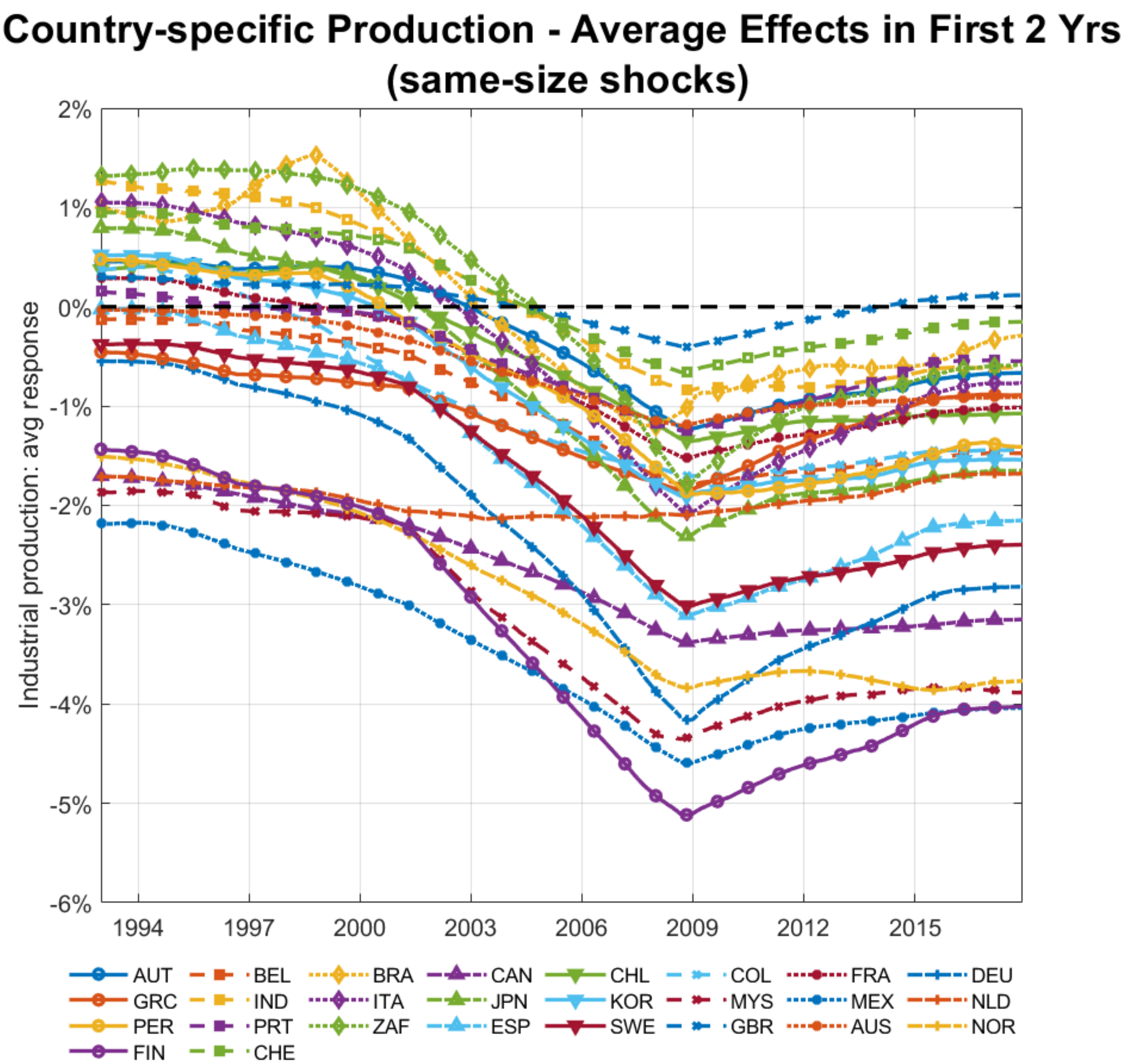
Notes: Responses to a contractionary U.S. monetary policy shock that induces a one percentage point increase in the federal fund rate in 1993M1. Countries: Peru, South Korea, Malaysia, India.

Figure F5: Country-specific impulse responses of industrial production: Finland, and Switzerland.



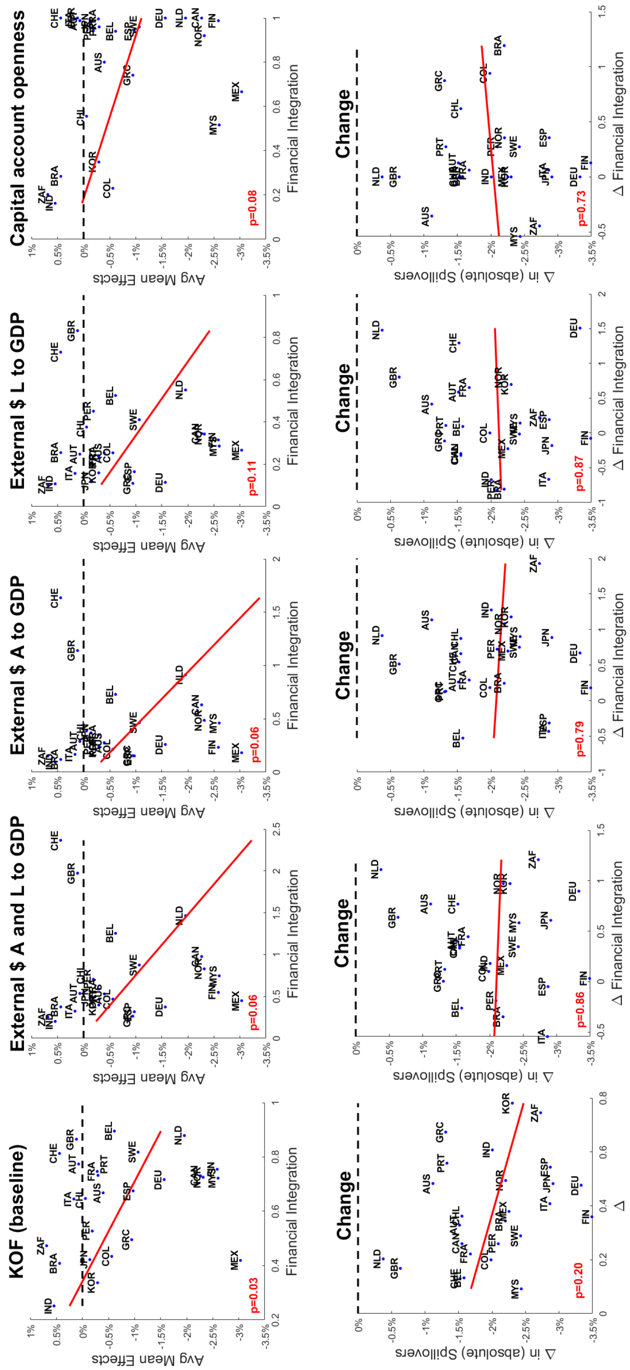
Notes: Responses to a contractionary U.S. monetary policy shock that induces a one percentage point increase in the federal fund rate in 1993M1.

Figure F6: Time-varying impulse responses of country-specific industrial production.



Notes: average effects of country-specific (non-U.S.) industrial production indices to a contractionary U.S. monetary policy in the first two years after the shock, using the "same size shocks" normalization.

Figure F7: Correlations of spillovers and additional financial integration measures.



Notes: Upper panels: Relationship between the average mean response of individual countries during the period 1993-2008 and their average financial during the same period. Lower panels: Relationship between the absolute change in country-specific mean response and the corresponding change in financial integration. Red lines: univariate regressions interpolating the points. Given our small sample, we reduce the impact of outliers by employing an iteratively reweighted least-squares algorithm.

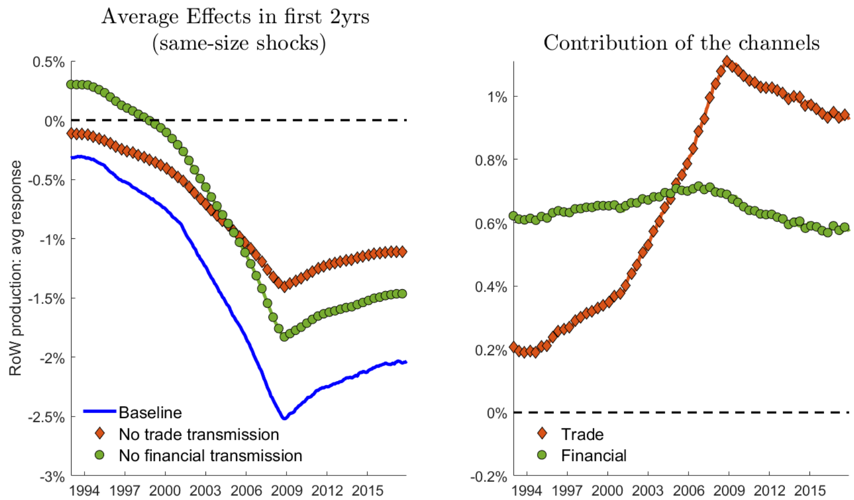
Table F1: Regression outcomes.

	Mean	*Peak*	$h = 0$	$h = 6$	$h = 12$	$h = 18$	$h = 24$
				Average			
Trade integration	**−1.40**	**−1.68**	**−2.91**	**−1.54**	**−1.41**	**−1.22**	−1.09
	(0.55)	(0.60)	(0.82)	(0.56)	(0.57)	(0.60)	(0.66)
Financial integration	**−2.72**	−1.91	−2.99	−2.16	**−2.77**	**−3.20**	**−3.48**
	(1.19)	(1.38)	(2.14)	(1.31)	(1.21)	(1.21)	(1.29)
				Time-variation			
Δ Trade integration	**−1.45**	−1.37	−0.39	**−1.96**	**−1.87**	−1.43	−1.09
	(0.71)	(1.08)	(0.89)	(0.87)	(0.83)	(0.70)	(0.65)
Δ Financial integration	−1.14	−0.85	−1.90	−1.34	−1.11	−1.00	−0.96
	(0.87)	(1.24)	(0.93)	(1.06)	(1.01)	(0.85)	(0.75)

Notes: Regression based on 26 observations with *average* or *time-variation* effects in industrial production per country as dependent variable. The independent variable is trade or financial integration, either at the level or as a log change. In parenthesis we report robust standard error of the slope coefficients. Bold numbers indicate statistical significance at 5% level.

