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# 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. Based on a cross-country analysis and counterfactual simulations, we argue that such amplification can primarily be attributed to the surge in trade integration, while the role of rising financial integration in explaining the time-variation is limited.

**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. 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-SVAR 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".

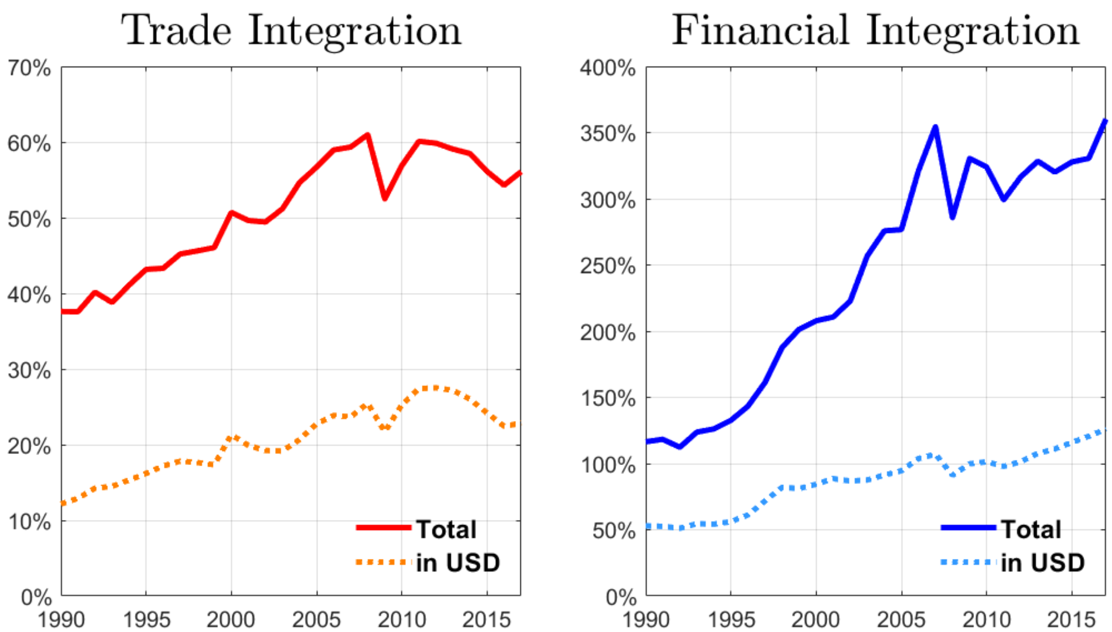
We document that globalization has led to substantial time-variation in the international ramifications of policy shocks: U.S. policy hikes generate stronger recessionary effects over time, with a flattening out of effects after the Great Recession. Whereas there is vast evidence showing that a monetary hike engineered by the Federal Reserve (Fed) generates global recessionary effects (e.g., 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; and Bräuning and Sheremirov, 2023), our results on the relevance of the time-variation is novel.

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 the global (excluding 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 are transmitted globally through trade and financial channels. The former refers to the traditional Mundell-Fleming effects, in which the current account adjusts because monetary policy affects both aggregate demand (of foreign goods) and the currency. A domestic policy tightening generates recessionary effects at home, which depresses demand for imports (rest of the world - RoW - export). This demand effect should be partly counteracted

[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).

**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 the External Wealth of Nations database (Lane and Milesi-Ferretti (2018)) and Bénétrix et al. (2019)).

through 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 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. Still, the U.S. dollar is not only the dominant invoicing currency for U.S. trade but for global trade in general (Goldberg and Tille, 2008; Gopinath, 2015; Gopinath et al., 2020; 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 negative output spillovers in response to the appreciation of the dollar (Gopinath and Neiman, 2014; Gopinath et al., 2020; Georgiadis and Schumann, 2021; Cook and Patel, 2023).[2] Monetary policy also operates globally through a financial channel by affecting asset prices, which either relaxes or binds balance sheet or leverage constraints and thus shapes investors' risk aversion (see, inter alia, Farhi and Werning, 2014; Bruno and Shin, 2015; Rey, 2016). This, in turn,

[2] 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 *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. Cook and Patel (2023) argue that global value chains lead to asymmetric adjustments in trade of the dominant-currency economy compared to regional economies.

has significant effects on cross-border capital flows, funding costs of agents, and eventually feeds back into asset prices. This causes a strong case for an international risk channel of monetary policy (Miranda-Agrippino and Rey, 2020). Particularly, the existence of the global financial cycle (see Rey, 2015, 2016) acts as a transmitter for the financial channel. 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., 2021, 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 spillover sizes only stabilized 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 (pp) hike in monetary policy leads to about a $-0.6\%$ contraction in RoW industrial production in 1993, a same-sized shock in 2008 results in a downturn of about $-3.2\%$. When we look at the transmission via the trade and financial channel, we find that monetary policy shocks generate global trade contractions and financial frictions. However, we find a disconnect between the two channels 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. Finally, rising spillovers in real activity are somehow mirrored by a strengthening of domestic effects on U.S. prices and output. Greater international spillovers imply a greater potential for these effects to spillback to the U.S. economy. Based on previous estimates by Breitenlechner et al. (2022), we argue that spillover-spillback loop effects are the likely driver of the rising effects of U.S. policy shocks on the U.S. economy that we find in the data. This calls for incorporating spillback effects in the calibration of U.S. policy decisions today more than in the past.

Given the disconnect in the evolution of effects on global trade (highly time-dependent) and financial conditions (relatively constant over time), our estimations suggest that trade integration is the primary factor driving 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 have 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 22 countries and estimate the 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 such association emerges in relation to heightened financial integration. We also find evidence that rising financial integration is not responsible for time-variation in real spillovers by running counterfactual simulations. To do so, we first identify a *global financial shock* following the approach of Gilchrist and Zakrajšek (2012), and then simulate counterfactual scenarios in which U.S. monetary disturbances do not impact global financial conditions (Sims and Zha, 2006; Antolín-Díaz et al., 2021; McKay and Wolf, 2023). By shutting off the financial channel in the transmission of U.S. monetary policy shocks, we obtain two results. On the one hand, the financial channel is important *on average*, accounting for about 30 – 40% of the overall response of RoW industrial production. When it comes to explaining the time-variation in the RoW industrial production response, on the other hand, its role is again found to be limited.

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.[3] We are closely related to the contributions of Miranda-Agrippino and Rey (2020), Dedola et al. (2017), Iacoviello and Navarro (2019), and Degasperi et al. (2020). While the first paper extensively investigates the impact of U.S. monetary policy on the global financial cycle, the others focus on examining the heterogeneity of spillovers in real economic activity. Differently, we use a time-varying model and document the increasing international spillovers of U.S. policy shocks.[4]

[3] Canova (2005), Maćkowiak (2007), Georgiadis (2016), Feldkircher and Huber (2016), Dedola et al. (2017), Iacoviello and Navarro (2019), Degasperi et al. (2020), Georgiadis and Schumann (2021), and Ca'Zorzi et al. (2023) analyze the foreign responses of real activity to U.S. policy decisions. See Rey (2016), Gerko and Rey (2017), Jordà et al. (2019), Habib and Venditti (2019), Dées and Galesi (2021), and Miranda-Agrippino and Rey (2020) for the transmission to international financial conditions. Obstfeld (2020) provides a comprehensive overview on the global dimension of U.S. monetary policy.

[4] The time-varying nature of monetary policy shocks on *domestic outcomes* has been documented in many studies, see for instance Cogley and Sargent (2005), Primiceri (2005), Boivin and Giannoni (2006), Canova and Gambetti (2009), Galí and Gambetti (2015) and Aastveit et al. (2023).

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

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 empirically 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 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).[5] 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.[6] 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 the growth rate and keep the remaining variables in levels. Our monthly dataset covers the time span from 1980M1 to 2017M12, which we split in two parts.

[5] 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.

[6] 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.

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.[7] 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.[8] 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.

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; Jarociński and Karadi, 2020).[9] Our choice is guided by the fact that such instrument has been shown to generate stable and well identified responses in small-scale models featuring a limited number of lags such as ours (see Miranda-Agrippino and Ricco, 2023).[10]

### *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

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

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

[7] Coibion (2012) shows that the identification of monetary policy shocks is sensitive to the inclusion of specific observations in the early 1980s. We have verified that our results are unchanged when the training sample only starts in 1984.

[8] 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 Section 3.6 we show that the our results are similar when excluding the zero lower bound period from the estimation sample.

[9] 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.

[10] This approach is also robust to a possible misspecification highlighted by Caldara and Herbst (2019), who argue that monetary policy VARs need to include credit spreads indices. Miranda-Agrippino and Ricco (2023) show that the exclusion of credit spreads does not generate any instability in the responses when using the instrument of Miranda-Agrippino and Ricco (2021), which we use here. 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).

policy instrument) with its respective $n \times 1$ time-varying coefficients $\boldsymbol{B}_t$. By 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) make it difficult to estimate heavily parameterized models, and stochastic volatility adds another significant estimation burden (see Paul (2020) for the same choice). 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 Equation 2.1. The dynamics 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 (Equation 2.1) and the state equation (Equation 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).[11] 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 10,000 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*

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 in $t$. We are interested in the effects of U.S. monetary policy and

[11] See Noh (2019) for a formal justification of this procedure. As a robustness check, we have also used an uninformative prior on $\boldsymbol{B}_t$. Results are robust to this choice and available upon request.

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 $\varepsilon_{1t}$ 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 the shock via

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

where $\zeta_t$ is orthogonal to all other variables. This assumption implies that the relation between the instrument and the monetary policy shock is not time-varying.[12]

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 equivalent to the external instrument approach (Stock and Watson, 2012 and Mertens and Ravn, 2013) under mild conditions.[13] The posterior quantities of $\boldsymbol{A}_{jt}$ ($j = 1, \ldots, p$; $t = 1, \ldots, T$) are then used to trace out subsequent impulse responses. We normalize the impact effect of a contractionary monetary policy shock on 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 Equation 2.3, see Paul, 2020).

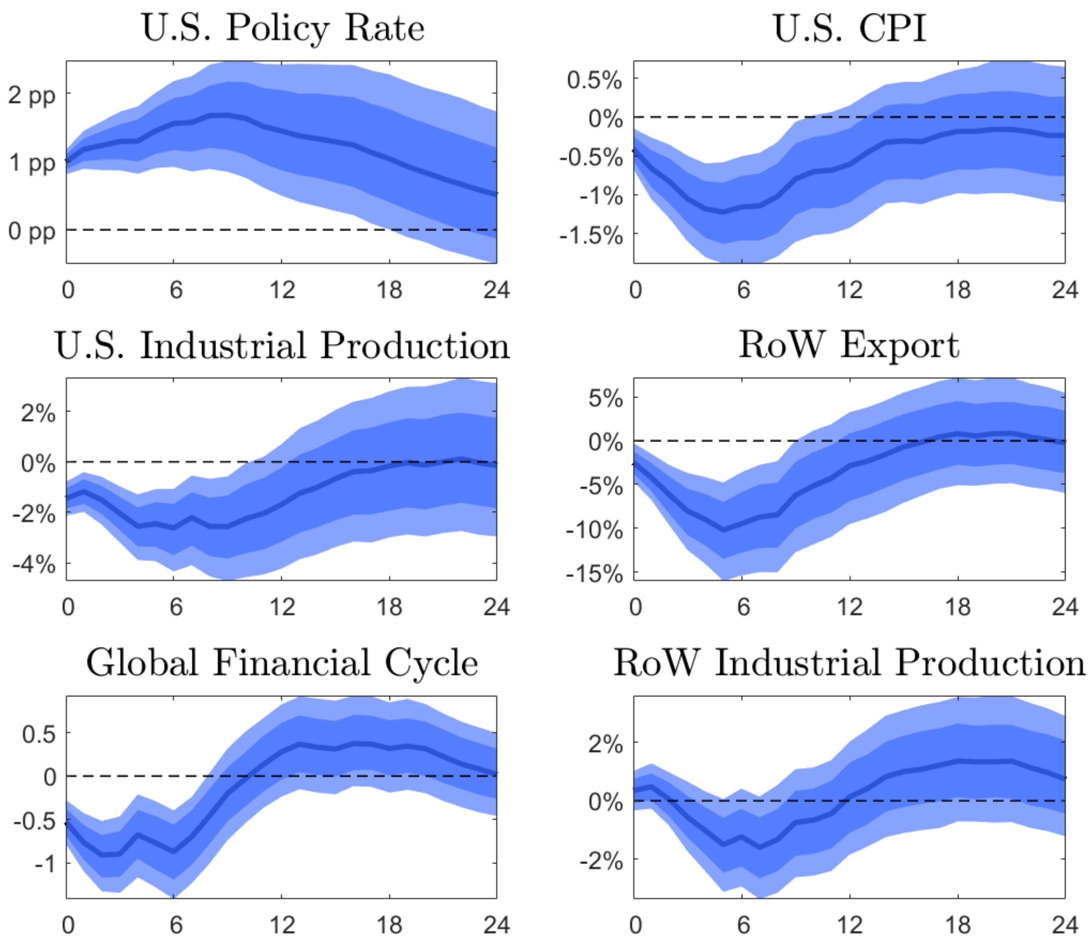
## 3. Evidence

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

[12] 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).