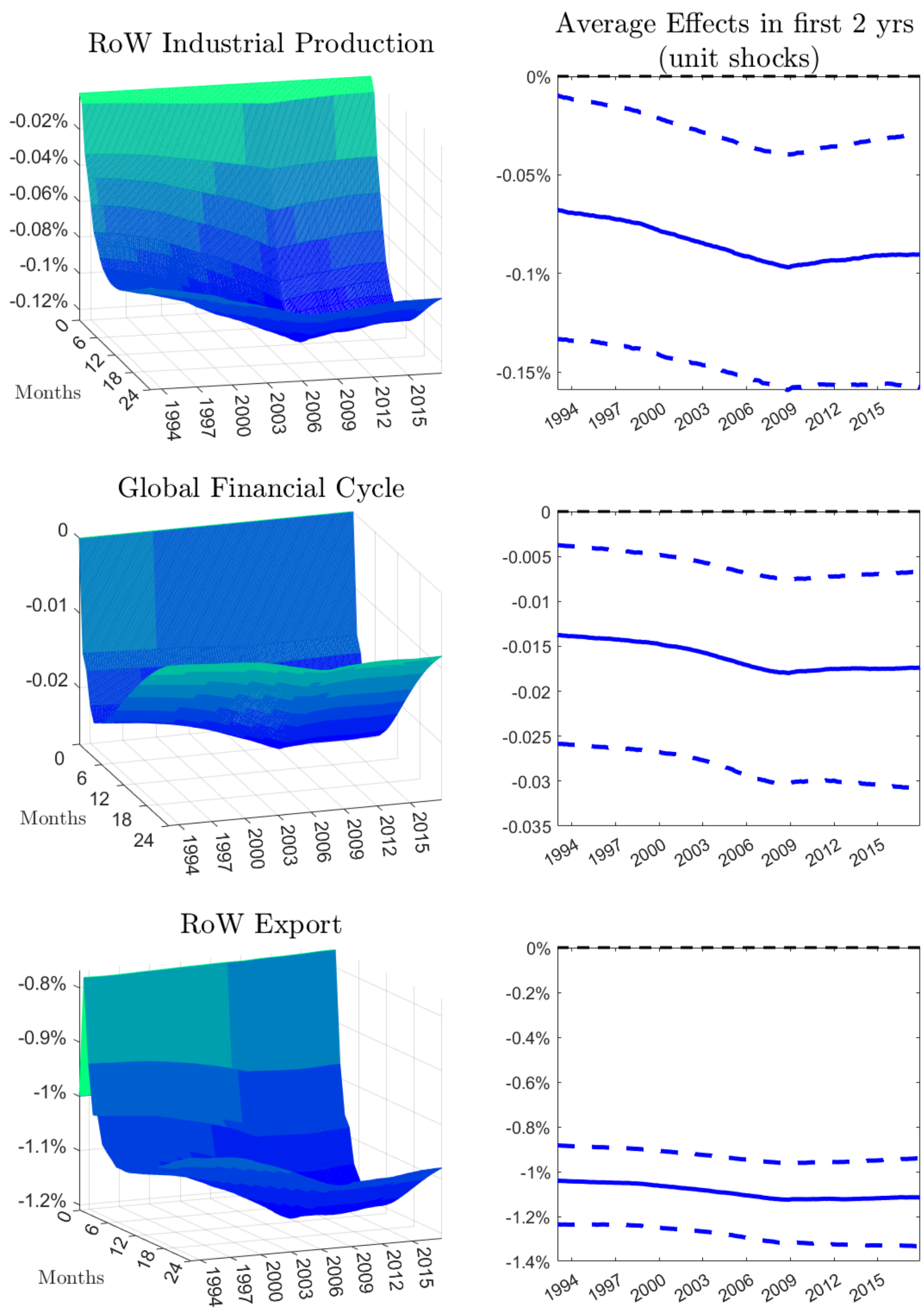
G. Responses of global trade and global financial shocks

Figure G1: Counterfactual analysis – relaxing the identifying assumptions.



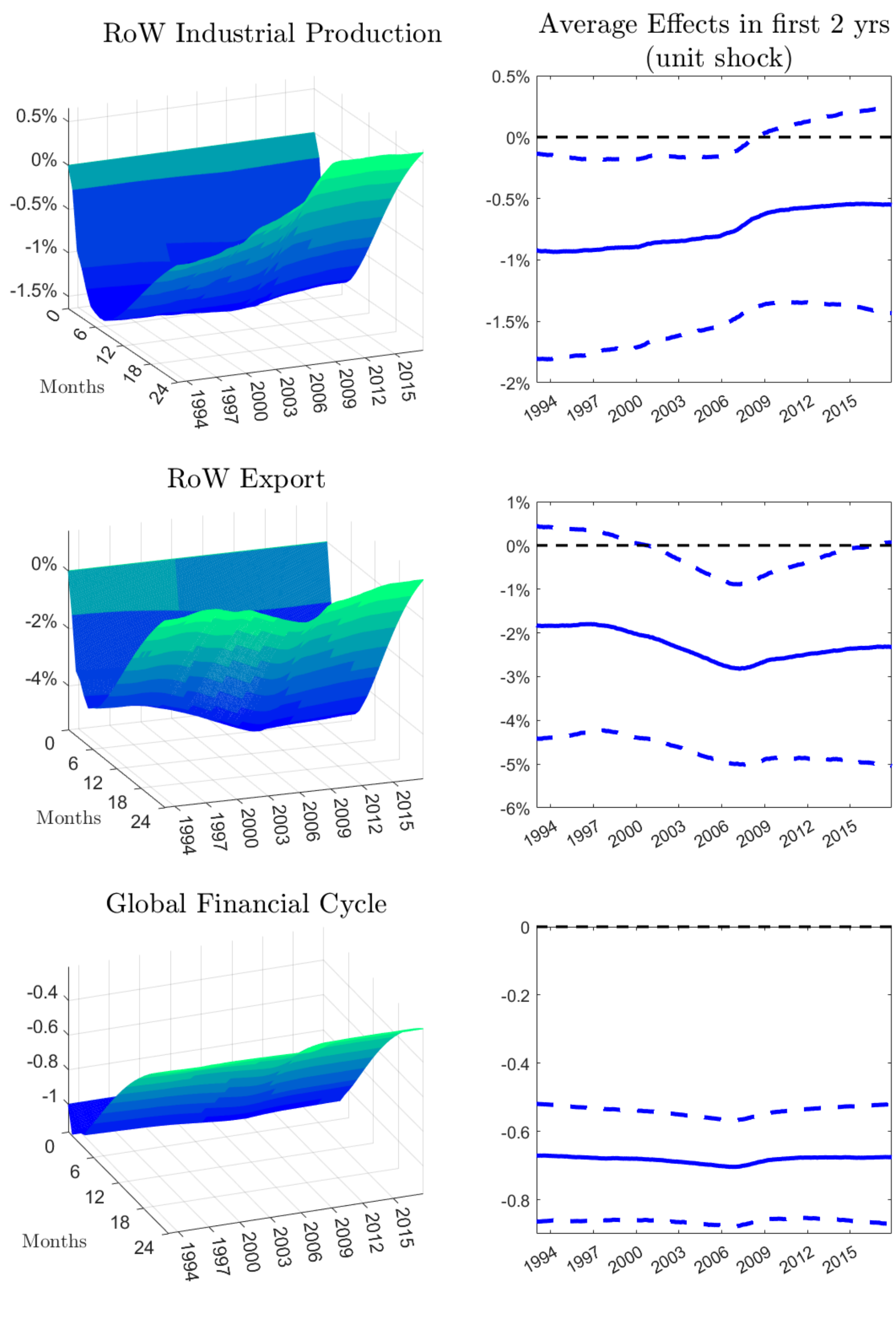
Notes: Global trade and financial shocks are identified employing recursive VARs where the indicators – RoW export and global financial cycle – are placed in second position (after RoW industrial production). Left panel: Peak responses of world industrial production to U.S. monetary shocks (blue line: baseline estimations) versus counterfactual scenarios obtained by zeroing out the response of RoW exports (orange diamond lines: "no trade transmission") and the global financial cycle's response to U.S. monetary shocks via global financial shocks (green circled lines: "no financial transmission"). Right panel: Absolute contribution of the two channels over time.

Figure G2: Global trade shock: time-varying impulse responses functions of global variables.



Notes: Responses of global variables to a contractionary global trade shock, normalized to generate a 1% impact contraction in RoW export at *t* (notice that the normalization of the global trade shocks has no impact on the results of the counterfactual simulation). Left column: evolution of the median impulse-response functions over time. Right column: average effects in the first two years after the shock.

Figure G3: Global financial shock: time-varying impulse responses functions of global variables.



Notes: Responses of global variables to a contractionary global financial shock, normalized to generate a -1 point impact contraction in the Global Financial Cycle at t (notice that the normalization of the global financial cycle shocks has no impact on the results of the counterfactual simulation). Left column: evolution of the median impulse-response functions over time. Right column: average effects in the first two years after the shock.

H. Our results through the lens of a two-country model with trade and financial frictions

As our empirical results are novel and point to a stronger transmission via the trade channel over time, this section proposes a theoretical model that rationalizes the link between rising output and trade spillovers. The theoretical model is a two-country New Keynesian DSGE model, which features international trade and international financial market frictions. The model consists of two countries, which we label U and R, as stand-ins for the U.S. and RoW, respectively. The two countries are symmetric in most aspects, but we assume two asymmetries, which mimic the central role of the U.S. in the international trade and financial landscape. For international trade, we assume dominant-currency pricing (Gopinath, 2015; Gopinath et al., 2020). On the financial market, we assume, additionally to the financial accelerator framework of Gertler and Karadi (2011), an additional adverse feedback loop arising from foreign-currency debt holdings as described in Akinci and Queralto (2024). A few remarks are in order. One limitation of the model is that, to keep it simple, we abstract from third-country trade (trade within RoW), which can create additional output spillovers through asymmetric trade invoicing (Georgiadis and Schumann, 2021), network effects (Bräuning and Sheremirov, 2023), or composition effects of emerging markets and advanced economies. As the U.S. does not account for the major share of trade across the globe, the amount of spillovers is limited in this model setup. Another limitation is the specific structure of international financial markets we assume in the model. A key aspect of the model is that financial imperfections limit the ability of borrowers to obtain foreign-currency denominated financing. This aspect is likewise more relevant for emerging market economies than for advanced economies. Furthermore, we abstain from the possibility of allowing for trade in claims on foreign assets, and we thus do not consider the role of safe dollar assets in international finance (He and Krishnamurthy, 2019; Jiang et al., 2021, 2024). Our choices are driven by the goal of keeping the model as simple as possible and focus more on the intuition behind the model predictions. For a fully fledged model estimated to match the empirical VAR evidence, see Camara et al. (2024) – where they find a dominant role of trade in explaining the international spillovers of U.S. monetary policy shocks.

H.1 Model description

In the description of the model, we will focus on country U and highlight only the differences to country R. We list all equilibrium conditions in section H.4.

Households. Households in country U consume, supply labor, save in domestic deposits, and own the capital stock. The household maximizes utility, which is derived from consumption and leisure

$$\mathcal{U}^U \qquad = \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left(\frac{(C_t^U - hC_{t-1}^U)^{1-\sigma}}{1-\sigma} - \kappa_L \frac{(L_t^U)^{1+\varphi}}{1+\varphi} \right),$$

where C_t^U is consumption and L_t^U is labor supply in country U. Parameter h governs habit formation in consumption, σ is the coefficient of relative risk aversion, φ is the inverse of the Frisch elasticity, κ_L is the relative weight of labor in the utility function, and β denotes the discount factor. The household budget constraint of country U reads

$$C_t^U + D_{r,t}^U + Q_t^U K_t^U \leq R_{r,t}^U D_{r,t-1}^U + R_{k,t}^U Q_{t-1}^U K_{t-1}^U + W_{r,t}^U L_t^U + \Gamma_t^U - T_t^U,$$

where $R_{r,t}^U$ is the real interest rate on the household's real deposits, $D_{r,t}^U$. The households in country U own directly the capital stock, K_t^U, where Q_t^U denotes the price of capital, and $R_{k,t}^U$ its return. $W_{r,t}^U$ is the real wage set by the household. The term Γ_t^U denotes profits and T_t^U tax receipts. The budget constraint in country R is different, as households do not hold capital directly. It is then defined as

$$C_t^R + D_{r,t}^R \leq R_{r,t}^R D_{r,t-1}^R + W_{r,t}^R L_t^R + \Gamma_t^R - T_t^R.$$

This setting in country U and R implies the following first-order conditions:

$$\varrho_t^U = (C_t^U - hC_{t-1}^U)^{-\sigma} - \beta h\left(\mathbb{E}_t C_{t+1}^U - hC_t^U\right)^{-\sigma}, \tag{H.1}$$

$$\Lambda_{t,t+1}^U = \frac{\varrho_{t+1}^U}{\varrho_t^U}, \tag{H.2}$$

$$W_{r,t}^U = \kappa_L (L_t^U)^{\varphi} (C_t^U)^{\sigma}, \tag{H.3}$$

$$1 = \beta R_{r,t+1}^U \mathbb{E}_t\left[\Lambda_{t,t+1}\right], \tag{H.4}$$

$$1 = \beta \mathbb{E}_t\left[\Lambda_{t,t+1} R_{k,t+1}^U\right], \tag{H.5}$$

$$R_{r,t+1}^U = \frac{R_{t+1}^U}{\mathbb{E}_t \Pi_{t+1}^U}, \tag{H.6}$$

$$\varrho_t^R = (C_t^R - hC_{t-1}^R)^{-\sigma} - \beta h\left(\mathbb{E}_t C_{t+1}^R - hC_t^R\right)^{-\sigma}, \tag{H.35}$$

$$\Lambda_{t,t+1}^R = \frac{\varrho_{t+1}^R}{\varrho_t^R}, \tag{H.36}$$

$$W_{r,t}^R = \kappa_L (L_t^R)^{\varphi} (C_t^R)^{\sigma}, \tag{H.37}$$

$$1 = \beta R_{r,t+1}^R \mathbb{E}_t\left[\Lambda_{t,t+1}^R\right], \tag{H.38}$$

$$R_{r,t+1}^R = \frac{R_{t+1}^R}{\mathbb{E}_t \Pi_{t+1}^R}. \tag{H.39}$$

Here ϱ_t^U and ϱ_t^R denote the marginal utility of consumption. We assume that the Fisher equation holds, where $\Pi_t^U = P_t^U / P_{t-1}^U$ and $\Pi_t^R = P_t^R / P_{t-1}^R$ denote the gross consumer price inflation rate.

Production and price setting. Intermediate-goods producers produce goods for the domestic and the export market. They employ labor and capital as input for the production of goods, using a Cobb-Douglas production function. The production function and optimality conditions regarding the input factors take the following form

$$P_t^U = (K_{t-1}^U)^\alpha (L_t^U)^{1-\alpha}, \tag{H.7}$$

$$R_{z,t}^U = \alpha \mathcal{MC}_t^U \frac{P_t^U}{K_{t-1}^U}, \tag{H.8}$$

$$W_{r,t}^U = (1-\alpha)\mathcal{MC}_t^U \frac{P_t^U}{L_t^U}, \tag{H.9}$$

$$P_t^R = (K_{t-1}^R)^\alpha (L_t^R)^{1-\alpha}, \tag{H.40}$$

$$R_{z,t}^R = \alpha \mathcal{MC}_t^R \frac{P_t^R}{K_{t-1}^R}, \tag{H.41}$$

$$W_{r,t}^R = (1-\alpha)\mathcal{MC}_t^R \frac{P_t^R}{L_t^R}. \tag{H.42}$$

Here, $P_t^U(j)$ denotes the production volume. $R_{z,t}^U$ is the rental rate of capital, $W_{r,t}^U$ the real wage rate, and $\mathcal{MC}_t^U$ is the real marginal cost of production. Firms produce for the domestic as well as the export market, and thus $P_t^U(j) = H_t^U(j) + \frac{1-n}{n} X_t^U(j)$, where H_t^U denotes output produced for the domestic market and X_t^U denotes goods produced for export to country R. We account for different country sizes with the parameter n. Accordingly, they face two demand functions for their goods, $H_t^U(j) = \left(\frac{P_{H,t}^U(j)}{P_{H,t}^U}\right)^{-\epsilon} H_t^U$, and $X_t^U(j) = \left(\frac{\mathcal{E}_t P_{X,t}^U(j)}{\mathcal{E}_t P_{X,t}^U}\right)^{-\epsilon} X_t^U$, where ϵ denotes the elasticity of substitution between goods produced in the same country. $P_{H,t}^U(j)$ denotes the price for the domestically produced good of variety j and $P_{H,t}^U$ is a price index of domestically produced goods. Similarly, $P_{X,t}^U(j)$ is the price of the export good of variety j, while $P_{X,t}^U$ is the price index of export goods. Note that the export price index is given in foreign currency. Therefore, we use the bilateral nominal exchange rate $\mathcal{E}_t$ between country U and R to transform this back into the local currency. When we aggregate and account for price dispersion, production of final goods in either country equals the demand for these goods at home and abroad, which yields $P_t^U Y_t^U = P_{H,t}^U \Delta_{H,t}^U H_t^U + P_{M,t}^U \Delta_{M,t}^U M_t^U$. Here, $\Delta_{H,t}^U$ denotes price dispersion of domestically produced goods, while $\Delta_{X,t}^U$ denotes price dispersion of goods demanded from abroad.

As intermediate-goods producing firms sell goods both in the domestic and export market, their profit function reads $\Gamma_t^{IG,U}(j) = \frac{P_{H,t}^U(j)}{P_t^U} H_t^U(j) + \frac{1-n}{n} \frac{\mathcal{E}_t P_{X,t}^U(j)}{P_t^U} X_t^U(j) - W_{r,t}^U L_t^U(j) - R_{z,t}^U K_{t-1}^U(j)$. Firms face monopolistic competition and set their prices with a markup over their marginal costs in both markets. Price rigidities are modelled à la Calvo (1983) and firms can change prices freely but only with a certain probability $1-\theta$. We assume that firms in country U set their optimal export prices in their own currency, according to the producer currency price (PCP) paradigm. The first-order condition for the optimally set price, $P^{U,+}_{H,t}$, in the domestic market is

$$\mathbb{E}_t \sum_{k=0}^{\infty} (\beta\theta)^k \Lambda^U_{t,t+k} \left\{ \left(\frac{P^U_{H,t}}{P^U_{H,t+k}} \right)^{-\epsilon} H^U_{t+k} \left[\frac{P^{U,+}_{H,t}}{P^U_{t+k}} - \frac{\epsilon}{\epsilon - 1} \mathcal{MC}^U_{t+k} \right] \right\} = 0,$$

and the respective first-order condition for the optimally set price in the export market, $P^{U,+}_{X,t}$, can be derived as

$$\mathbb{E}_t \sum_{k=0}^{\infty} (\beta\theta)^k \left\{ \frac{1-n}{n} \Lambda^U_{t,t+k} \left(\frac{\mathcal{E}_t P^U_{X,t}}{\mathcal{E}_{t+k} P^U_{X,t+k}} \right)^{-\epsilon} X^U_{t+k} \left[\frac{\mathcal{E}_t P^{U,+}_{X,t}}{P^U_{t+k}} - \frac{\epsilon}{\epsilon - 1} \mathcal{MC}^U_{t+k} \right] \right\} = 0.$$

While both countries are symmetric in most aspects, here we introduce an asymmetry. Motivated by the dominance of the U.S. dollar (the currency of country U) in international goods trade, we assume that firms in country R set their export prices in terms of the currency of their export market, following the local currency paradigm (LCP). Hence, the first-order condition for the optimally set price in the export market, $P^{R,+}_{X,t}$, in country R reads

$$\mathbb{E}_t \sum_{k=0}^{\infty} (\beta\theta)^k \left\{ \Lambda^R_{t,t+k} \left(\frac{P^R_{X,t}}{P^R_{X,t+k}} \right)^{-\epsilon} X^R_{t+k} \left[\frac{P^{R,+}_{X,t}}{\mathcal{E}_{t+k} P^R_{t+k}} - \frac{\epsilon}{\epsilon - 1} \mathcal{MC}^R_{t+k} \right] \right\} = 0.$$

This yields the following set of equations for optimal set prices for the domestic and export markets

$$p^{U,+}_{H,t} = \frac{\epsilon}{\epsilon - 1} \frac{Z^{U,1}_{H,t}}{Z^{U,2}_{H,t}}, \tag{H.10}$$

$$p^{U,+}_{X,t} = \frac{\epsilon}{\epsilon - 1} \frac{Z^{U,1}_{X,t}}{Z^{U,2}_{X,t}}, \tag{H.11}$$

$$p^{R,+}_{H,t} = \frac{\epsilon}{\epsilon - 1} \frac{Z^{U,1}_{R,t}}{R^{R,2}_{H,t}}, \tag{H.43}$$

$$p^{R,+}_{X,t} = \frac{\epsilon}{\epsilon - 1} \frac{Z^{R,1}_{X,t}}{Z^{R,2}_{X,t}}. \tag{H.44}$$

These equations denote the optimal reset price relative to the overall price level in a given segment, i.e., the optimal domestic price is $p^{U,+}_{H,t} = P^{U,+}_{H,t} / P^U_{H,t}$ relative to the price index of domestic goods. Similarly, the optimal export price is defined as $p^{U,+}_{X,t} = P^{U,+}_{X,t} / P^U_{X,t}$ relative to the price index of export goods. The other variables are auxiliary variables to facilitate recursive price Phillips curves with Calvo pricing. These are given by

$$Z^{U,1}_{H,t} = H^U_t \mathcal{MC}^U_t + \beta\theta \mathbb{E}_t \Lambda^U_{t,t+1} (\Pi^U_{H,t+1})^{\epsilon} Z^{U,1}_{H,t+1}, \tag{H.12}$$

$$Z^{U,2}_{H,t} = H^U_t p^U_{H,t} + \beta\theta \mathbb{E}_t \Lambda^U_{t,t+1} (\Pi^U_{H,t+1})^{\epsilon-1} Z^{U,2}_{H,t+1}, \tag{H.13}$$

$$Z^{U,1}_{X,t} = X^U_t \mathcal{MC}^U_t + \beta\theta \mathbb{E}_t \Lambda^U_{t,t+1} (\Delta\mathcal{E}_{t+1} \Pi^U_{X,t+1})^{\epsilon} Z^{U,1}_{H,t+1}, \tag{H.14}$$

$$Z^{U,2}_{X,t} = X^U_t Q_t p^U_{X,t} + \beta\theta \mathbb{E}_t \Lambda^U_{t,t+1} (\Delta\mathcal{E}_{t+1} \Pi^U_{X,t+1})^{\epsilon-1} Z^{U,2}_{X,t+1}, \tag{H.15}$$

$$Z_{H,t}^{R,1} = H_t^R \mathcal{MC}_t^R + \beta\theta\mathbb{E}_t\Lambda_{t,t+1}^R(\Pi_{H,t+1}^R)^{\epsilon} Z_{H,t+1}^{R,1}, \tag{H.45}$$
$$Z_{H,t}^{R,2} = H_t^R p_{H,t}^R + \beta\theta\mathbb{E}_t\Lambda_{t,t+1}^R(\Pi_{H,t+1}^R)^{\epsilon-1} Z_{H,t+1}^{R,2}, \tag{H.46}$$
$$Z_{X,t}^{R,1} = X_t^R \mathcal{MC}_t^R + \beta\theta\mathbb{E}_t\Lambda_{t,t+1}^R(\Pi_{X,t+1}^R)^{\epsilon} Z_{X,t+1}^{R,1}, \tag{H.47}$$
$$Z_{X,t}^{R,2} = X_t^R \mathcal{Q}_t^{-1} p_{X,t}^R + \beta\theta\mathbb{E}_t\Lambda_{t,t+1}^R(\Pi_{X,t+1}^R)^{\epsilon-1} Z_{X,t+1}^{R,2}. \tag{H.48}$$