[13] 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.

**Figure 2:** Linear Impulse Responses Functions.




*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).

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.

### *3.1 Evidence from the Linear Specification*

We first estimate a linear VAR considering the variables described before.[14] We test the relevance of the policy instrument in the linear VAR and we find a first-stage F-statistic of 26.6, 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 (1 pp, i.e., 100 basis point) contractionary U.S. monetary policy shock. A monetary policy tightening is followed by strong

[14] 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).

domestic recessionary effects, consistent with standard macroeconomic theory: Domestic real activity deteriorates and prices 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. However, within the context of this initial linear framework, the evidence of a worldwide economic downturn due to U.S. monetary policy tightening is relatively mild. The adverse effects resulting from monetary shocks only attain statistical significance at a confidence level of 68%.

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.

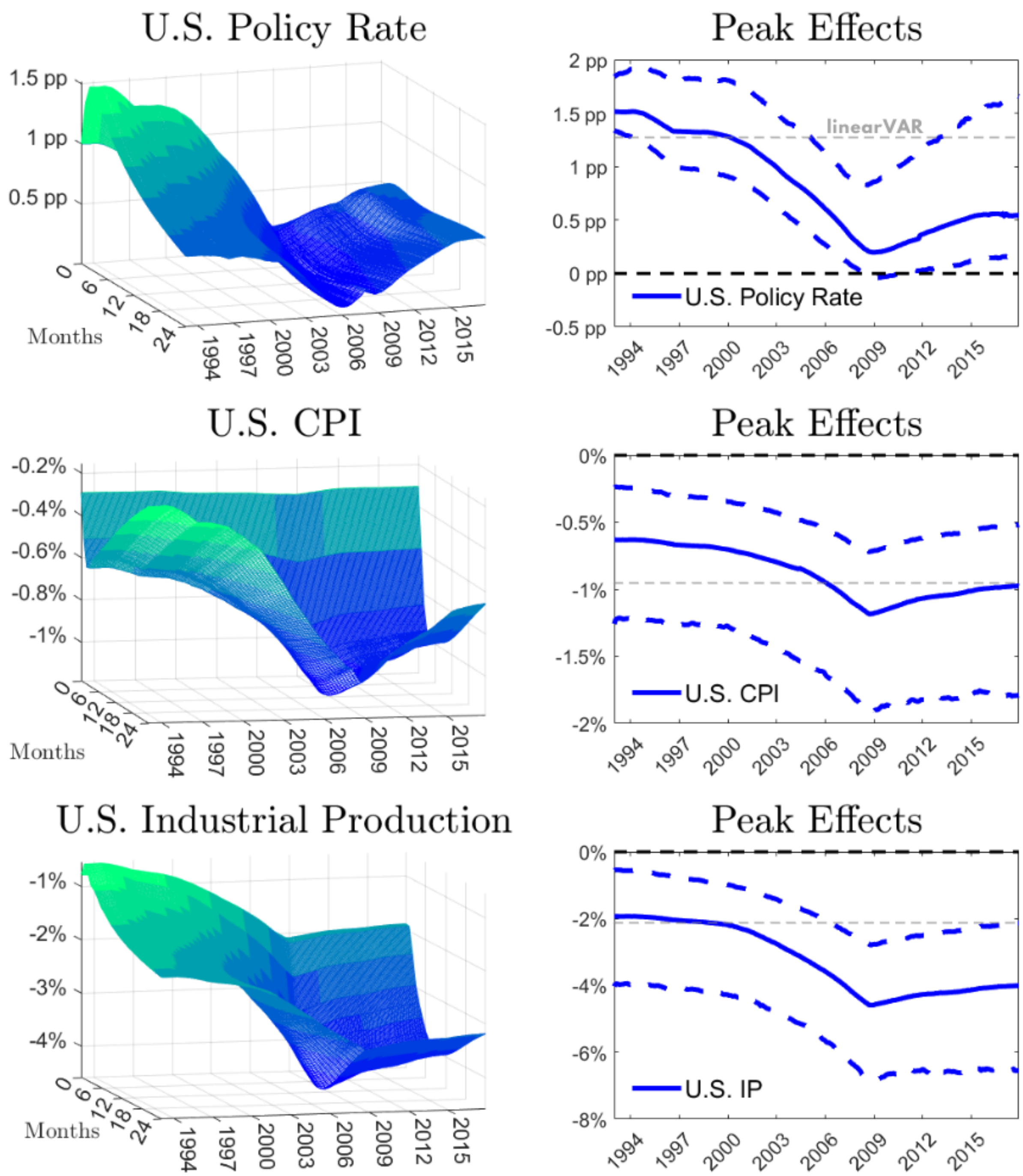
### *3.2 Evidence from the Time-Varying Parameter Specification*

We report the time-varying impulse response functions from the TVP-VAR in Figure 3 (U.S. domestic variables) and Figure 4 (global variables). 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 U.S. consumer prices, U.S. industrial production, RoW exports, and RoW industrial production. The global financial factor has no unit of scaling attached.

To facilitate interpretation, we show the results in two ways. First, we report the evolution over time of the median impulse-response functions in the left column. Second, we display the peak response of the variables (together with its 68% credibility sets) in the right column. For comparison, we also report the (constant) peak responses for equally sized shocks in the linear VAR (horizontal dotted grey lines). Since we only identify relative impulse responses, we normalize the response of the U.S. policy rate to a one percentage point increase on impact at 1993M1 (first month of the estimation sample).[15]

[15] Consistently, we use the same variation in $z_t$ underlying the TVP-VAR shocks to normalize the linear VAR's responses, which delivers a 0.76 percentage point increase in the policy rate within the linear specification.

**Figure 3:** Time-Varying Impulse Responses Functions of Domestic Variables.




*Notes*: Responses of domestic variables to a contractionary U.S. monetary policy shock that induces a one percentage point (1 pp, 100 basis points) increase in the federal fund rate in 1993M1. Left column reports the evolution over time of the median impulse-response functions. The right column reports the peak effects over time for each variable with 68% posterior credible sets. Grey dotted horizontal lines: (constant) peak effects in the linear VAR of a same-sized shock (which consists of a 0.76 pp increase in the policy rate in the linear specification).

Figure 3 confirms the findings obtained in the linear setting, with a monetary policy tightening being followed by conventional negative demand-type effects in the U.S. economy. A domestic contraction in industrial production goes along with a decline in prices. The magnitudes of these effects are consistent with the linear specification. Some comments, however, are in order. First, to obtain same-size shocks, the impact effect of the shocks on the policy rate itself diminishes over time. This pattern is consistent with a progressive decline of the long-run trend of the U.S. interest rate, which reaches its trough with the zero lower bound period (2009-2015).[16] Second, time-dependent patterns arise in the responses of U.S. aggregates. Peak contractions in domestic industrial production aggravate over time (rising from $-1.9\%$ in 1993 to $-4.4\%$ in 2008) and only stabilize at their lowest level with the onset of the Great Recession (this evolution is in line with the findings of Paul (2020), who consider a different specification). A similar pattern emerges for U.S. prices, which contract stronger over time. The impact in the initial periods of the sample is about $-0.6\%$, which grows in magnitude and reaches $-1.1\%$ during the Great Recession.

Figure 4 presents the dynamic responses of RoW exports, RoW industrial production, and the global financial cycle. 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 global trade, production, and the financial cycle. Throughout the period considered, all variables react negatively (and most of the time statistically significant so) to a monetary policy shock. The peak effect on RoW real activity strongly increases over time, increasing from $-0.6\%$ in 1993 to $-3.2\%$ in 2008.[17] In this regard, the linear VAR seems to capture well the mean effect over time, masking though the time-specific heterogeneity. Similarly, the response of RoW exports (as a measure of trade) is strongly growing in magnitude (from $-3.6\%$ to $-11.4\%$).[18] The time-variation in the impulse response of the global financial cycle is more limited but yet non-negligible from an economic point of view, with an increase over time from $-0.63$ to $-0.90$.

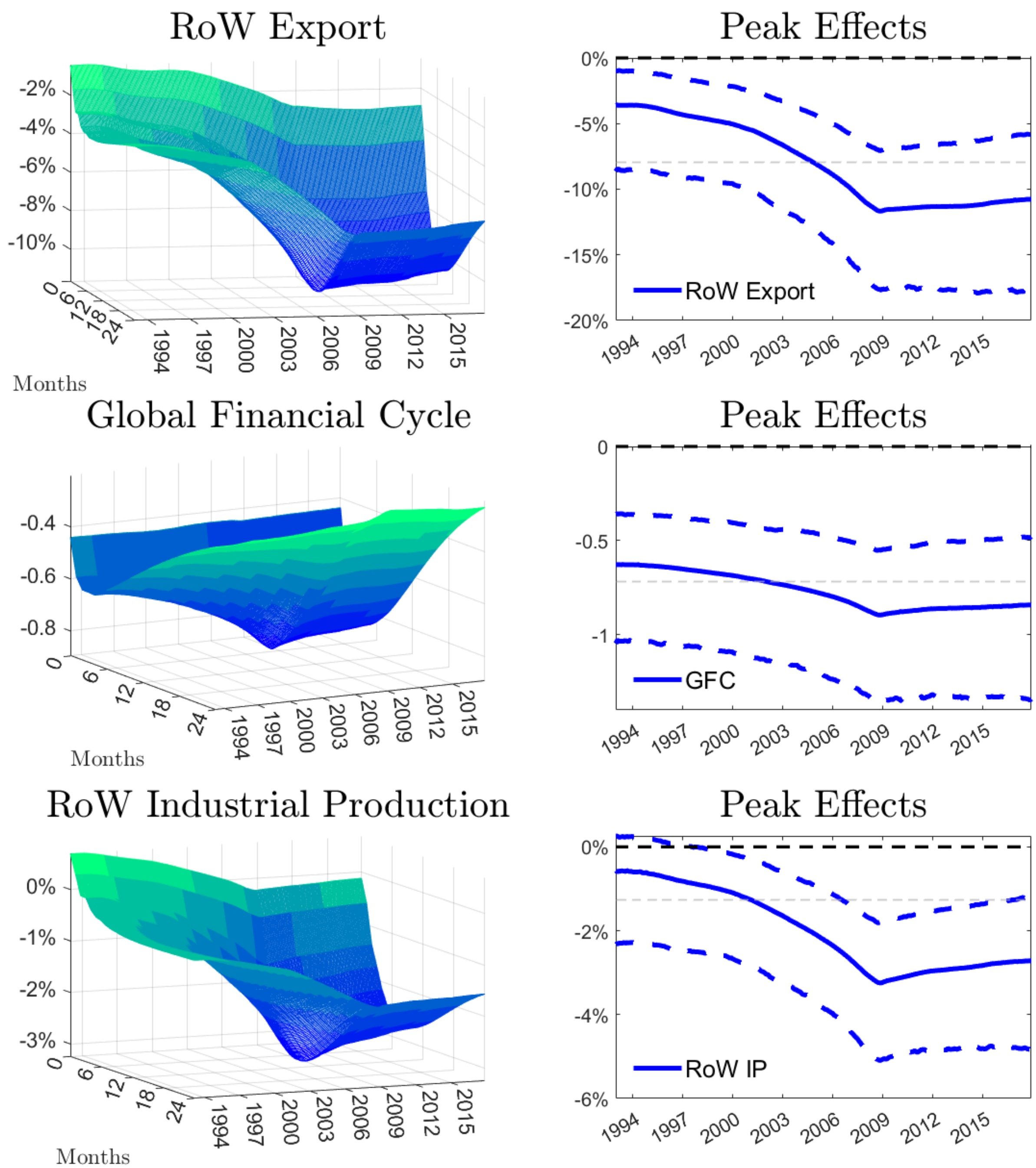
Although the magnitude of the global effects has been growing since the beginning of the sample, the pace of this growth notably accelerates in the early 2000s. This intensification of international spillovers coincides precisely with a period characterized by factors such as the trade boom, relaxed financial regulation and supervision of banks (Shin, 2012), and a sharp rise in

[16] Figure C1 reports the evolution of the long-run trend along with the on impact response of the U.S. policy rate underlying our specification. Accordingly, the long-run trend is relatively stable in 1990s, but starts decreasing in the 2000s, and becomes even negative during the zero lower bound period. See Appendix C for more details.