Here, $\Pi_{H,t}^{U,+}$ is the optimal reset inflation rate for domestic good inflation, while $\Pi_{H,t}^U = P_{H,t}^U / P_{H,t-1}^U$ is the gross inflation rate for domestic goods in country U. The (optimal) inflation rates for the export markets are similarly defined. We also have to define the relative prices involved: $p_{H,t}^U = P_{H,t}^U / P_t^U$ is the relative price of domestic goods to the overall CPI, $p_{X,t}^U = P_{X,t}^U / P_t^R$ is the relative price of export goods produced in U and exported to R relative to the overall price level in country R (as export prices are in foreign currency). $\mathcal{Q}_t = \mathcal{E}_t P_t^R / P_t^U$ denotes the real exchange rate between country U and R. $\Delta\mathcal{E}_t$ is the change in the nominal exchange rate. The dynamics of the four inflation rates and the respective price dispersion measures are given by

$$1 = (1-\theta)(p_{H,t}^{U,+})^{1-\epsilon} + \theta(\Pi_{H,t}^U)^{\epsilon-1}, \tag{H.16}$$
$$1 = (1-\theta)(p_{X,t}^{+,U})^{1-\epsilon} + \theta(\Delta\mathcal{E}_t\Pi_{X,t}^U)^{\epsilon-1}, \tag{H.17}$$
$$\Delta_{H,t}^U = (1-\theta)(\Pi_{H,t}^{U,+})^{-\epsilon} + \theta(\Pi_{H,t}^U)^{\epsilon}\Delta_{H,t-1}^U, \tag{H.18}$$
$$\Delta_{H,t}^R = (1-\theta)(\Pi_{H,t}^{R,+})^{-\epsilon} + \theta(\Pi_{H,t}^R)^{\epsilon}\Delta_{H,t-1}^R, \tag{H.19}$$
$$1 = (1-\theta)(p_{H,t}^{R,+})^{1-\epsilon} + \theta(\Pi_{H,t}^R)^{\epsilon-1}, \tag{H.49}$$
$$1 = (1-\theta)(p_{X,t}^{R,+})^{1-\epsilon} + \theta(\Pi_{X,t}^R)^{\epsilon-1}, \tag{H.50}$$
$$\Delta_{X,t}^U = (1-\theta)(\Pi_{X,t}^{U,+})^{-\epsilon}(\Pi_{X,t}^U)^{\epsilon} + \theta(\Pi_{X,t}^U)^{\epsilon}\Delta_{X,t-1}^U, \tag{H.51}$$
$$\Delta_{X,t}^R = (1-\theta)(\Pi_{X,t}^{R,+})^{-\epsilon}(\Pi_{X,t}^R)^{\epsilon} + \theta(\Pi_{X,t}^R)^{\epsilon}\Delta_{X,t-1}^R. \tag{H.52}$$

It remains to discuss capital producers. We adopt the assumption by Gertler and Karadi (2011) on the structure of the firm sector to disentangle dynamic price and investment decisions: Producing firms buy capital at the beginning of the period, and re-sell it after using it. Capital producing firms buy the used capital, repair it, and build new capital. The new and refurbished capital is then sold again to final-goods producers at the price Q_t^U. The demand for capital by final-goods producers thus depends on the marginal product of capital and the variations in the price of capital

$$R_{k,t}^U = \frac{R_{z,t}^U + (1-\delta)Q_t^U}{Q_{t-1}^U}, \tag{H.21}$$
$$R_{k,t}^R = \frac{R_{z,t}^R + (1-\delta)Q_t^R}{Q_{t-1}^R}, \tag{H.54}$$

where $R_{k,t}^U$ is the return on capital and $R_{Z,t}^U$ is the rental rate of capital. We assume that there are flow adjustment costs associated with producing new capital. We use investment adjustment costs, which give rise to the investment Euler equation

$$1 = Q_t^U \left[1 - \frac{\kappa_I}{2} \left(\frac{I_t^U}{I_{t-1}^U} - 1 \right)^2 - \kappa_I \left(\frac{I_t^U}{I_{t-1}^U} - 1 \right) \frac{I_t^U}{I_{t-1}^U} \right] \tag{H.22}$$

$$+ \beta \kappa_I \mathbb{E}_t \left\{ \Lambda_{t,t+1}^U \left[Q_{t+1}^U \left(\frac{I_{t+1}^U}{I_t^U} - 1 \right) \frac{(I_{t+1}^U)^2}{(I_t^U)^2} \right] \right\},$$

$$1 = Q_t^R \left[1 - \frac{\kappa_I}{2} \left(\frac{I_t^R}{I_{t-1}^R} - 1 \right)^2 - \kappa_I \left(\frac{I_t^R}{I_{t-1}^R} - 1 \right) \frac{I_t^R}{I_{t-1}^R} \right] \tag{H.55}$$

$$+ \beta \kappa_I \mathbb{E}_t \left\{ \Lambda_{t,t+1}^R \left[Q_{t+1}^R \left(\frac{I_{t+1}^R}{I_t^R} - 1 \right) \frac{(I_{t+1}^R)^2}{(I_t^R)^2} \right] \right\},$$

where I_t denotes investment. Finally, the law of motion of capital is given by

$$K_t^U = (1-\delta) K_{t-1}^U + \left[1 - \frac{\kappa_I}{2} \left(\frac{I_t^U}{I_{t-1}^U} - 1 \right)^2 \right] I_t^U, \tag{H.23}$$

$$K_t^R = (1-\delta) K_{t-1}^R + \left[1 - \frac{\kappa_I}{2} \left(\frac{I_t^R}{I_{t-1}^R} - 1 \right)^2 \right] I_t^R, \tag{H.56}$$

where δ is the depreciation rate and κ_I is a parameter governing the adjustment costs.

Aggregate demand and price indices. Demand for domestic or imported goods depends on relative prices, including import prices, and overall demand in a country, which includes private consumption, private investment, and government spending. The aggregate demand aggregator is a composite demand index, defined as follows

$$Y_t^U(j) \equiv \left[(1-(1-n)\nu^U)^{\frac{1}{\eta}} (H_t^U(j))^{\frac{\eta-1}{\eta}} + ((1-n)\nu^U)^{\frac{1}{\eta}} (M_t^U(j))^{\frac{\eta-1}{\eta}} \right]^{\frac{\eta}{\eta-1}},$$

where $Y_t^U(j)$ is aggregate demand for the final good of variety j. $H_t^U(j)$ denotes domestically produced goods, while $M_t^U(j)$ are imported final goods of variety j. The imports of country U are exports from country R, and thus $M_t^U(j) = X_t^R(j)$. The parameter n denotes the country size, $\nu^U \in [0,1]$ is a measure of openness, and the parameter $\eta > 0$ measures the substitutability between domestic and imported goods. This is optimized given the constraint $P_t^U Y_t^U(j) = P_{H,t}^U H_t^U(j) + P_{M,t}^U M_t^U(j)$, where P_t^U denotes the overall price level, $P_{H,t}^U$ is the price of the domestically produced good, and $P_{M,t}^U$ is the price of the imported good, both expressed in domestic currency. The demand for domestic and imported goods in both countries and the decomposition of demand in private consumption, private investment, and government spending give the following equilibrium conditions:

$$Y_t^U = C_t^U + I_t^U + G_t^U, \tag{H.24}$$

$$\Delta_{H,t}^U H_t^U = (1-\nu^U)(p_{H,t}^U)^{-\eta} Y_t^U, \tag{H.25}$$

$$\Delta_{M,t}^U M_t^U = \nu^U (p_{M,t}^U)^{-\eta} Y_t^U, \tag{H.26}$$

$$Y_t^R = C_t^R + I_t^R + G_t^R, \tag{H.57}$$

$$\Delta_{H,t}^R H_t^R = (1-\nu^R)(p_{H,t}^R)^{-\eta} Y_t^R, \tag{H.58}$$

$$\Delta_{M,t}^R M_t^R = \nu^R (p_{M,t}^R)^{-\eta} Y_t^R, \tag{H.59}$$

where $\Delta_{H,t}^U$ accounts for price dispersion. The aggregate price indices and the link between the overall CPI and production prices in each country are given by

$$1 = (1-\nu^U)(p_{H,t}^U)^{1-\eta} + \nu^U (p_{M,t}^U)^{1-\eta}, \tag{H.28}$$

$$\Pi_{H,t}^U = \frac{p_{H,t}^U}{p_{H,t-1}^U} \Pi_t^U, \tag{H.29}$$

$$\Pi_{M,t}^U = \frac{p_{M,t}^U}{p_{M,t-1}^U} \Pi_t^U, \tag{H.30}$$

$$1 = (1-\nu^R)(p_{H,t}^R)^{1-\eta} + \nu^R (p_{M,t}^R)^{1-\eta}, \tag{H.61}$$

$$\Pi_{H,t}^R = \frac{p_{H,t}^R}{p_{H,t-1}^R} \Pi_t^R, \tag{H.62}$$

$$\Pi_{M,t}^R = \frac{p_{M,t}^R}{p_{M,t-1}^R} \Pi_t^R, \tag{H.63}$$

where $p_{H,t}^U = P_{H,t}^U / P_t^U$ and $p_{M,t}^U = P_{M,t}^U / P_t^U$ are the relative prices of domestic and import goods to the overall CPI level. Note that the relative import price of country U, $p_{M,t}^U$, is the same as the relative export price of country R, $p_{X,t}^R = P_{X,t}^R / P_t^U$, which is given in the currency of country U.

Financial sector. In country U, households directly hold capital and there is no financial sector. In country R, there is a financial sector subject to financial frictions. Households consist of workers and bankers. At any moment in time a fraction $1-f$ of the household members are workers and the fraction f are bankers. Each banker in the household operates a financial intermediary. Bankers exit randomly: any banker operating in period t continues into period $t+1$ with exogenous probability χ. When the banker exits, she rebates her earnings to the household and begins a career as a worker. At the same time, workers in the household become bankers with probability $(1-\chi)\frac{f}{1-f}$, so a measure of $(1-\chi)f$ of new bankers enters each period and exactly offsets the number that has exited. Entrant bankers receive a small equity endowment from the household so they can start operations.

Banker j chooses assets $S_t^R(j)$ (financial claims on non-financial firms), deposits issued to domestic households in the local currency $D_{r,t}^R(j)$, and deposits issued to country U households in foreign currency $D_{r,t}^U(j)$. The assets/financial claims to non-financial firms $S_t^R(j)$ consist of claims on country R's physical capital. The optimization problem facing banker j is to choose a state-contingent sequence $\{S_{t+k}^R(j), D_{r,t+k}^R(j), D_{r,t+k}^U(j)\}_{k=0}^{\infty}$ to maximize

$$V_t^R(j) = \mathbb{E}_t\left\{\beta\Lambda_{t,t+1}\left[(1-\chi)N_{t+1}^R(j) + \chi V_{t+1}^R(j)\right]\right\},$$

where $N_{t+k}^R(j)$ is terminal net worth if the banker exits at $t+k$. The banker's objective function is the expected value of its payout to the household, evaluated using the household's stochastic discount factor.

The maximization is subject to two constraints. The first constraint is the banker's budget constraint, given by

$$Q_t^R S_t^R(j) + R_{r,t}^R D_{r,t-1}^R(j) + R_{r,t}^U \frac{D_{r,t-1}^U(j)}{Q_t} \leq R_{k,t}^R Q_{t-1}^R S_{t-1}^R(j) + D_{r,t}^R(j) + \frac{D_{r,t}^U(j)}{Q_t},$$

where the left-hand side is banker i's uses of funds, consisting of loans to non-financial firms $(Q_t^R S_t^R(j))$ plus deposit repayments (both domestic, $R_{r,t}^R D_{t-1}^R(j)$, and foreign, $R_{r,t}^U Q_t^{-1} D_{r,t-1}^U(j)$). The right-hand side is the source of funds, consisting of returns from past loans to non-financial firms $(R_{k,t}^R Q_{t-1}^R S_t^R(j))$ plus deposits issued (to domestic residents, $D_{r,t}^R(j)$, and foreign households, $Q_t^{-1} D_{r,t}^U(j)$). The balance sheet identity for banker i is given by

$$Q_t^R S_t^R(j) = D_{r,t}^R(j) + \frac{D_{r,t}^U(j)}{Q_t} + N_t^R(j),$$

which states that the value of the banker's assets equals the value of its liabilities.

The second constraint arises due to moral hazard. After borrowing funds, the banker may decide to divert assets for personal gain, rather than honoring obligations with creditors. Diverting means selling a fraction of $\Theta_t \in (0,1)$ of assets secretly in secondary markets. The remaining assets are then seized by the banker's creditors in bankruptcy proceedings. We assume that Θ_t depends on the composition of the banker's liability portfolio

$$\Theta_t = \Theta\left(\frac{Q_t^{-1} D_{r,t}^U(j)}{Q_t^R S_t^R(j)}\right) = \Theta\left(x_t^R(j)\right) = \theta_F\left(1 + \frac{\gamma_F}{2}(x_t^R(j))^2\right),$$

where $\Theta(\cdot)$ is a function satisfying $\Theta' > 0$. We define the external share of liabilities as $x_t^R(j) = Q_t^{-1} D_{r,t}^U(j)/(Q_t^R S_t^R(j))$ and the functional form of Θ is a quadratic one. Thus, the banker is able to divert more assets when the fraction of her assets financed by *foreign* liabilities (equal to $Q_t^{-1} D_{r,t}^U(j)$) is larger. The assumption that $\Theta' > 0$ captures the notion that legal and institutional environments in country R, as well as the nature of the country R capital, which serves as collateral, effectively make it more difficult for foreign creditors to recover assets from a defaulting borrower compared to domestic creditors (Akinci and Queralto, 2024).

We define the leverage ratio as $\phi_t^R(j) = Q_t^R S_t^R(j)/N_t^R(j)$. Then, we guess the solution to the banker's problem to be linear in the banker's net worth $V_t^R(j) = \alpha_t N_t^R(j)$ and get

$$\begin{aligned}
\alpha_t^R = & \underbrace{\mathbb{E}_t\left[\beta\Lambda_{t,t+1}^R\left(1-\chi+\chi\alpha_{t+1}^R\right)\left(R_{k,t+1}^R - R_{r,t+1}^R\right)}_{=\mu_t^R}\phi_t^R(j)\right] \\
& + \underbrace{\mathbb{E}_t\left[\beta\Lambda_{t,t+1}^R\left(1-\chi+\chi\alpha_{t+1}^R\right)\left(R_{r,t+1}^R - R_{r,t+1}^U\frac{Q_t}{Q_{t+1}}\right)\right]}_{=\eta_t^R} x_t^R(j)\phi_t^R(j) \\
& + \underbrace{\mathbb{E}_t\left[\beta\Lambda_{t,t+1}^R\left(1-\chi+\chi\alpha_{t+1}^R\right)R_{r,t+1}^R\right]}_{=\xi_t^R}.
\end{aligned}$$

Hence, we introduce three coefficients, μ_t, η_t, and ξ_t. Those are given by:

$$\mu_t^R = \mathbb{E}_t\left[\beta\Lambda_{t,t+1}^R\left(1-\chi+\chi\alpha_{t+1}^R\right)\left(R_{k,t+1}^R - R_{r,t+1}^R\right)\right], \tag{H.66}$$

$$\eta_t^R = \mathbb{E}_t\left[\beta\Lambda_{t,t+1}^R\left(1-\chi+\chi\alpha_{t+1}^R\right)\left(R_{r,t+1}^R - R_{r,t+1}^U\frac{Q_t}{Q_{t+1}}\right)\right], \tag{H.67}$$

$$\xi_t^R = \mathbb{E}_t\left[\beta\Lambda_{t,t+1}^R\left(1-\chi+\chi\alpha_{t+1}^R\right)R_{r,t+1}^R\right]. \tag{H.68}$$

Their interpretation is as follows. μ_t^R is the excess marginal value of assets over deposits to the banker; η_t^R is the excess marginal cost of domestic relative to foreign funding (equivalently, the marginal value of foreign relative to domestic funding); and ξ_t^R is the marginal cost of domestic funding.

This allows us to set up the maximization problem as follows

$$\begin{aligned}
\alpha_t^R = & \max_{\phi_t^R(j), x_t^R(j)} \quad \left(\mu_t^R + \eta_t^R x_t^R(j)\right)\phi_t^R(j) + \xi_t^R \\
s.t. \quad & \left(\mu_t^R + \eta_t^R x_t^R(j)\right)\phi_t^R(j) + \xi_t^R \geq \Theta_t\left(x_t^R(j)\right)\phi_t^R(j)
\end{aligned}$$

with the following first-order conditions

$$\begin{aligned}
\mu_t + \eta_t x_t^R(j) &= \frac{\lambda_t^B}{1+\lambda_t^B}\Theta\left(x_t^R(j)\right), \\
\eta_t &= \frac{\lambda_t^B}{1+\lambda_t^B}\Theta'\left(x_t^R(j)\right), \\
\left(\mu_t + \eta_t x_t^R(j)\right)\phi_t^R(j) + \xi_t &= \Theta_t\left(x_t^R(j)\right)\phi_t^R(j).
\end{aligned}$$

Note that none of the marginal values is bank-specific, and thus we can drop the bank indicators. Hence, $x_t^R(j) = x_t^R$ and $\phi_t^R(j) = \phi_t^R$. We can combine the first two first-order conditions to get

$$\eta_t = \left(\frac{\Theta_t'(x_t^R)}{\Theta_t(x_t^R) - \Theta_t'(x_t^R)x_t^R}\right)\mu_t.$$

This condition equates the marginal value of foreign relative to domestic funding, η_t, to its marginal cost, which relates to the fact that a larger x_t^R tightens the incentive constraint. Plugging in the function, $\Theta(x_t^R) = \theta_F\left(1 + \frac{\gamma_F}{2}(x_t^R)^2\right)$, and its first derivative, $\Theta'(x_t^R) = \theta_F\gamma_F x_t^R$, yields

$$x_t^R = \frac{\mu_t}{\eta_t}\left(-1 + \sqrt{1 + \frac{2}{\gamma_F}\left(\frac{\eta_t}{\mu_t}\right)^2}\right) \tag{H.69}$$

Using the third first-order condition, we can express the leverage ratio as

$$\phi_t^R = \frac{\xi_t}{\Theta(x_t^R) - (\mu_t + \eta_t x_t^R)}, \tag{H.70}$$

which highlights that the leverage ratio, ϕ_t^R, is increasing in ξ_t, the saving to the bank in deposit costs from an extra unit of net worth, and in $\mu_t + \eta_t x_t^R$, the discounted total excess return on the bank's assets; and decreasing in the fraction of funds banks are able to divert, $\Theta(x_t^R)$. This also provides a solution for α_t^R using the optimization function:

$$\alpha_t^R = (\mu_t^R + \eta_t^R x_t^R)\phi_t^R + \xi_t^R. \tag{H.71}$$

Given that banks' leverage ratio ϕ_t^R and foreign funding ratio x_t^R do not depend on bank-specific factors, aggregating across banks yields the following relationships between the country R's aggregate assets and foreign debt ($S_t^R = \int_0^f S_t^R(j)dj$ and $D_{r,t}^U = \int_0^f D_{r,t}^U(j)dj$, respectively) and aggregate net worth $N_t^R = \int_0^f N_t^R(j)dj$:

$$Q_t^R S_t^R = \phi_t^R N_t^R, \tag{H.72}$$

$$Q_t^{-1} D_{r,t}^U = x_t^R \phi_t^R N_t^R. \tag{H.73}$$

If bank j is a new entrant, its net worth is given by $N_{n,t}^R = \iota Q_{t-1}^R S_{t-1}^R$. Using this condition, we can express the the evolution of banker's net worth as

$$\begin{aligned} N_t^R = \chi\Bigg[\left(R_{k,t}^R - R_{r,t}^R\right)Q_{t-1}^R S_{t-1}^R + \left(R_{r,t}^R - R_{r,t}^U\frac{Q_{t-1}}{Q_t}\right)Q_{t-1}^{-1}D_{r,t-1}^U + R_{r,t}^R N_{t-1}^R\Bigg] \\ + (1-\chi)\iota Q_{t-1}^R S_{t-1}^R. \end{aligned} \tag{H.74}$$

Lastly, we define the credit spread as

$$sp_t^R = \mathbb{E}_t\left[R_{k,t+1}^R - R_{r,t+1}^R\right]. \tag{H.75}$$

Monetary and fiscal policy. The central bank follows a Taylor rule in both countries

$$\left(\frac{R_t^U}{\overline{R}^U}\right) = \left[\frac{R_{t-1}^U}{\overline{R}^U}\right]^{\rho_r} \left[\left(\frac{\Pi_t^U}{\overline{\Pi}^U}\right)^{\phi_\pi} \left(\frac{Y_t^U}{\overline{Y}^U}\right)^{\phi_y}\right]^{1-\rho_r} \exp\{\varepsilon_{r,t}^U\}, \qquad \varepsilon_{r,t}^U \sim \mathcal{N}(0, \sigma_r^2) \tag{H.33}$$

$$\left(\frac{R_t^R}{\overline{R}^R}\right) = \left[\frac{R_{t-1}^R}{\overline{R}^R}\right]^{\rho_r} \left[\left(\frac{\Pi_t^R}{\overline{\Pi}^R}\right)^{\phi_\pi} \left(\frac{Y_t^R}{\overline{Y}^R}\right)^{\phi_y}\right]^{1-\rho_r}, \tag{H.76}$$

and the fiscal authority keeps a balanced budget

$$T_t^U = G^U, \tag{H.34}$$

$$T_t^R = G^R, \tag{H.77}$$

where government spending G^U is set exogenously.