[17] Given that the shocks are normalized to have the same magnitude, the effects over time are comparable. However, the shape of the policy rate's responses different. To eliminate any concern, we report in Figure D1 the ratios of the peak responses for RoW vs. U.S. industrial production (see Hofmann and Peersman (2017) for a similar use of ratio impulse response functions). In each $t$, the shock hitting the two variables is the same. We obtain an increase in the ratio response over time, which again points towards significant time-variation in global spillovers.

[18] In unreported checks available upon request, we find extremely similar results when considering RoW imports. This is consistent with a symmetric contraction of RoW export and import in response to U.S. policy shocks (Gopinath et al., 2020; Degasperi et al., 2020).

**Figure 4:** Time-Varying Impulse Responses Functions of Global Variables.




*Notes*: Responses of domestic variables to a contractionary U.S. monetary policy shock that induces a one percentage point (1 pp, 100 basis points) increase in the federal fund rate in 1993M1. Left column reports the evolution over time of the median impulse-response functions. The right column reports the peak effects over time for each variable with 68% posterior credible sets. Grey dotted horizontal lines: (constant) peak effects in the linear VAR of a same-sized shock (which consists of a 0.76 pp increase in the policy rate in the linear specification).

dollar-denominated cross-border positions held by international actors (Rey, 2016). All these factors indicate a heightened role of international linkages in transmitting U.S. monetary policy decisions abroad, which is evident in the data. 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, 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, which could potentially impact our estimates. Our empirical model appears to effectively capture this economic narrative.

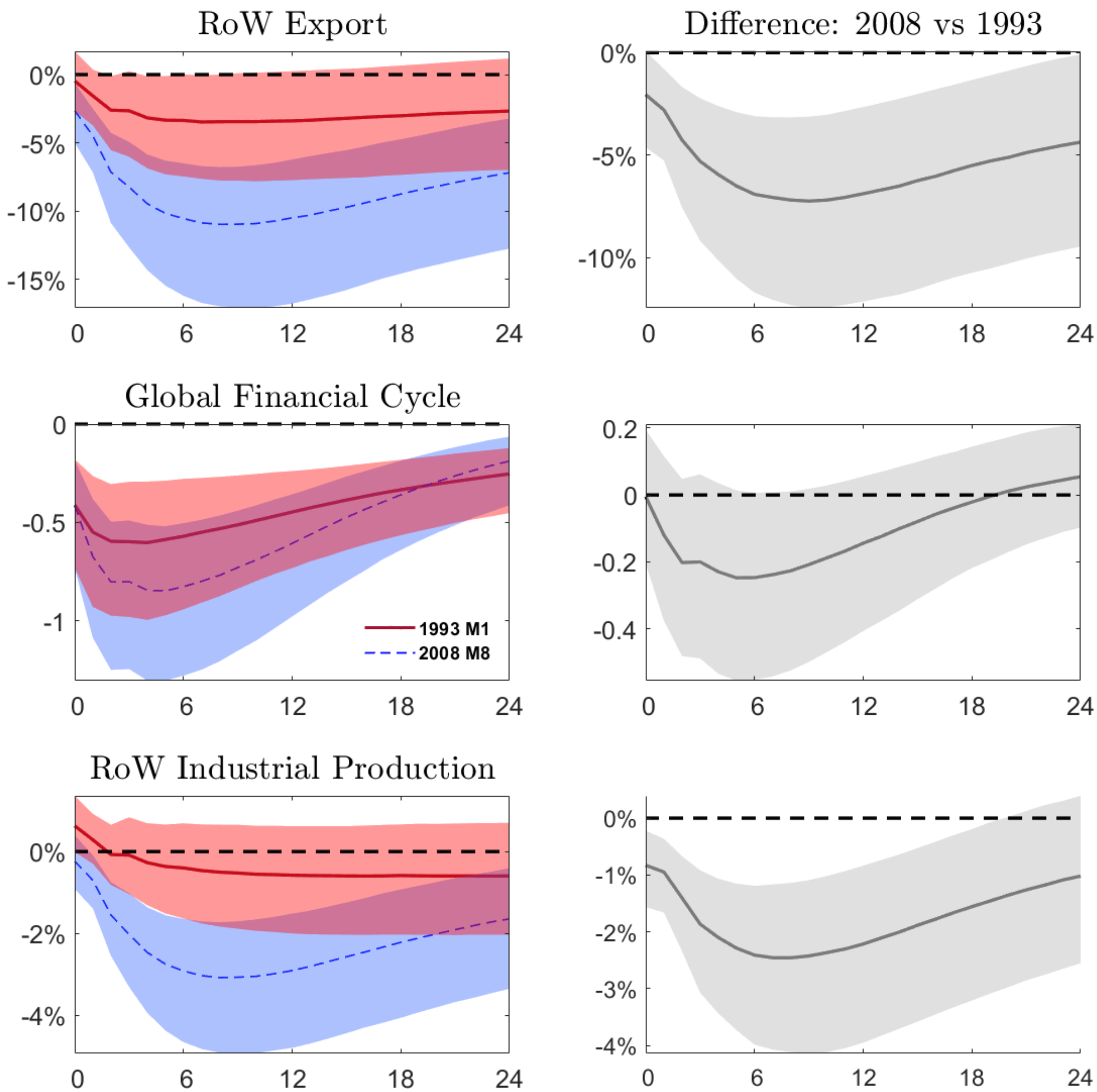
The stronger global effects of U.S. monetary policy shocks can also trigger spillover-spillback loop effects. This means that international recessionary effects spillback to the domestic economy and affect its economic aggregates. Evidence for an active spillback mechanism is provided by Breitenlechner et al. (2022), who find that spillbacks account for nearly half of the overall effect of U.S. monetary policy on domestic real activity using counterfactual simulations. In our estimations, such spillover-spillback loops are the likely explanation of the time-varying effects that we find in the response of U.S. domestic variables.

Finally, we investigate whether time-variation is statistically significant by focusing on particular episodes in the sample. 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, which is somehow in line with the results in the linear model. 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 sharp contrast, the evidence for the global financial cycle is weak: While U.S. disturbances generate significant financial frictions in both periods, the difference in the impulse-response functions is not statistically different from zero.

### *3.3 Wider Propagation Channels*

To get a better understanding of the time-varying transmission of U.S. monetary policy shocks, we analyze the effects on a range of relevant macroeconomic and financial variables. 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 (we refer to Appendix A for the exact variable definitions and transformations).

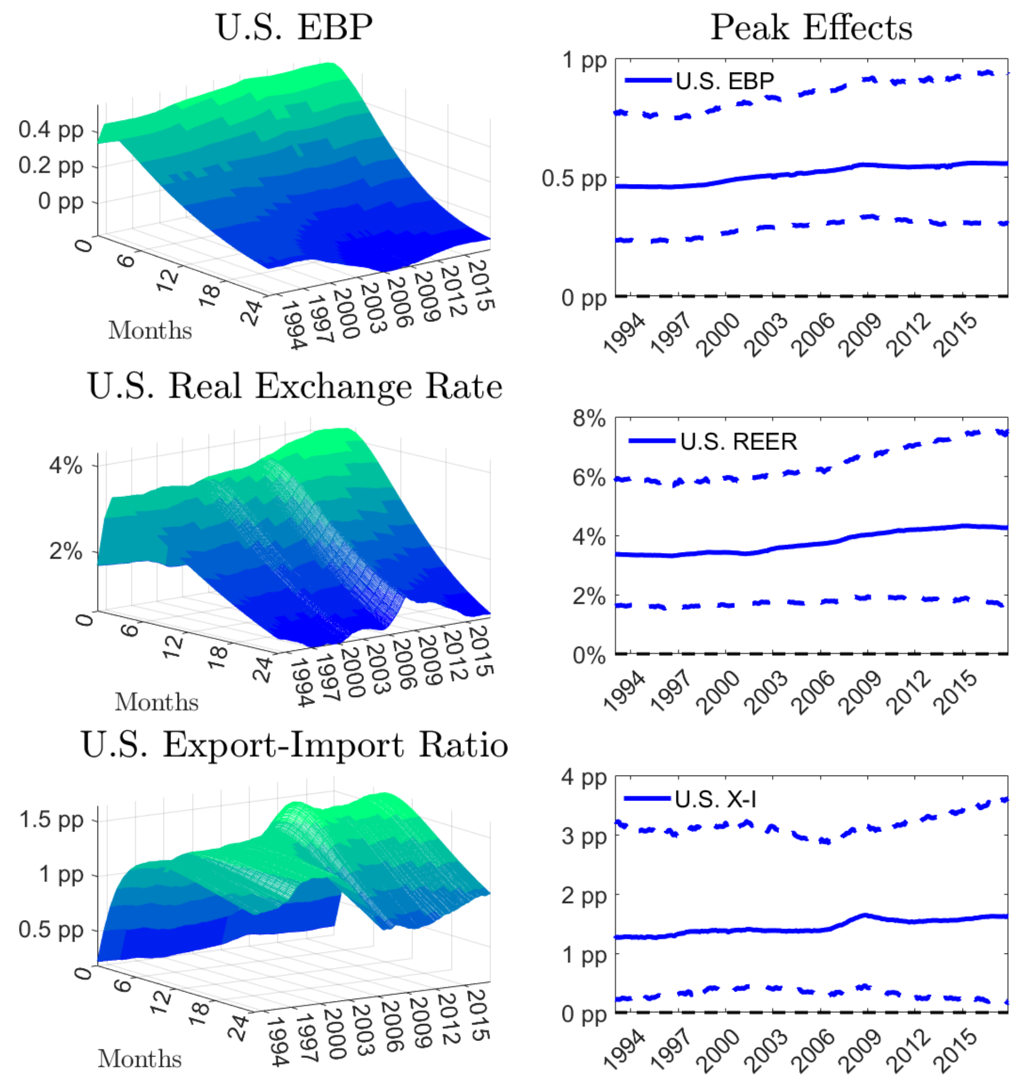
**Figure 5:** Differences in Impulse Responses: 1993M1 vs. 2008M8




*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).

We report the results of these extensions in Figure 6 and Figure 7. 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; and iii) a rise in the U.S. export import ratio. The sign of these effects are as expected. First, Caldara and Herbst (2019) highlight the role of financial conditions in

**Figure 6:** Impulse Responses of Extensions to the Baseline Specification.

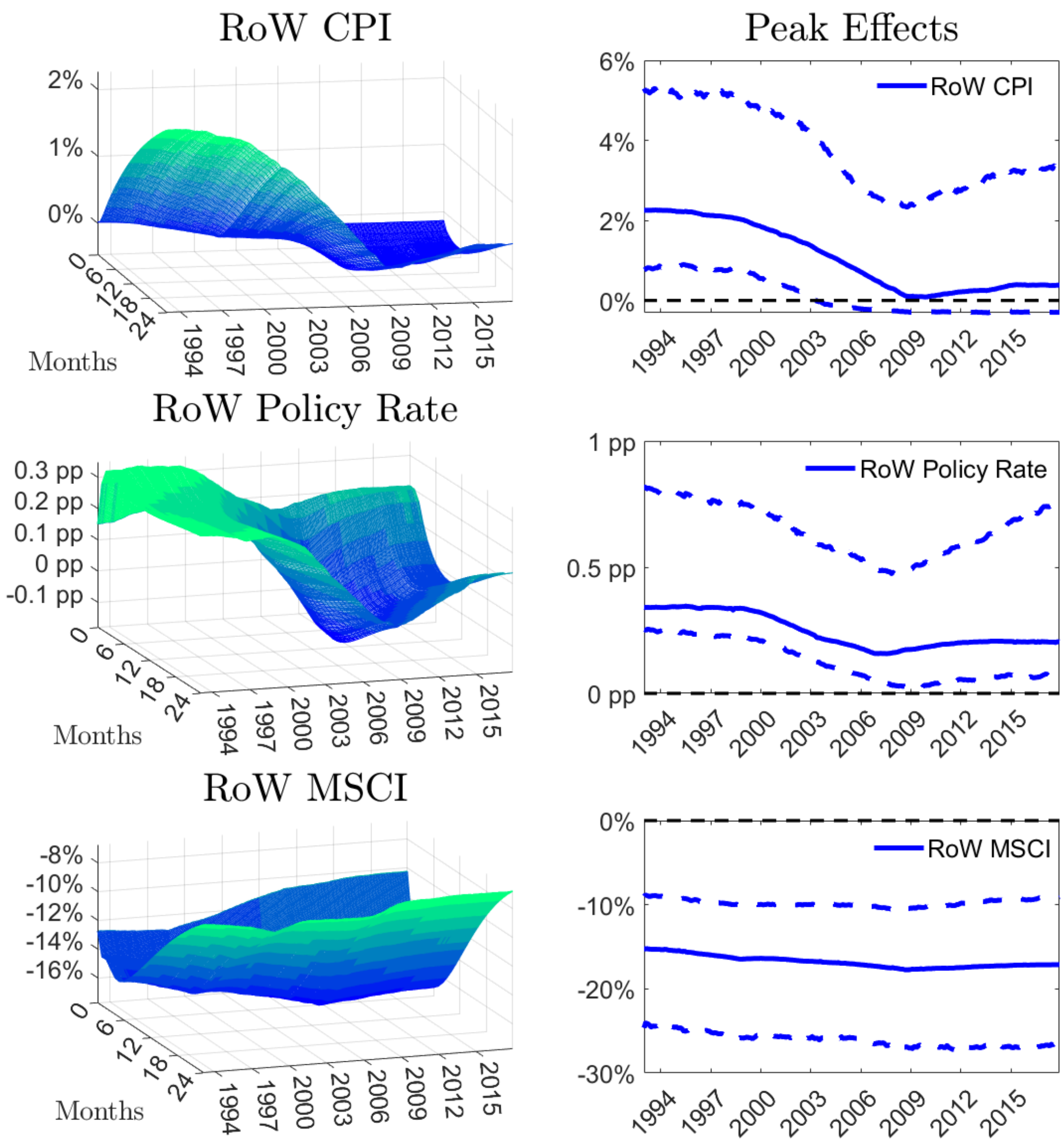



*Notes*: Responses of additional variables to a contractionary U.S. monetary policy shock that induces a one percentage point (1 pp, 100 basis points) increase in the federal fund rate in 1993M1. Left column reports the evolution over time of the median impulse-response functions. The right column reports the peak effects over time for each variable with 68% posterior credible sets.