Market clearing and international linkages. Both countries are linked through trade in the goods sector and the holding of foreign-currency debt in the financial sector. Market clearing in the goods market implies

$$P_t^U = \Delta_{H,t}^U H_t^U + \frac{1-n}{n}\Delta_{X,t}^U X_t^U, \tag{H.78}$$

$$P_t^R = \Delta_{H,t}^R H_t^R + \frac{n}{1-n}\Delta_{X,t}^R X_t^R, \tag{H.79}$$

where we adjust for relative population sizes. The real trade balance of country R reads

$$TB_t^R = -\left(\frac{D_{r,t}^U}{Q_t} - R_{r,t}^U \frac{D_{r,t-1}^U}{Q_t}\right), \tag{H.80}$$

$$TB_t^R = \frac{n}{1-n} Q_t^{-1} p_{X,t}^R \Delta_{X,t}^R X_t^R - p_{M,t}^R \Delta_{M,t}^R M_t^R, \tag{H.81}$$

which states that the net accumulation of foreign liabilities in country R (the left-hand side), equals the value of country R's exports minus the value of its imports (the right-hand side). Both sides are expressed in the foreign-currency of country R and adjusted by relative population sizes. Lastly, we can express the uncovered interest rate (UIP) premium and the link between the real and nominal exchange rates as

$$\Delta\mathcal{E}_t = \frac{Q_t}{Q_{t-1}} \frac{\Pi_t^U}{\Pi_t^R}, \tag{H.82}$$

$$UIP_t = \mathbb{E}_t\left[R_{r,t+1}^U - R_{r,t+1}^R \frac{Q_t}{Q_{t+1}}\right]. \tag{H.83}$$

Summary. In total, these are 83 equations for the behavior of 36 variables of country **U**: $\{Y_t^U, P_t^U, H_t^U, C_t^U, I_t^U, T_t^U, X_t^U, M_t^U, D_{r,t}^U, L_t^U, K_t^U, \varrho_t^U, \Lambda_{t,t+1}^U, \mathcal{MC}_t^U, Q_t^U, R_t^U, R_{r,t}^U, R_{z,t}^U, R_{k,t}^U, W_{r,t}^U, \Pi_t^U,$ $\Pi_{H,t}^U, \Pi_{X,t}^U, \Pi_{M,t}^U, p_{H,t}^U, p_{X,t}^U, p_{M,t}^U, p_{H,t}^{U,+}, p_{X,t}^{U,+}, Z_{H,t}^{U,1}, Z_{H,t}^{U,2}, Z_{X,t}^{U,1}, Z_{X,t}^{U,2}, \Delta_{H,t}^U, \Delta_{X,t}^U, \Delta_{M,t}^U\}$ and 44 variables of country **R**: $\{Y_t^R, P_t^R, H_t^R, C_t^R, I_t^R, T_t^R, X_t^R, M_t^R, TB_t^R, L_t^R, K_t^R, \varrho_t^R, \Lambda_{t,t+1}^R, \mathcal{MC}_t^R, Q_t^R, R_t^R, R_{r,t}^R, R_{z,t}^R,$

$R^R_{k,t}, W^R_{r,t}, \Pi^R_t, \Pi^R_{H,t}, \Pi^R_{X,t}, \Pi^R_{M,t}, p^R_{H,t}, p^R_{X,t}, p^R_{M,t}, p^{R,+}_{H,t}, p^{R,+}_{X,t}, Z^{R,1}_{H,t}, Z^{R,2}_{H,t}, Z^{R,1}_{X,t}, Z^{R,2}_{X,t}, \Delta^R_{H,t}, \Delta^R_{X,t}, \Delta^R_{M,t}, N^R_t,$ $\phi^R_t, x^R_t, sp^R_t, \alpha^R_t, \mu^R_t, \eta^R_t, \xi^R_t\}$ and three international prices $\Delta\mathcal{E}_t, Q_t, UIP_t$. The model has one exogenous variable $\varepsilon^U_{r,t}$ and the following set of parameters $\{n, \alpha, \beta, \delta, \sigma, \kappa_D, \kappa_I, \kappa_L, \varphi, \eta, \nu^U, \nu^R, \epsilon, \theta, \chi, \iota,$ $\theta_F, \gamma_F, \phi_\pi, \phi_y, \rho_r, G^U, G^R\}$. This concludes the description of the model.

H.2 Calibration

We report the baseline calibration in Table H1. We follow Akinci and Queralto (2024) and set the U.S. discount factor β^U to 0.9950, implying a steady state real interest rate of 2% per year. We set the RoW discount factor β^R to 0.9925, implying a steady state real interest rate of 3% per year. The preference and production parameters $\sigma, h, \varphi, \alpha, \delta, \epsilon, \theta, \kappa_I$ are set to conventional values. The Taylor rule parameters for the U.S. are set to conventional values, with $\phi^U_\pi = 1.5$ and $\phi^U_y = 0.125$. For the RoW countries, we assume that they only react to inflation with the conventional value of $\phi^R_\pi = 1.5$ and do not react to an opening up of the output gap. The parameter governing interest rate smoothing is set to 0.82 and we assume that the share of government spending in steady state is 20%.

Turning to the parameters related to trade, we set them in the baseline to the following values. We set the trade elasticity $\eta^U = \eta^R = 1.5$ in both countries to 1.5, which is commonly used (Erceg et al., 2007; Caldara et al., 2020; Akinci and Queralto, 2024). We set the country size of the U.S. to $n = 0.30$, as the U.S. share in world GDP fluctuates around this number in our sample period. The parameter ν, governing the degree of home bias in preferences, is usually calibrated to the import share. In the baseline, we set this parameter to 0.25, which is above the U.S. import share but clearly below values commonly used in the context of small open economies of $0.3 - 0.4$.

It remains to discuss the parameters related to the financial market frictions. We set the survival rate of bankers χ to 0.95, which is a value around the midpoint of the range found in the literature (Gertler and Karadi, 2011; Akinci and Queralto, 2024). The remaining three parameters are set to match three steady state targets: a leverage ratio of 4, a credit spread of 100 basis points (1 pp), and a ratio of foreign-currency debt to domestic debt of 30%. The target leverage ratio is a rough average across different sectors. We choose a credit spread of 100 basis points to have a mid-range between emerging markets with higher credit spreads and advanced economies. These targets imply that $\iota = 0.1583$, $\theta_F = 0.3007$, and $\gamma_F = 5.3793$. These values are similar in size to those in Akinci and Queralto (2024).

H.3 Spillovers of monetary policy shocks in the model

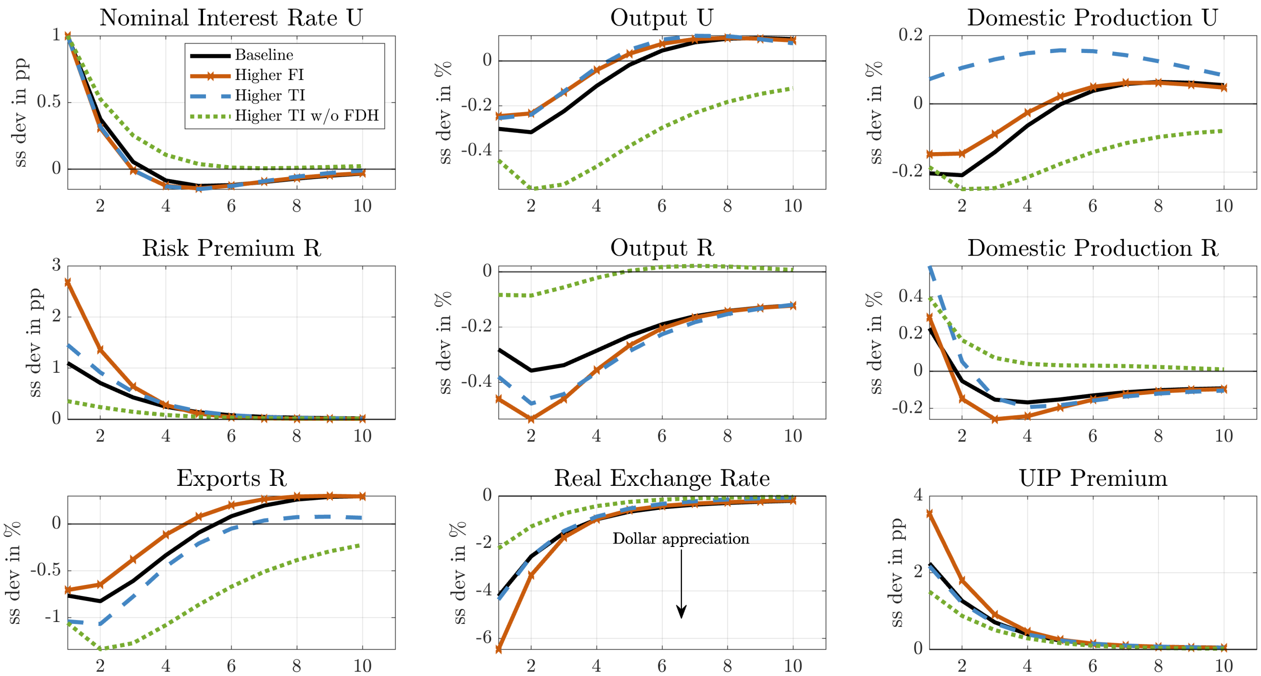
The theoretical analysis focuses on a contractionary monetary policy shock in country U. Similar to the empirical evidence, we are interested in the spillovers on RoW output and the transmission

Table H1: Parameter values.

Parameter	Value	Description	Source
Household preferences			
σ	2	Coefficient of relative risk aversion	Standard value
β^U	0.9950	Discount factor country U	Akinci and Queralto (2024)
β^R	0.9925	Discount factor country R	Akinci and Queralto (2024)
h	0.78	Habit parameter	Akinci and Queralto (2024)
φ	3.79	Frisch elasticity of labor supply	Akinci and Queralto (2024)
ν	0.25	Degree of home bias in preferences	
η	1.5	Elasticity of substitution between home and foreign goods	Erceg et al. (2007); Caldara et al. (2020); Akinci and Queralto (2024)
n	0.30	Size of country U	Matched to U.S. share on world GDP
Production			
α	0.33	Capital share in production	Standard value
δ	0.025	Depreciation rate	Standard value
ϵ	6	Elasticity of substitution between good varieties	Standard value
θ	0.75	Calvo parameter; probability of keeping prices fixed	Standard value
κ_I	2.48	Investment adjustment cost parameter	Standard value
Financial Intermediaries			
χ	0.95	Survival rate of banker	Akinci and Queralto (2024)
ι	0.1583	Transfer rate to new bank entrants	matched to leverage ratio of 4
θ_F	0.3007	Fraction of divertable capital	matched to credit spread of $1pp$ (per year)
γ_F	5.3793	Home bias in bank funding	matched to foreign-currency debt of 30%
Monetary and fiscal policy			
ϕ_π^U, ϕ_π^R	1.5	Response to inflation	Standard value
ϕ_y^U	0.125	Response to output in country U	Standard value
ϕ_y^R	0.0	Response to output in country R	Akinci and Queralto (2024)
ρ_r^U, ρ_r^R	0.82	Interest rate smoothing	Justiniano et al. (2010)
G/Y	0.2	Share of government spending in steady state	Standard value

via the trade and financial channel. In the model, we implement a monetary policy shock in country U of 1 percentage point (pp). We investigate the impulse responses using the calibration laid out in the previous section as a baseline. Then, we examine three experiments in the model: i) higher financial integration (FI), ii) higher trade integration (TI), and iii) higher trade integration but shutting down the additional amplification mechanism via foreign-currency debt holdings (FDH). We define higher trade integration as increasing the degree of home bias in preferences by 67% (ν^U: from 0.25 to 0.42). This increase corresponds in magnitude to the rise of the U.S. import share of GDP in the sample we consider. To keep it comparable, we implement a similar-sized increase for financial integration by increasing the leverage ratio ($\bar{\phi}$: from 4 to 6.67) and foreign-currency debt holdings (FDH: 30% to 50%). In the last exercise, we increase trade integration as before but set the foreign-currency debt holdings to a small value (FDH: 30% to 0.1%). To be clear, we

Figure H1: Model responses to a contractionary monetary policy shock in country U.