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.[19] Second, the appreciation of the dollar is expected by the uncovered interest rate parity. Third, the positive response of the U.S. export import ratio implies no discernible expenditure-switching channel, as expected in the dominant currency paradigm (Gopinath et al.,

[19] As mentioned earlier, and in contrast to Caldara and Herbst (2019), our specification remains robust to the inclusion of credit spreads due to the choice of the instrument (Miranda-Agrippino and Ricco, 2021; Miranda-Agrippino and Ricco, 2023).

**Figure 7:** Impulse Responses of Extensions to the Baseline Specification.




*Notes*: Responses of additional variables to a contractionary U.S. monetary policy shock that induces a one percentage point (1 pp, 100 basis points) increase in the federal fund rate in 1993M1. Left column reports the evolution over time of the median impulse-response functions. The right column reports the peak effects over time for each variable with 68% posterior credible sets.

2020). While U.S. exports decline in response to less aggregate demand, U.S. imports do not outweigh this force. 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 for similar results). We uncover that these responses, while being statistically significant throughout the whole sample, exhibit very mild evidence of time-variation.

We now turn to the global responses in Figure 7. 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). 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.[20] 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. Finally, we look at the responses of the global stock market, measured through the RoW MSCI index. The indicator shows an abrupt decline in response to U.S. disturbances, with peak responses being stable over time (about −15%). These findings are in line with the response of the global financial factor of Miranda-Agrippino and Rey (2020) in our benchmark specification: Again, we do not find evidence of time-variation in the effects of U.S. shocks on global financial markets.

### *3.4 Heterogeneity Analysis*

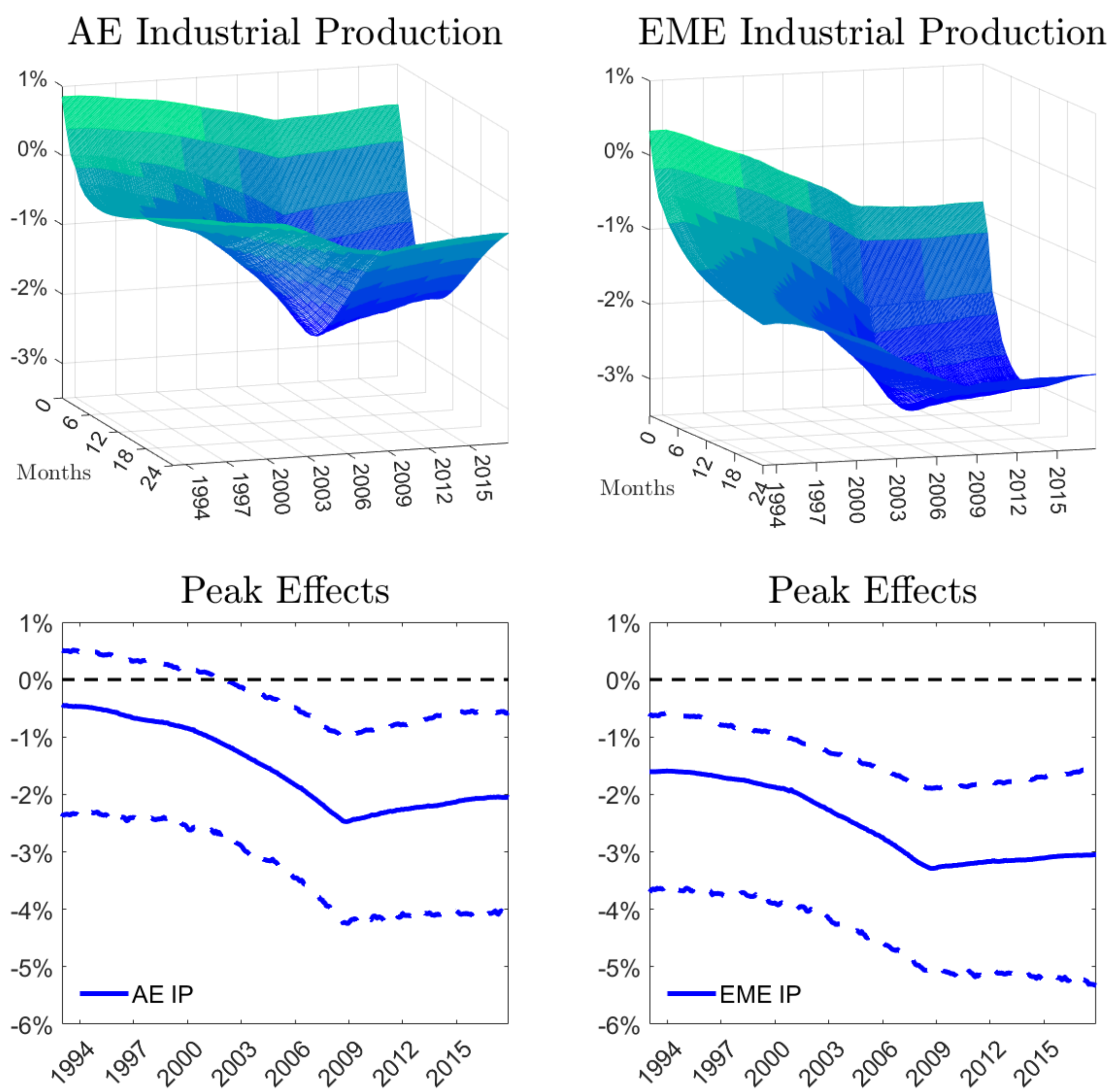
We have documented a significant time-variation in the international spillovers of U.S. monetary policy. In this and the following section, we conduct additional exercises to explain the findings and link it to different channels. As highlighted by Kalemli-Özcan (2019), De Leo et al. (2022), and Ca'Zorzi 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 of 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.[21] The results, shown in Figure 8, reveal more pronounced recessionary

[20] 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.

[21] 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 8:** 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.

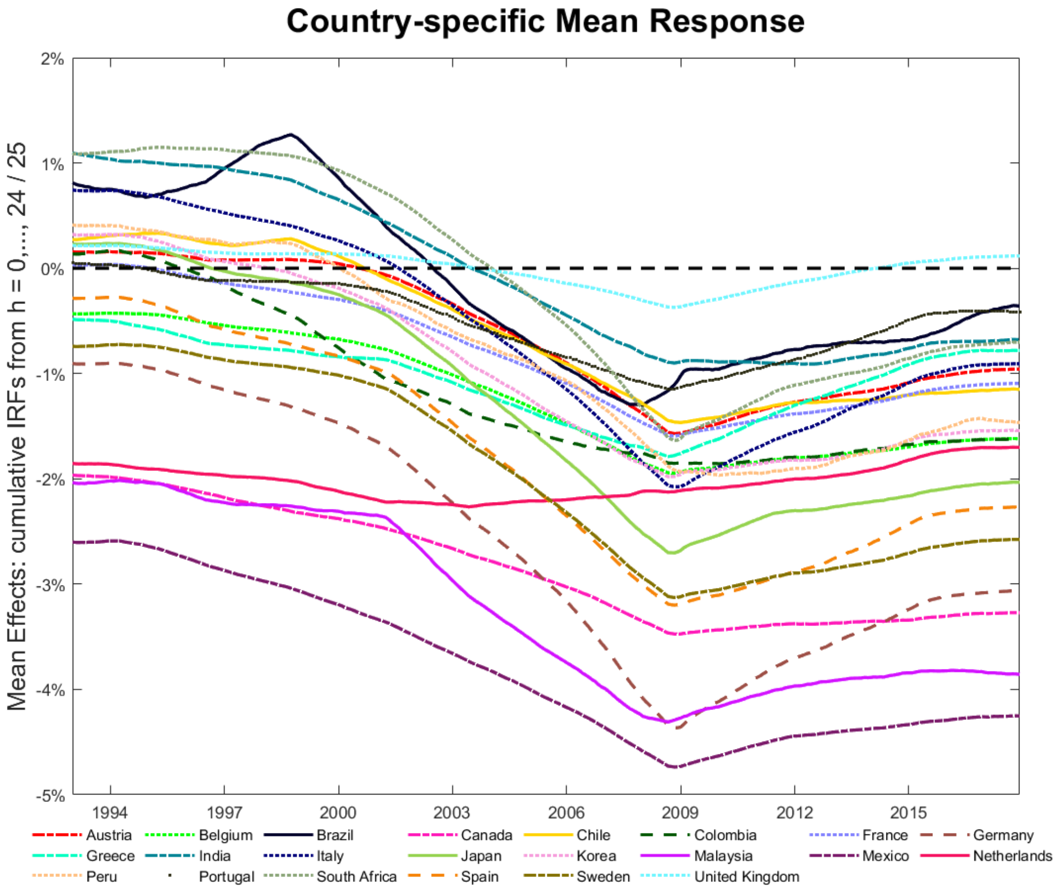
effects in emerging economies - consistent with existing studies. The peak effect on emerging (advanced) countries' industrial production is $-1.6\%$ ($-0.5\%$) in 1993 and $-3.2\%$ ($-2.4\%$) in 2008. In those years, the peak response of RoW industrial production, which combines both emerging and advanced economies, is $-0.6\%$ and $-3.2\%$, respectively. While in the beginning of the sample period the response of RoW industrial production is strongly tilted towards the one of advanced economies (emerging economies have little relevance in the overall index), the response in 2008 aligns more with the effect in EMEs (of course, estimation uncertainty must be take into account). Furthermore, the dynamic responses reveal that the response of AEs' industrial production returns back to the zero line relatively quickly, while the contraction in EMEs' industrial production is far more persistent. Since the composition of RoW industrial production between 1993 and 2008 has strongly changed (in favour 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 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 national-level indices. This allows us to examine the response at a more granular level. We consider a total of 22 countries in our analysis. These countries were selected based on two criteria: i) they are included in the aggregate RoW production measure of Federal Reserve Bank of Dallas and, ii) monthly data is available from 1980 onwards. Our sample includes: Austria, Belgium, Brazil, Canada, Chile, Colombia, France, Germany, Greece, India, Italy, South Korea, Japan, Malaysia, Mexico, Netherlands, Peru, Portugal, South Africa, Spain, Sweden, UK.[22]

To save space, we report each country's median time-varying impulse response functions in Appendix E (Figure E1-Figure E4). We also report the evolution of country-specific mean effects over time in Figure 9. While we have so far considered peak effects as summary statistics of impulse response functions, we now shift to mean effects (i.e., we report the mean effect from $h = 0, \ldots, 24$ for each $t$) to account for the heterogeneity in the shapes of the effects. Given that we observe a positive short-run response of industrial production in a significant share of countries - which would not be considered when focusing on peak effects, mean effects seem to be a more adequate measure of the effects in a given country. (However, the results are very similar when considering peak effects.) The findings in Figure 9 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 peak effect in country $i$. Secondly, the *increase* in recessionary effects can be more or less substantial, and this dimension can be

[22] 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#tab2$

**Figure 9:** Mean Effects of Country-Specific Industrial Production.




*Notes*: Mean effects in the responses of country-specific (non-U.S.) industrial production indices to a contractionary U.S. monetary policy shock that induces a one percentage point (pp) increase in the federal fund rate in 1993M1.