Notes: Model impulse responses to a contractionary monetary policy shock in country U. Baseline: Black line. Higher financial integration (FI): orange line with stars. Higher trade integration (TI): blue dashed line. Higher trade integration (TI) without foreign-currency debt holdings (FDH): dotted green line. Variables: nominal interest rate U (R_t^U), output U (Y_t^U), domestic production U (H_t^U), risk premium R (sp_t^R), output R (Y_t^R), domestic production R (H_t^R), exports R (X_t^R), real exchange rate (Q_t), and the UIP premium (UIP_t). Horizontal axis measures impulse response horizon in quarters. Vertical axis are in percent (output U, domestic production U, output R, domestic production R, exports R, and the real exchange rate) or in percentage points (nominal interest rate U, risk premium R, UIP premium).

translate higher trade and financial integration to a reduction of either trade or financial frictions in the model.

In Figure H1, we show the model responses of the nominal interest rate in U, output in U, domestic production in U, the risk premium in R, output in R, domestic production in R, exports in R, the real exchange rate, and the premium on the uncovered interest rate parity (UIP). We now discuss in detail the baseline and the three scenarios in the model. Each of the scenarios provides model predictions which we compare qualitatively to the empirical evidence.

In the baseline, a contractionary monetary policy shock in country U leads to an increase in the nominal interest rate, which we have normalized to a one percentage point increase. On the domestic side, this leads to a contraction of demand and decreasing inflation. As demand contracts, country U is importing less from abroad. Hence, exports in country R decrease. In a model without financial imperfections, country R would start switching to higher domestic good production. This is indeed what we observe for a couple of periods, before also domestic production contracts. This is due to the financial accelerator in the model, which works as follows. The increase in the nominal interest rate leads also to an appreciation of the dollar, which translates into a decline in the real exchange rate. This leads to a significant tightening of the balance sheet of banks in country R, driving up the risk premium, a sharp reduction in investment spending, and a contraction in output. The presence of strong feedback effects due to the financial market friction is key to understanding this reaction in the model. First, we observe the *standard financial accelerator*: a decline in intermediary net worth weakens investment spending and Tobin's Q, which depresses net worth further. Here, the model adds another feedback effect through the exchange rate and foreign-currency borrowing. A lowering of banks' net worth is associated with an opening up of the UIP premium and a stronger reaction of the exchange rate. Akinci and Queralto (2024) show that a negative shift in intermediaries' net worth is linked to an increase in the risk premium, a drop in asset prices, a wider UIP premium, and a stronger depreciation of the local currency *vis-á-vis* the dollar.

We translate the process of globalization into a reduction (increase) in either trade or financial frictions (integration) in the model. Is the model now able to predict the stronger spillovers we observe in the data? In terms of output, the answer is *yes* for both channels. Let us assume the reduction in financial frictions (i.e., an increase in financial integration) is responsible for this effect. As agents take on more foreign-currency debt, the initial decline in net worth is stronger, which leads to a higher risk premium and a wider UIP premium. The foreign-currency depreciates even stronger. This additional feedback loop creates a stronger drop in demand, causing stronger output spillovers. Notably, on top of the stronger output spillovers, we observe that exports in R actually decrease less than in the baseline scenario. This is due to an adjustment of prices. In response to weakened demand in country R (and the dollar appreciation that makes markups on exports temporarily too high in country R), firms reduce their domestic and export prices, boosting their exports and, at the same time, households demand less imports, opening up the trade balance. However, this prediction of the model – namely, that exports in R decline less – is difficult to reconcile with the empirical evidence, which points to substantially stronger spillovers on both output *and trade* (Figure 3 and Figure 4). Moreover, higher financial integration in the model would also imply a stronger response of the risk premium and a more pronounced exchange rate appreciation—this is not what we observe in the data, as shown in Figure 4 and Figure E2.

The model's predictions are instead more in line with the empirical evidence when reducing trade frictions (i.e., an increase in trade integration). The decline of demand in country U now translates to a stronger decline in demand for foreign goods. Exports in country R thus contract much stronger. The strong decline in trade leads to an increase in domestic goods production in country U. The strong effects via a reduction in imports are also shown in Camara et al. (2024). This opens up the trade balance and creates stronger output spillovers in country R, while the exchange rate and the risk premium are not strongly affected in comparison to the baseline (and in line with our empirical evidence).

However, we note that these strong effects via an increase in trade integration are dependent on a certain level of financial integration. To examine this more carefully in the model, we consider a third scenario in which we use the *high trade integration* calibration, but shut down foreign-currency debt holdings (loosely speaking, no financial integration). The appreciation of the dollar does not lead to valuation effects on banks' balance sheets. The financial accelerator is not (strongly) set in motion and demand contracts only slightly in country R – essentially, the model would predict no output spillovers (while trade is reduced markedly, country R can substitute this with domestic goods production). Hence, some degree of financial integration is *necessary* for the model to generate strong spillovers. In the absence of financial integration, the role of trade integration is significantly diminished (in this sense, there is a substantial interaction between trade and financial integration, which somewhat echoes the results in Wu et al., 2024). This once again highlights the overall importance of the financial channel in the global transmission of U.S. monetary policy shocks—which we do not dispute. Our point is not that financial transmission is not important in general, but rather that, when it comes to explaining the *time variation in spillovers*, the trade channel appears to play a more relevant role.

To conclude, the results from the model are consistent with the empirical facts we have documented. Specifically, the model predicts that an increase in trade integration — given a certain level of financial integration — leads to: i) larger output spillovers, ii) stronger effects on RoW exports, and iii) only mildly stronger financial repercussions. Therefore, an increase in trade integration is able to jointly account for the three stylized facts observed in the data, whereas an increase in financial integration cannot.

H.4 Full set of equilibrium conditions

We provide an overview of all equations in the model. In total, the model consists of 83 equations.

Equations for country U:

$$\varrho_t^U = (C_t^U - hC_{t-1}^U)^{-\sigma} - \beta h\left(\mathbb{E}_t C_{t+1}^U - hC_t^U\right)^{-\sigma} \tag{H.1}$$

$$\Lambda_{t,t+1}^U = \frac{\varrho_{t+1}^U}{\varrho_t^U} \tag{H.2}$$

$$W_{r,t}^U = \kappa_L (L_t^U)^{\varphi} (C_t^U)^{\sigma} \tag{H.3}$$

$$1 = \beta R_{r,t+1}^U \mathbb{E}_t\left[\Lambda_{t,t+1}\right] \tag{H.4}$$

$$1 = \beta \mathbb{E}_t\left[\Lambda_{t,t+1} R_{k,t+1}^U\right] \tag{H.5}$$

$$R^U_{r,t+1} = \frac{R^U_{t+1}}{\mathbb{E}_t \Pi^U_{t+1}} \tag{H.6}$$

$$P^U_t = (K^U_{t-1})^\alpha (L^U_t)^{1-\alpha} \tag{H.7}$$

$$W^U_{r,t} = (1-\alpha)\mathcal{MC}^U_t \frac{P^U_t}{L^U_t} \tag{H.8}$$

$$R^U_{z,t} = \alpha \mathcal{MC}^U_t \frac{P^U_t}{K^U_{t-1}} \tag{H.9}$$

$$p^{U,+}_{H,t} = \frac{\epsilon}{\epsilon - 1} \frac{Z^{U,1}_{H,t}}{Z^{U,2}_{H,t}} \tag{H.10}$$

$$p^{U,+}_{X,t} = \frac{\epsilon}{\epsilon - 1} \frac{Z^{U,1}_{X,t}}{Z^{U,2}_{X,t}} \tag{H.11}$$

$$Z^{U,1}_{H,t} = H^U_t \mathcal{MC}^U_t + \beta\theta \mathbb{E}_t \Lambda^U_{t,t+1} (\Pi^U_{H,t+1})^\epsilon Z^{U,1}_{H,t+1} \tag{H.12}$$

$$Z^{U,2}_{H,t} = H^U_t p^U_{H,t} + \beta\theta \mathbb{E}_t \Lambda^U_{t,t+1} (\Pi^U_{H,t+1})^{\epsilon-1} Z^{U,2}_{H,t+1} \tag{H.13}$$

$$Z^{U,1}_{X,t} = X^U_t \mathcal{MC}^U_t + \beta\theta \mathbb{E}_t \Lambda^U_{t,t+1} (\Delta\mathcal{E}_{t+1} \Pi^U_{X,t+1})^\epsilon Z^{U,1}_{H,t+1} \tag{H.14}$$

$$Z^{U,2}_{X,t} = X^U_t \mathcal{Q}_t p^U_{X,t} + \beta\theta \mathbb{E}_t \Lambda^U_{t,t+1} (\Delta\mathcal{E}_{t+1} \Pi^U_{X,t+1})^{\epsilon-1} Z^{U,2}_{X,t+1} \tag{H.15}$$

$$1 = (1-\theta)(p^{U,+}_{H,t})^{1-\epsilon} + \theta(\Pi^U_{H,t})^{\epsilon-1} \tag{H.16}$$

$$1 = (1-\theta)(p^{+,U}_{X,t})^{1-\epsilon} + \theta(\Delta\mathcal{E}_t \Pi^U_{X,t})^{\epsilon-1} \tag{H.17}$$

$$\Delta^U_{H,t} = (1-\theta)(p^{U,+}_{H,t})^{-\epsilon} + \theta(\Pi^U_{H,t})\Delta^U_{H,t-1} \tag{H.18}$$

$$\Delta^U_{X,t} = (1-\theta)(p^{U,+}_{X,t})^{-\epsilon} + \theta(\Delta\mathcal{E}_t \Pi^U_{X,t})^\epsilon \Delta^U_{X,t-1} \tag{H.19}$$

$$\Delta^U_{M,t} = \Delta^R_{X,t} \tag{H.20}$$

$$R^U_{k,t} = \frac{R^U_{z,t} + (1-\delta)Q^U_t}{Q^U_{t-1}} \tag{H.21}$$

$$\begin{aligned} 1 = {} & Q^U_t \left[1 - \frac{\kappa_I}{2} \left(\frac{I^U_t}{I^U_{t-1}} - 1 \right)^2 - \kappa_I \left(\frac{I^U_t}{I^U_{t-1}} - 1 \right) \frac{I^U_t}{I^U_{t-1}} \right] \\ & + \beta\kappa_I \mathbb{E}_t \left\{ \Lambda^U_{t,t+1} \left[Q^U_{t+1} \left(\frac{I^U_{t+1}}{I^U_t} - 1 \right) \frac{(I^U_{t+1})^2}{(I^U_t)^2} \right] \right\} \end{aligned} \tag{H.22}$$

$$K^U_t = (1-\delta)K^U_{t-1} + \left[1 - \frac{\kappa_I}{2} \left(\frac{I^U_t}{I^U_{t-1}} - 1 \right)^2 \right] I^U_t \tag{H.23}$$