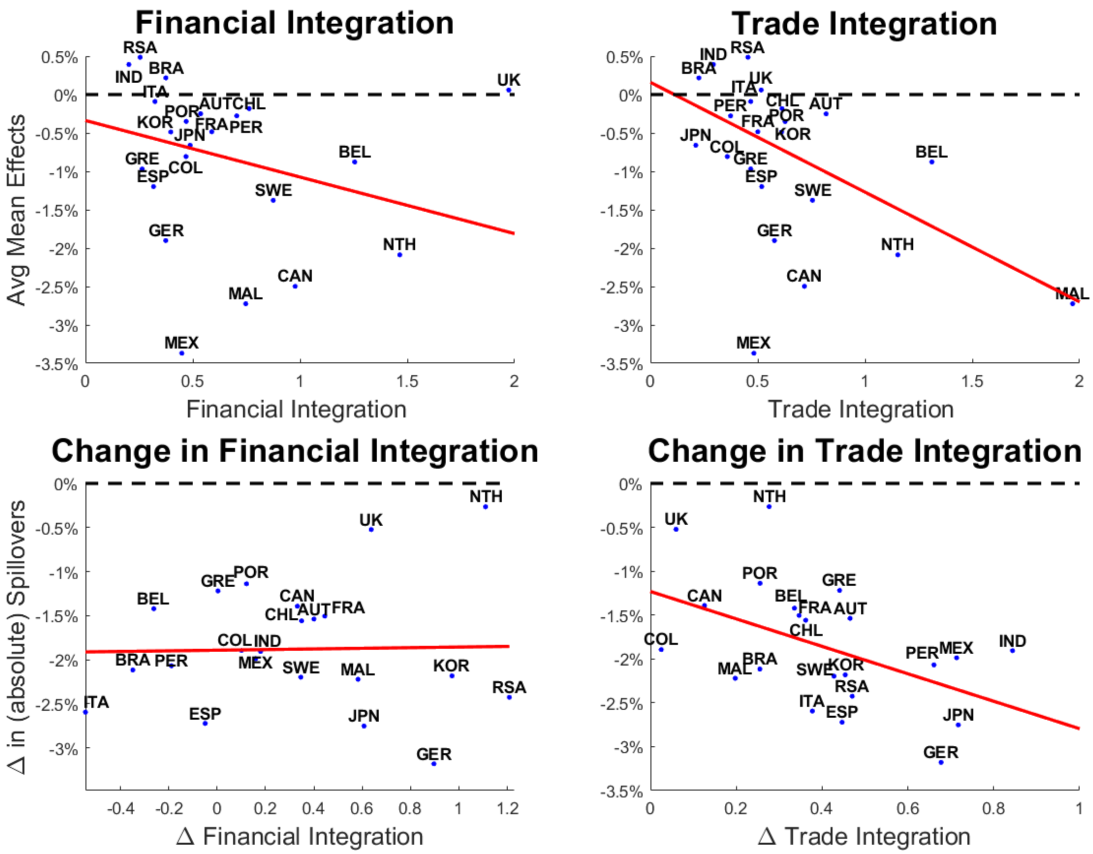
summarized by comparing the recessionary effects in country *i* at the end of the sample with those at the beginning. Time-variation is pervasive in Japan, Germany, and Spain, while it is more limited in the Netherlands and UK.

### *3.5 Evaluation of the Channels*

Our empirical analysis again points towards significant time patterns in global spillovers. But what are the main drivers of such dynamics? We shed light on this aspect by establishing a connection between country-specific effects and country-specific information on trade and financial integration. We define financial integration as the ratio of countries' (dollar-denominated) external assets and liabilities to GDP, while trade integration is determined by the ratio of total trade to GDP (for a similar choice, see Ca'Zorzi et al., 2023).[23] Financial and trade integration are then computed

[23] We retrieve the data for financial integration from Bénétrix et al. (2019). Data for trade integration is instead taken from the World Bank (World Development Indicators). Given that country-specific shares of dollar invoiced trade

**Figure 10:** Correlations of Country-Specific Responses, 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.

as the historical average of yearly measures for each country. Given that our aim 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.

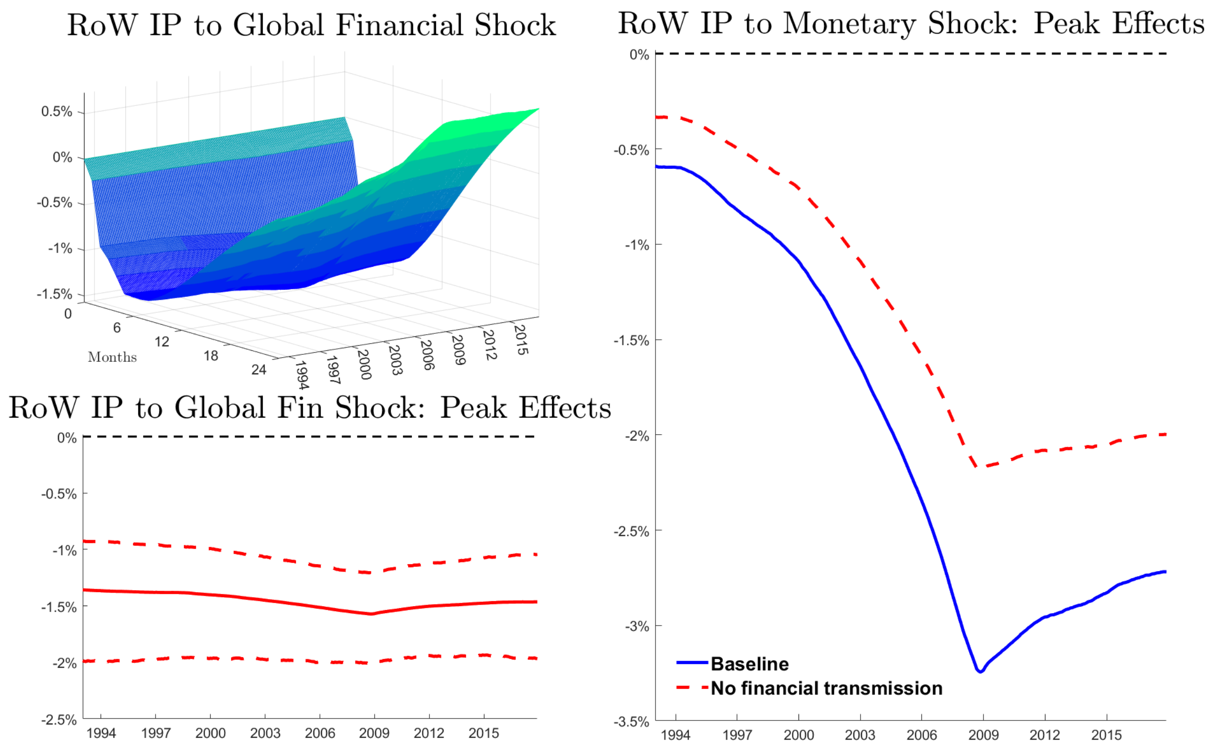
In Figure 10, we examine the correlation between the two dimensions of spillover heterogeneity and country-specific trade and financial integration measures (both in terms of *average* and *time-variation*). In the upper panel, we report the correlation of average effects with average integration,

are rarely available for the 1990s and early 2000s (see Boz et al., 2022), we consider the overall trade to GDP ratio (irrespective of currency) as a proxy for overall trade integration.

while in the lower panels we show the correlation of variation in effects with variation in integration. Looking at the upper panels, we observe that, on average, higher levels of financial and trade integration are associated with higher international spillover effects. 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*. However, when it comes to explain the *time-variation* in the effects, our estimations suggest that the trade dimension plays a more significant role (lower panels of Figure 10). Countries that undergo a substantial increase in trade integration, such as Japan and Germany, 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 (the p-value of a two-sided test is 0.03). Conversely, there is no relationship between changes in financial integration and changes in recessionary effects (p-value = 0.92). Overall, countries that witness a notable rise in financial integration are not susceptible to larger time-variation in the effects of U.S. monetary policy shocks. Table E1 in Appendix E documents that these results are confirmed when looking at peak effects or various conventional horizons. 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.

Taken at face value, this aligns with the estimates in the baseline specification. Considering the estimates presented in Figure 4, we observe a striking disconnect in the amount of time-variation observed in the effects on global trade (highly time-dependent) and financial conditions (relatively constant over time). Given this evidence, we argue that the transformation of the international financial landscape, although potentially influential, cannot account for the observed time-variation in real spillovers. While the financial transmission is likely one of the main international propagation channels *on average* (Rey, 2016; Miranda-Agrippino and Rey, 2020), it does not seem to be the main driver of the observed dynamics. This argument is based on the assumption of a constant *financial multiplier*, which is the amplification of the real consequences resulting from monetary shocks generated by global financial frictions. To substantiate this claim, we provide corroborating evidence by isolating a *global financial shock*. In the spirit of Gilchrist and Zakrajšek (2012), we estimate a recursive TVP-VAR in which the financial indicator (the global financial cycle in our case) is positioned after real activity and price indicators (slow-moving variables, which are only

**Figure 11:** Global Financial Shock: Impulse Response Functions and Counterfactual Simulations.




On the left: Responses of RoW industrial production to a "global financial shock" normalized to induce a unit deterioration of the global financial cycle. Median responses (upper panel) and peak responses (lower panel) over time. Recursive VAR with the following ordering: RoW industrial production, RoW export, U.S. industrial production, U.S. CPI, global financial cycle, U.S. policy rate. On the right: peak responses of world industrial production to US monetary shocks (blue line; baseline estimations) vs. counterfactual scenario obtained by zeroing out the response of the global financial cycle to U.S. monetary shocks via a global financial shocks (Sims and Zha, 2006).

affected with a lag), but before the U.S. policy rate (fast-moving variable, which is allowed to react contemporaneously).[24] We keep everything else as in the baseline estimations.

Figure 11 illustrates that the recessionary effects of this shock on RoW industrial production have remained constant over time (plots on the left). Digging deeper, we conduct counterfactual simulations aimed at shutting down the financial transmission of U.S. monetary policy (in the right plot of Figure 11). The blue line represents the baseline peak responses of RoW industrial production, while the red dotted lines represent the counterfactual responses in hypothetical scenarios in which U.S. monetary disturbances are no longer transmitted through global financial conditions. We construct the counterfactual simulation by generating a sequence of global financial shocks that completely offset the impact of the U.S. policy shocks on global financial cycle (see Sims and

[24] Abbate et al. (2016) employ a similar approach and find that global financial shocks have a considerable global impact. They report that changes in the transmission of global financial shocks have slightly increased, but that these changes are not statistically significant.

Zha, 2006); that is, we document how the RoW production would have been affected by monetary shocks in the absence of financial transmission.[25] Our findings reveal that, on average, the financial channel is relevant, contributing to more than one-third of the real effects. However, this channel fails to explain the observed changes in real spillovers over time. When we consider the scenario where global financial frictions are eliminated, we observe that the extent of time-variation in RoW industrial production remains comparable to the baseline estimations.

We hence conjecture that most of the variation is attributable to macroeconomic and trade-related factors. A battery of different results support this. First, the increasing response of RoW export (as opposed to a constant response of global financial conditions) signals a more pronounced real transmission through trade channels. Second, we show the potential relevance for spillovers of the economic rise of emerging economies via composition effects. Third, cross-sectional correlation analysis indicates rising trade integration as a likely driver of rising spillovers. Fourth, counterfactual simulations aimed at shutting down the role of financial transmission of U.S. policy disturbances attribute a limited role to financial linkages as a potential diver of the observed dynamics in spillovers.

### *3.6 Sensitivity Analysis*

In this final section, we assess the robustness of our main findings by exploring different specifications. All results are reported in Appendix F.

**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). A comparison of the peak effects in the baseline to the alternative specification can be found in Figure F1. The results align well with our baseline estimations.

**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: $\boldsymbol{\Sigma}_t = \boldsymbol{A}_t^{-1}\text{diag}\left(\exp(h_{1t}), \ldots, \exp(h_{nt})\right) \boldsymbol{A}_t^{-1\prime}$. 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. The results of this exercise are presented in Figure F2 (where we also report further estimation details). If we compare the global responses to the baseline estimates, the qualitative pattern is the same.

[25] McKay and Wolf (2023) propose a refinement that involves subjecting the economy to financial shocks only at horizon $h = 0$ in order to closely approximate the counterfactual scenario. However, that approach would require the identification of multiple global financial shocks. This proves difficult in our application.

Time-variation in real spillover is present, while this is not the case when it comes to the effects on global financial conditions.

**Alternative Monetary Policy Instrument**. There exists a range of high-frequency instruments for the identification of monetary policy. The first versions of the instruments do not control for the *information effect* (Melosi, 2017; Nakamura and Steinsson, 2018). Therefore, we use the monetary policy instrument of Miranda-Agrippino and Ricco (2021) which controls for this by projecting market-based monetary surprises on the Fed's information set (proxied by Greenbook forecasts). To provide robustness, we also re-estimate the model using the original instruments of Gertler and Karadi (2015) and the "poor-man" refinement proposed in Jarociński and Karadi (2020) (that exploits high-frequency co-movements of federal funds futures surprises and stock price to eliminate the *information effect*). Results of these checks, which strongly overlap with the estimates of the baseline model, are reported in Figure F3.

**Prior Selection**. We investigate the sensitivity of our results to different calibrations of $\kappa_Q^2$, 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). Figure F4 show that the responses are qualitatively similar when using both tighter ($\kappa_Q^2 = 0.01$) and wider choices ($\kappa_Q^2 = 0.02$).

## 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 TVP-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 of 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 the one 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 transmission of U.S. policy disturbances.

Finally, a policy implication of this paper is that in a world with increasing spillovers, policies in the Rest of the World (RoW) need careful calibration in response to possibly time-varying output/ inflation effects of U.S. policies. Furthermore, the rising spillovers likely lead to a rise in domestic repercussions. Our estimations reveal that the influence of U.S. policy shocks on the domestic front has intensified over time. This phenomenon may be a direct result of the increasing international spillover effects, which subsequently magnify the corresponding domestic consequences.

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## A. Data

All series were gathered from the sources listed below, particularly the FRED and several papers. In Table A1, we define the exact transformations of the variables used in the estimation of the TVP-VAR. Note that we use month-on-month growth rates.

In specific cases (indicated by stars $^*$ in the Table A1), we are missing data points in the pre-sample period. In order to estimate the TVP-VAR on the same estimation sample (for which all data points are available), we pursue the following strategy. If we have information on a similar series for the same time period, we assume the same growth scenario. Specifically, we observe industrial production of EMEs only from 1987M1 to 2017M12. Hence, we assume that industrial production in EMEs grows as the corresponding series in the RoW during the time frame 1980M1 to 1986M12. This allows us to extend the series backwards.

**Table A1:** Variable Definitions.

| Variable | Transformation in TVP-VAR | Details |
|---|---|---|
| $ffr_t$ | $\texttt{FFR}_t$ | $\texttt{FFR}_t$ is the federal funds rate (Fred), replaced with the shadow rate of Wu and Xia (2016) during the ZLB period (2008M12-2015M12). |
| $cpi_t^{US}$ | $100 \times \left[\ln\left(\texttt{CPI}_t^{US}\right) - \ln\left(CPI_{t-1}^{US}\right)\right]$ | $\texttt{CPI}_t^{US}$ is U.S. consumer price index (Fred). |
| $ip_t^{US}$ | $100 \times \left[\ln\left(\texttt{IP}_t^{US}\right) - \ln\left(\texttt{IP}_{t-1}^{US}\right)\right]$ | $\texttt{IP}_t^{US}$ is U. S. industrial production (Fred). |
| $xim_t^{US}$ | $\texttt{XIM}_t^{US}$ | $\texttt{XIM}_t^{US}$ is the ratio between U.S. export and import in goods and services (Fred), interpolated from quarterly to monthly using a shape-preserving piecewise cubic interpolation. |
| $ebp_t^{US}$ | $\texttt{EBP}_t^{US}$ | $\texttt{EBP}_t^{US}$ is the U.S. excess bond premium by Gilchrist and Zakrajšek (2012). |
| $gfc_t$ | $\texttt{GFC}_t$ | $\texttt{GFC}_t$ is the global financial cycle indicator by Miranda-Agrippino and Rey (2020). |
| $reer_t$ | $100 \times \ln\left(\texttt{REER}_t^{US}\right)$ | $\texttt{REER}_t^{US}$ is the U.S. real effective exchange rate from the BIS. |
| $cpi_t^{RoW}$ | $100 \times \left[\ln\left(\texttt{CPI}_t^{RoW}\right) - \ln\left(\texttt{CPI}_{t-1}^{RoW}\right)\right]$ | $\texttt{CPI}_t^{RoW}$ is RoW consumer prices (excluding the U.S.) Grossman et al. (2014). |
| $ip_t^{RoW}$ | $100 \times \left[\ln\left(\texttt{IP}_t^{RoW}\right) - \ln\left(\texttt{IP}_{t-1}^{RoW}\right)\right]$ | $\texttt{IP}_t^{RoW}$ is the RoW industrial production (excluding the U.S.) by Grossman et al. (2014). |
| $msci_t^{RoW}$ | $100 \times \left[\ln\left(\texttt{MSCI}_t^{RoW}\right) - \ln\left(\texttt{MSCI}_{t-1}^{RoW}\right)\right]$ | $\texttt{MSCI}_t^{RoW}$ is RoW MSCI (excluding the U.S.). |
| $ex_t^{RoW}$ | $100 \times \left[\ln\left(\texttt{EX}_t^{RoW}\right) - \ln\left(\texttt{EX}_{t-1}^{RoW}\right)\right]$ | $\texttt{EX}_t^{RoW}$ is RoW exports (excluding the U.S.) by Grossman et al. (2014). |
| $ip_t^{AE}$ | $100 \times \left[\ln\left(\texttt{IP}_t^{AE}\right) - \ln\left(\texttt{IP}_{t-1}^{AE}\right)\right]$ | $\texttt{IP}_t^{AE}$ is the advanced economies' industrial production (excluding the U.S.) by Grossman et al. (2014). |
| $ex_t^{AE}$ | $100 \times \left[\ln\left(\texttt{EX}_t^{AE}\right) - \ln\left(\texttt{EX}_{t-1}^{AE}\right)\right]$ | $\texttt{EX}_t^{AE}$ is AE exports (excluding the U.S.) by Grossman et al. (2014). |
| $ip_t^{EME}$ | $100 \times \left[\ln\left(\texttt{IP}_t^{EME}\right) - \ln\left(\texttt{IP}_{t-1}^{EME}\right)\right]$ | $\texttt{IP}_t^{EME}$ is the emerging market economies' industrial production by Grossman et al. (2014).* |
| $ex_t^{EME}$ | $100 \times \left[\ln\left(\texttt{EX}_t^{EME}\right) - \ln\left(\texttt{EX}_{t-1}^{EME}\right)\right]$ | $\texttt{EX}_t^{EME}$ 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: $\texttt{EME IP}_t$ 1987M1 to 2017M12.

## B. TVP-VAR: Prior Settings and 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 prior densities for the covariance matrices $\boldsymbol{Q}$, where we use an inverse-Wishart prior density.[26] Last, we also need a prior distribution for the covariance matrix of the VAR $\boldsymbol{\Sigma}$, which is again following 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. 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 $\kappa_Q^2$, 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. An additional note is in order: We observe the high-frequency surprises not until 1991M1. Following the procedure of Paul (2020), we plug in zeros for the observations prior to this period. This should not cause a bias in OLS as long as those zeros are from a random sample. This should be indeed the case for the monetary policy surprises. Nevertheless, we estimate those coefficients with less precision.

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 $\boldsymbol{\theta}^T = (\boldsymbol{\theta}_1', \dots, \boldsymbol{\theta}_T')'$. 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.

[26] We denote by $iW(S, d)$ an inverse-Wishart distribution with degrees of freedom $d$ and scale matrix $S$.

## C. Same-Sized Shocks to the U.S. Policy Rate

In this section, we take a closer look at the on impact response of the U.S. policy rate. Same-sized shocks are retrieved from the procedure outlined in Paul (2020), which identifies *relative* impulse responses. The critical equation in this procedure is Equation 2.3, which assumes a constant relationship between the instrument and the structural monetary policy shock.

As the right panel of Figure C1 suggests, the on impact response of the U.S. policy rate declines over the sample period. We normalize it to one percentage point (1 pp) in the first period of the sample, 1993M1. Afterwards, given our assumptions, the on impact response shows same-sized shocks over time. In the 2000s the on impact response starts to drastically decline before it levels out at about 0.2 pp in 2009, around the Great Financial Crisis. Afterwards, it slowly starts to increase again. The interpretation of a same-sized shock is that 1 pp monetary policy contraction in 1993M1 has the same economic size as a 0.2 pp monetary policy contraction in 2009M1.

We relate this to the long-run trend in the U.S. policy rate. Similar to the analysis in Liu et al. (2022), we first transform the TVP-VAR in Equation 2.1 in its companion form

$$\boldsymbol{z}_t = \boldsymbol{\mu}_t + \boldsymbol{\alpha}_t \boldsymbol{z}_{t-1} + \boldsymbol{\eta}_t,$$
$$\boldsymbol{z}_t := \begin{bmatrix} \boldsymbol{y}_t \\ \boldsymbol{y}_{t-1} \\ \vdots \\ \boldsymbol{y}_{t-p+1} \end{bmatrix}, \quad \boldsymbol{\mu}_t := \begin{bmatrix} \boldsymbol{c}_t \\ \boldsymbol{0} \\ \vdots \\ \boldsymbol{0} \end{bmatrix}, \quad \boldsymbol{\alpha}_t := \begin{bmatrix} \boldsymbol{A}_{1t} & \boldsymbol{A}_{2t} & \dots & \boldsymbol{A}_{pt} \\ \boldsymbol{I}_n & \boldsymbol{0} & \dots & \boldsymbol{0} \\ \vdots & \ddots & \ddots & \vdots \\ \boldsymbol{0} & \dots & \boldsymbol{I}_n & \boldsymbol{0} \end{bmatrix}, \quad \boldsymbol{\eta}_t := \begin{bmatrix} \boldsymbol{u}_t \\ \boldsymbol{0} \\ \vdots \\ \boldsymbol{0} \end{bmatrix}. \tag{C.1}$$

The stability condition implies that the roots of the polynomial $\varphi(z) = \det\left(\boldsymbol{I}_n - \sum_{j=1}^{p} z^p \boldsymbol{A}_{jt}\right)$ are below one. Following Giraitis et al. (2018), this allows us to approximate the companion form by an VMA($\infty$) process of the form

$$\boldsymbol{y}_t = (\boldsymbol{I}_n - \boldsymbol{\alpha}_t)^{-1}\boldsymbol{\mu}_t + \sum_{h=0}^{\infty} \boldsymbol{\alpha}_t^h \boldsymbol{\eta}_{t-j} + o_p(1). \tag{C.2}$$

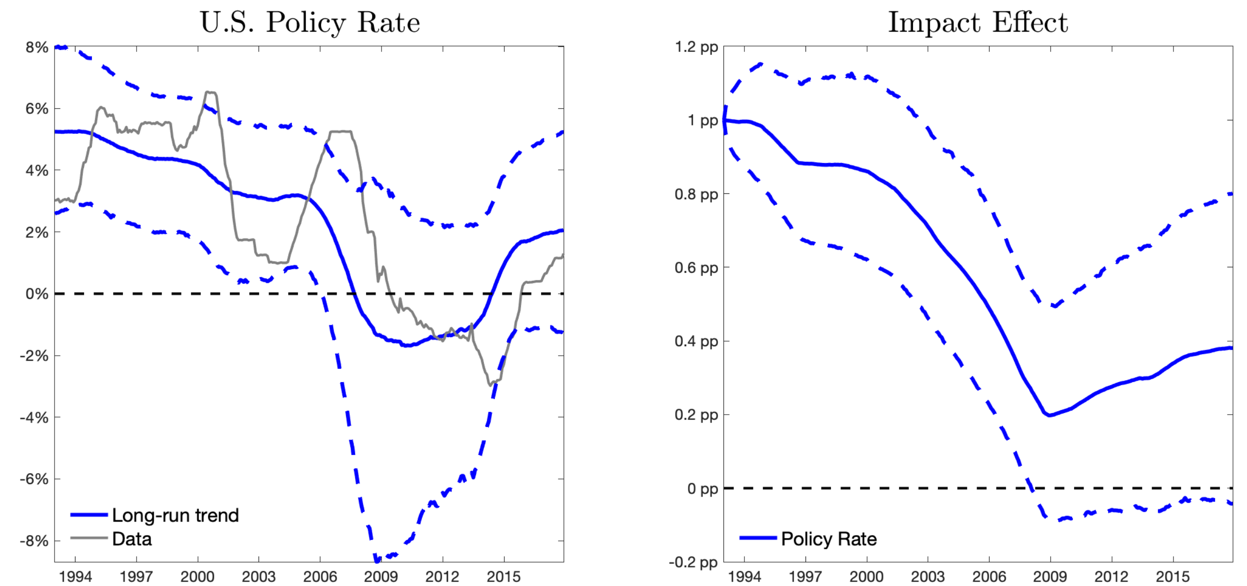
We use this approximation to compute the implied trends of the model's variables:

$$\boldsymbol{\tau}_t = (\boldsymbol{I}_n - \boldsymbol{\alpha}_t)^{-1}\boldsymbol{\mu}_t, \tag{C.3}$$

where the elements of $\boldsymbol{\tau}_t$ can be interpreted as the long-run economic expectations or infinite-horizon forecasts implied by the model. These have been used by Cogley and Sargent (2005) and Liu et al. (2022) to study the natural rates or long-run trends. Specifically, we are interested in the implied long-run, or natural, trend of the U.S. policy rate.