$$Y^U_t = C^U_t + I^U_t + G^U_t \tag{H.24}$$

$$\Delta^U_{H,t} H^U_t = (1-\nu^U)(p^U_{H,t})^{-\eta} Y^U_t \tag{H.25}$$

$$\Delta^U_{M,t} M^U_t = \nu^U (p^U_{M,t})^{-\eta} Y^U_t \tag{H.26}$$

$$X_t^U = M_t^R \tag{H.27}$$

$$1 = (1-\nu^U)(p_{H,t}^U)^{1-\eta} + \nu^U (p_{M,t}^U)^{1-\eta} \tag{H.28}$$

$$\Pi_{H,t}^U = \frac{p_{H,t}^U}{p_{H,t-1}^U}\Pi_t^U \tag{H.29}$$

$$\Pi_{M,t}^U = \frac{p_{M,t}^U}{p_{M,t-1}^U}\Pi_t^U \tag{H.30}$$

$$p_{M,t}^U = p_{X,t}^R \tag{H.31}$$

$$\Pi_{M,t}^U = \Pi_{X,t}^R \tag{H.32}$$

$$\left(\frac{R_t^U}{\overline{R}^U}\right) = \left[\frac{R_{t-1}^U}{\overline{R}^U}\right]^{\rho_r} \left[\left(\frac{\Pi_t^U}{\overline{\Pi}^U}\right)^{\phi_\pi} \left(\frac{Y_t^U}{\overline{Y}^U}\right)^{\phi_y}\right]^{1-\rho_r} \exp\{\varepsilon_{r,t}^U\}, \qquad \varepsilon_{r,t}^U \sim \mathcal{N}(0,\sigma_r^2), \tag{H.33}$$

$$T_t^U = G^U \tag{H.34}$$

Equations for country R:

$$\varrho_t^R = (C_t^R - hC_{t-1}^R)^{-\sigma} - \beta h\left(\mathbb{E}_t C_{t+1}^R - hC_t^R\right)^{-\sigma} \tag{H.35}$$

$$\Lambda_{t,t+1}^R = \frac{\varrho_{t+1}^R}{\varrho_t^R} \tag{H.36}$$

$$W_{r,t}^R = \kappa_L (L_t^R)^{\varphi} (C_t^R)^{\sigma} \tag{H.37}$$

$$1 = \beta R_{r,t+1}^R \mathbb{E}_t\left[\Lambda_{t,t+1}^R\right] \tag{H.38}$$

$$R_{r,t+1}^R = \frac{R_{t+1}^R}{\mathbb{E}_t \Pi_{t+1}^R} \tag{H.39}$$

$$P_t^R = (K_{t-1}^R)^{\alpha} (L_t^R)^{1-\alpha} \tag{H.40}$$

$$R_{z,t}^R = \alpha \mathcal{M}C_t^R \frac{P_t^R}{K_{t-1}^R} \tag{H.41}$$

$$W_{r,t}^R = (1-\alpha)\mathcal{M}C_t^R \frac{P_t^R}{L_t^R} \tag{H.42}$$

$$p_{H,t}^{R,+} = \frac{\epsilon}{\epsilon-1}\frac{Z_{H,t}^{R,1}}{Z_{H,t}^{R,2}} \tag{H.43}$$

$$p_{X,t}^{R,+} = \frac{\epsilon}{\epsilon-1}\frac{Z_{X,t}^{R,1}}{Z_{X,t}^{R,2}} \tag{H.44}$$

$$Z_{H,t}^{R,1} = H_t^R \mathcal{M}C_t^R + \beta\theta \mathbb{E}_t \Lambda_{t,t+1}^R (\Pi_{H,t+1}^R)^{\epsilon} Z_{H,t+1}^{R,1} \tag{H.45}$$

$$Z_{H,t}^{R,2} = H_t^R p_{H,t}^R + \beta\theta \mathbb{E}_t \Lambda_{t,t+1}^R (\Pi_{H,t+1}^R)^{\epsilon-1} Z_{H,t+1}^{R,2} \tag{H.46}$$

$$Z_{X,t}^{R,1} = X_t^R \mathcal{M}C_t^R + \beta\theta\mathbb{E}_t\Lambda_{t,t+1}^R(\Pi_{X,t+1}^R)^{\epsilon} Z_{X,t+1}^{R,1} \tag{H.47}$$

$$Z_{X,t}^{R,2} = X_t^R Q_t^{-1} p_{X,t}^R + \beta\theta\mathbb{E}_t\Lambda_{t,t+1}^R(\Pi_{X,t+1}^R)^{\epsilon-1} Z_{X,t+1}^{R,2} \tag{H.48}$$

$$1 = (1-\theta)(p_{H,t}^{R,+})^{1-\epsilon} + \theta(\Pi_{H,t}^R)^{\epsilon-1} \tag{H.49}$$

$$1 = (1-\theta)(p_{X,t}^{R,+})^{1-\epsilon} + \theta(\Pi_{X,t}^R)^{\epsilon-1} \tag{H.50}$$

$$\Delta_{H,t}^R = (1-\theta)(p_{H,t}^{R,+})^{-\epsilon} + \theta(\Pi_{H,t}^R)^{\epsilon}\Delta_{H,t-1}^R \tag{H.51}$$

$$\Delta_{X,t}^R = (1-\theta)(p_{X,t}^{R,+})^{-\epsilon} + \theta(\Pi_{X,t}^R)^{\epsilon}\Delta_{X,t-1}^R \tag{H.52}$$

$$\Delta_{M,t}^R = \Delta_{X,t}^U \tag{H.53}$$

$$R_{k,t}^R = \frac{R_{z,t}^R + (1-\delta)Q_t^R}{Q_{t-1}^R} \tag{H.54}$$

$$1 = Q_t^R \left[1 - \frac{\kappa_I}{2}\left(\frac{I_t^R}{I_{t-1}^R} - 1\right)^2 - \kappa_I\left(\frac{I_t^R}{I_{t-1}^R} - 1\right)\frac{I_t^R}{I_{t-1}^R}\right] \tag{H.55}$$

$$K_t^R = (1-\delta)K_{t-1}^R + \left[1 - \frac{\kappa_I}{2}\left(\frac{I_t^R}{I_{t-1}^R} - 1\right)^2\right] I_t^R \tag{H.56}$$

$$Y_t^R = C_t^R + I_t^R + G_t^R \tag{H.57}$$

$$\Delta_{H,t}^R H_t^R = (1-\nu^R)(p_{H,t}^R)^{-\eta} Y_t^R \tag{H.58}$$

$$\Delta_{M,t}^R M_t^R = \nu^R (p_{M,t}^R)^{-\eta} Y_t^R \tag{H.59}$$

$$X_t^R = M_t^U \tag{H.60}$$

$$1 = (1-\nu^R)(p_{H,t}^R)^{1-\eta} + \nu^R(p_{M,t}^R)^{1-\eta} \tag{H.61}$$

$$\Pi_{H,t}^R = \frac{p_{H,t}^R}{p_{H,t-1}^R}\Pi_t^R \tag{H.62}$$

$$\Pi_{M,t}^R = \frac{p_{M,t}^R}{p_{M,t-1}^R}\Pi_t^R \tag{H.63}$$

$$p_{M,t}^R = p_{X,t}^U \tag{H.64}$$

$$\Pi_{M,t}^R = \Pi_{X,t}^U \tag{H.65}$$

$$\mu_t^R = \mathbb{E}_t\left[\beta\Lambda_{t,t+1}^R\left(1-\chi+\chi\alpha_{t+1}^R\right)\left(R_{k,t+1}^R - R_{r,t+1}^R\right)\right. \tag{H.66}$$

$$\eta_t^R = \mathbb{E}_t\left[\beta\Lambda_{t,t+1}^R\left(1-\chi+\chi\alpha_{t+1}^R\right)\left(R_{r,t+1}^R - R_{r,t+1}^U\frac{Q_t}{Q_{t+1}}\right)\right] \tag{H.67}$$

$$\xi_t^R = \mathbb{E}_t\left[\beta\Lambda_{t,t+1}^R\left(1-\chi+\chi\alpha_{t+1}^R\right)R_{r,t+1}^R\right] \tag{H.68}$$

$$x_t^R = \frac{\mu_t^R}{\eta_t^R}\left(-1 + \sqrt{1 + \frac{2}{\gamma_F}\left(\frac{\eta_t^R}{\mu_t^R}\right)^2}\right) \tag{H.69}$$

$$\phi_t^R = \frac{\xi_t^R}{\theta_F(1+\frac{\gamma_F}{2}(x_t^R)^2) - (\mu_t^R + \eta_t^R x_t^R)} \tag{H.70}$$

$$\alpha_t^R = (\mu_t^R + \eta_t^R x_t^R)\phi_t^R + \xi_t^R \tag{H.71}$$

$$Q_t^R K_t^R = \phi_t^R N_t^R \tag{H.72}$$

$$Q_t^{-1} D_{F,r,t}^U = x_t^R \phi_t^R N_t^R \tag{H.73}$$

$$N_t^R = \chi\left[\left(R_{k,t}^R - R_{r,t}^R\right)Q_{t-1}^R K_{t-1}^R + \left(R_{r,t}^R - R_{r,t}^U \frac{Q_{t-1}}{Q_t}\right)Q_{t-1}^{-1} D_{F,r,t-1}^U + R_{r,t}^R N_{t-1}^R\right] \tag{H.74}$$

$$+ (1-\chi)\iota Q_{t-1}^R K_{t-1}^R$$

$$sp_t^R = \mathbb{E}_t\left[R_{k,t+1}^R - R_{r,t+1}^R\right] \tag{H.75}$$

$$+ \beta\kappa_I \mathbb{E}_t\left\{\Lambda_{t,t+1}^R\left[Q_{t+1}^R\left(\frac{I_{t+1}^R}{I_t^R} - 1\right)\frac{(I_{t+1}^R)^2}{(I_t^R)^2}\right]\right\}$$

$$\left(\frac{R_t^R}{\overline{R}^R}\right) = \left[\frac{R_{t-1}^R}{\overline{R}^R}\right]^{\rho_r}\left[\left(\frac{\Pi_t^R}{\overline{\Pi}^R}\right)^{\phi_\pi}\left(\frac{Y_t^U}{\overline{Y}^R}\right)^{\phi_y}\right]^{1-\rho_r} \tag{H.76}$$

$$T_t^R = G^R \tag{H.77}$$

Market clearing and definitions:

$$P_t^U = \Delta_{H,t}^U H_t^U + \frac{1-n}{n}\Delta_{X,t}^U X_t^U \tag{H.78}$$

$$P_t^R = \Delta_{H,t}^R H_t^R + \frac{n}{1-n}\Delta_{X,t}^R X_t^R \tag{H.79}$$

$$TB_t^R = -\left[\frac{D_{r,t}^U}{Q_t} - R_{r,t}^U \frac{D_{r,t-1}^U}{Q_t}\right] \tag{H.80}$$

$$TB_t^R = \frac{n}{1-n} Q_t^{-1} p_{X,t}^R \Delta_{X,t}^R X_t^R - p_{M,t}^R \Delta_{M,t}^R M_t^R \tag{H.81}$$

$$\Delta\mathcal{E}_t = \frac{Q_t}{Q_{t-1}}\frac{\Pi_t^U}{\Pi_t^R} \tag{H.82}$$

$$UIP_t = \mathbb{E}_t\left[R_{r,t+1}^U - R_{r,t+1}^R \frac{Q_t}{Q_{t+1}}\right] \tag{H.83}$$

Appendix References

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