The left panel of Figure C1 shows the estimates of the implied long-run trend of the U.S. policy rate along with the 68% credible sets and the underlying data (federal funds shadow rate). Similar to other findings in the literature, we also find that the long-run trend in the federal funds rate declines

**Figure C1:** Long-Run Trend and On Impact Response of U.S. Policy Rate.

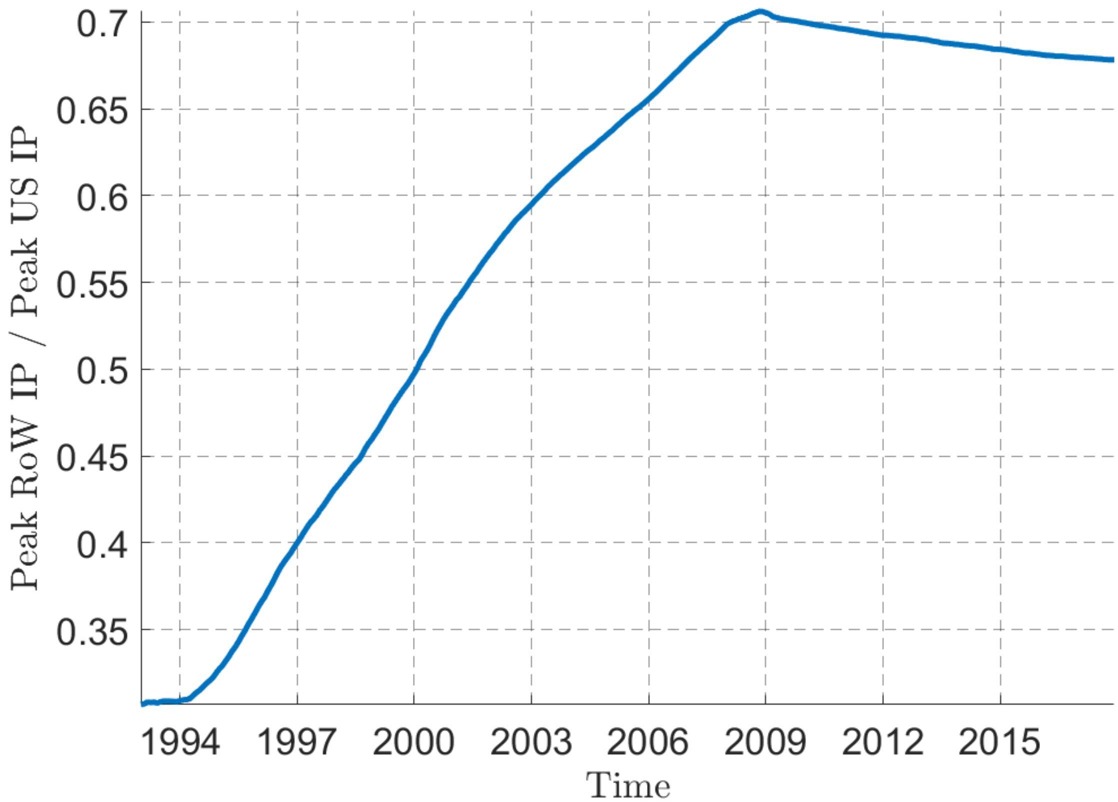



*Notes*: The left plot shows the actual U.S. policy rate along with the estimate of its long-run trend (unconditional mean). Bands depict the 68% credible intervals. The right plot shows the on impact response of the U.S. policy rate along with its 68% credible sets.

over time. While it was around 5% in 1993, it declines to 3% before the onset of the Great Recession before it drastically falls even below the zero line. Only in the later periods of the sample, the trend breaks again through the zero line. With a reduction in the long-run mean, there is also a decrease in volatility. We argue that this decrease is responsible for the decline in the on impact responses of same-sized monetary policy shocks.

## D. Additional Results of the Baseline Specification

**Figure D1:** Ratio of Peak Responses: RoW IP and U.S. IP.



## E. Additional Results of Country-Specific Effects

In this section, we report additional results of country-specific effects. Figure E1 - Figure E4 report the country-specific impulse responses of industrial production for each country separately.

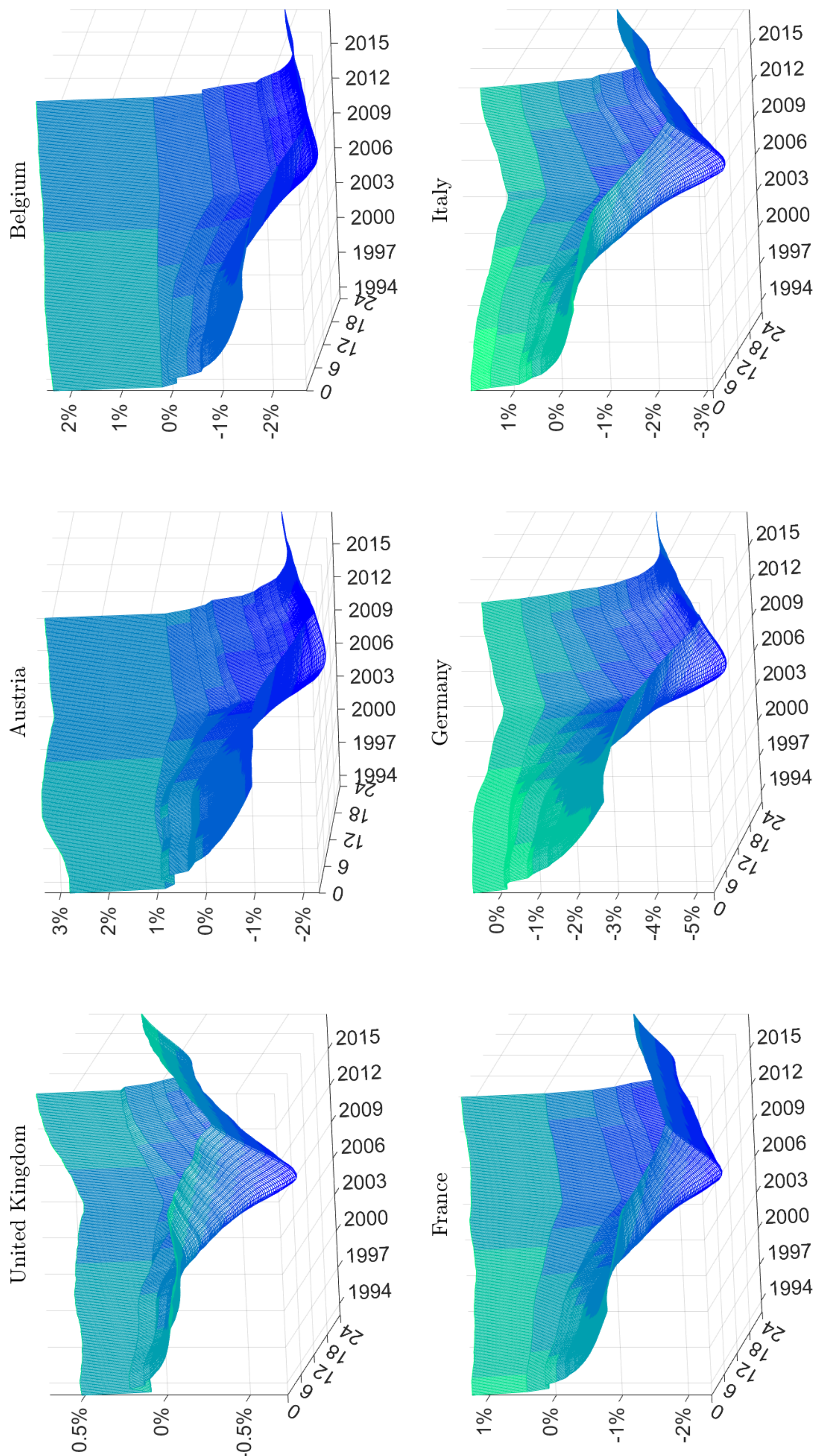
In Figure 10, we report the graphical illustration of two sets of regressions: i) average mean effects in country $i$ as a function of average financial and trade integration in country $i$; ii) change in the effects in country $i$ as a function of change in financial and trade integration in country $i$. (Given our small sample, we reduce the impact of outliers by employing an iteratively reweighted least-squares algorithm.) These findings are detailed in Table E1. Furthermore, our results remain consistent when we shift focus from mean effects (calculated as the cumulative impulse response from $h = 0$ to $h = 24$ months, divided by 25) to peak effects and effects at selected horizons (0, 6, 12, 18, 24 months).

**Table E1:** Regression Outcomes.

| | *Mean* | *Peak* | $h = 0$ | $h = 6$ | $h = 12$ | $h = 18$ | $h = 24$ |
|---|---|---|---|---|---|---|---|
| | | *Average* | | | | | |
| Trade integration | **−1.43** | **−1.62** | **−3.07** | **−1.57** | **−1.40** | **−1.25** | −1.12 |
| | (0.47) | (0.40) | (0.82) | (0.55) | (0.51) | (0.53) | (0.59) |
| Financial integration | −0.74 | −0.75 | −0.65 | −0.41 | **−1.20** | **−1.35** | −1.03 |
| | (0.54) | (0.60) | (0.95) | (0.60) | (0.54) | (0.52) | (0.58) |
| | | *Time-variation* | | | | | |
| Δ Trade integration | **−1.56** | −1.25 | −0.70 | **−1.85** | **−1.93** | **−1.64** | **−1.43** |
| | (0.68) | (1.10) | (0.88) | (0.87) | (0.79) | (0.66) | (0.58) |
| Δ Financial integration | 0.04 | 0.07 | −0.18 | −0.15 | 0.06 | 0.22 | 0.28 |
| | (0.35) | (0.52) | (0.41) | (0.44) | (0.44) | (0.34) | (0.30) |

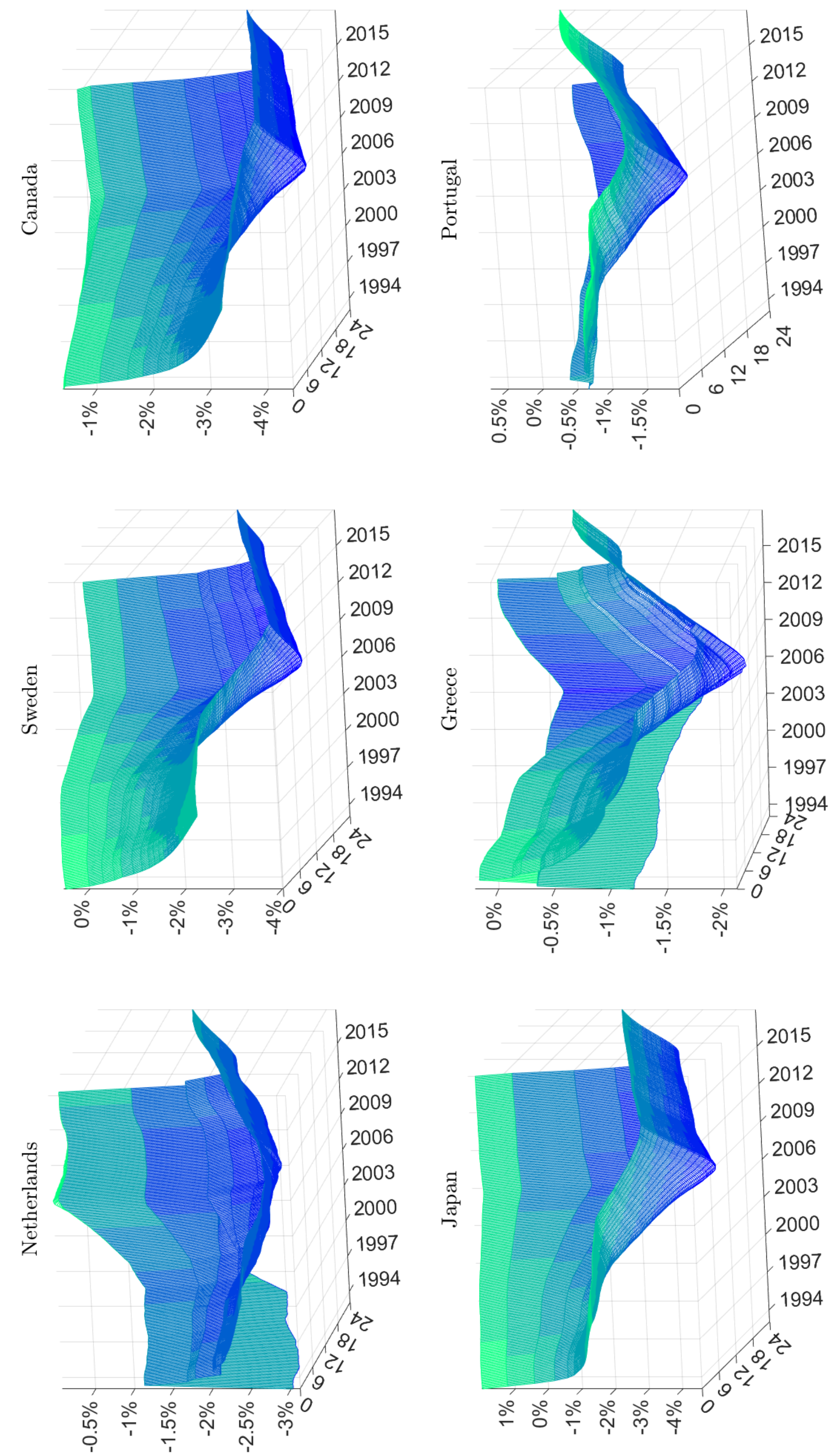
*Notes:* Regression based on 22 observations with *average* or *time-variation* effects in industrial production per country as dependent variable. Independent variable is trade or financial integration, either in the level or as log change. Number in parenthesis report robust standard error of slope coefficients. Bold numbers indicate statistical significance at 5% level.

**Figure E1:** Country-Specific Industrial Production Impulse Responses.



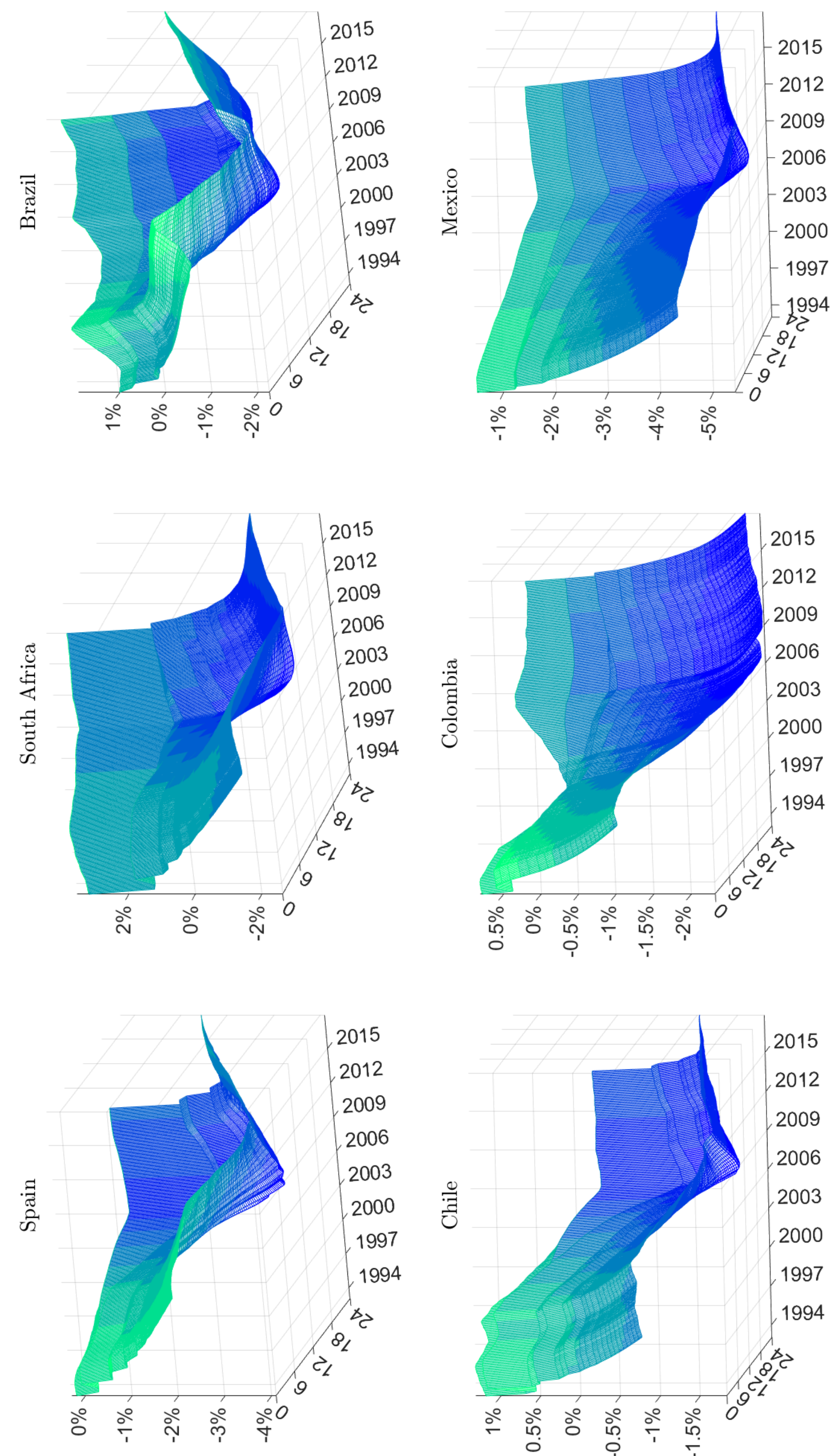

*Notes*: Responses to a contractionary U.S. monetary policy shock that induces a one percentage point (pp) increase in the federal fund rate in 1993M1. Countries: UK, Austria, Belgium, Canada, France, Germany, and Italy.

**Figure E2:** Country-Specific Industrial Production Impulse Responses.



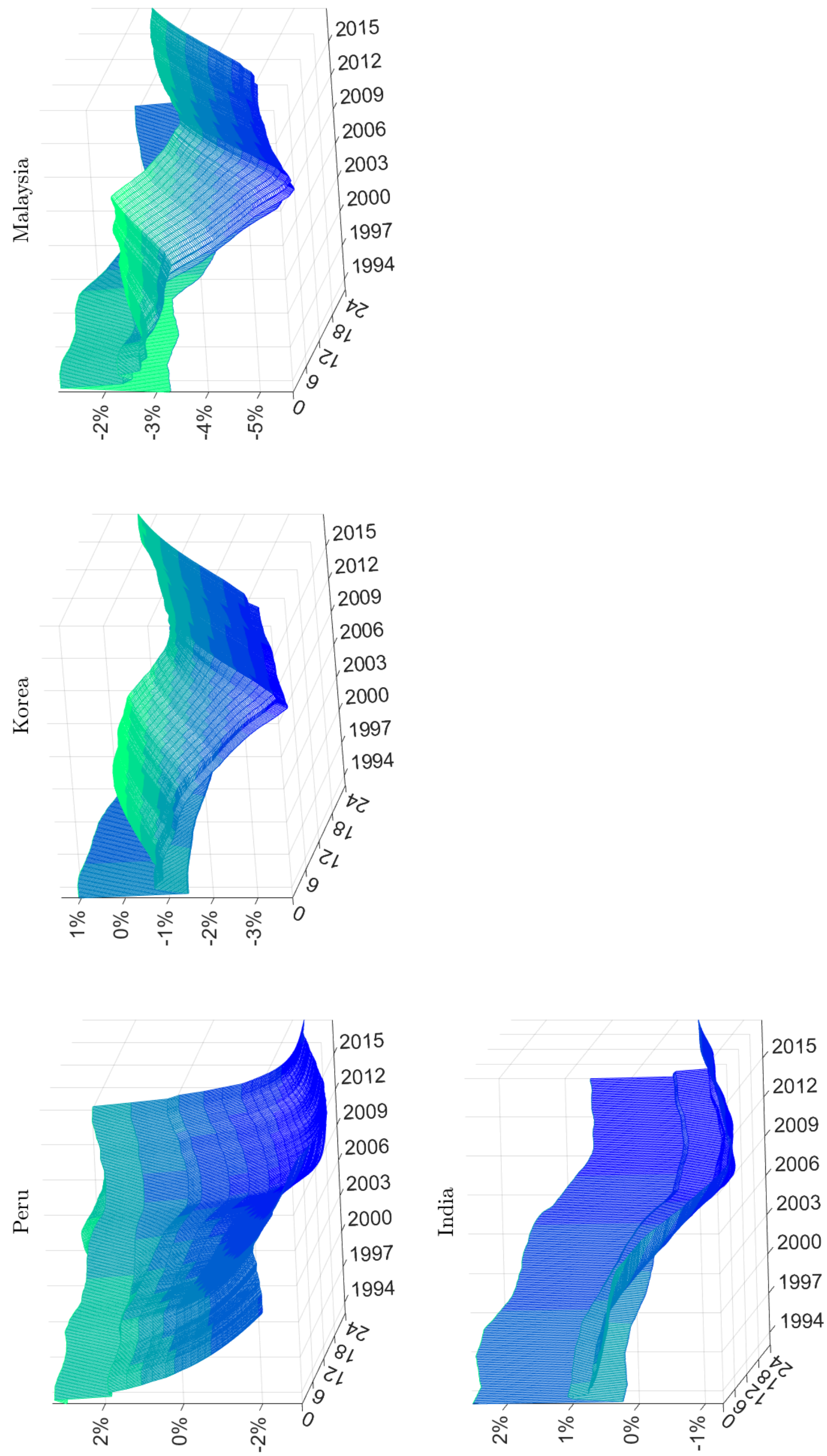

*Notes*: Responses to a contractionary U.S. monetary policy shock that induces a one percentage point increase in the federal fund rate in 1993M1. Countries: Netherlands, Sweden, Canada, Japan, Greece, Portugal.

**Figure E3:** Country-Specific Industrial Production Impulse Responses.




*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: Spain, South Africa, Brazil, Chile, Columbia, Mexico.

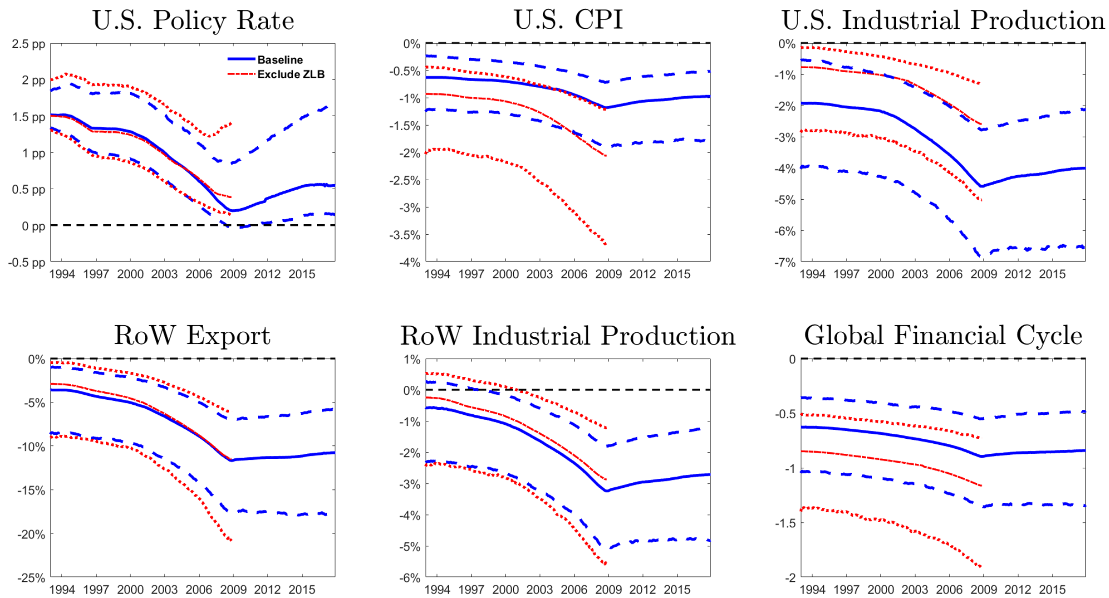
**Figure E4:** Country-Specific Industrial Production Impulse Responses.




*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.

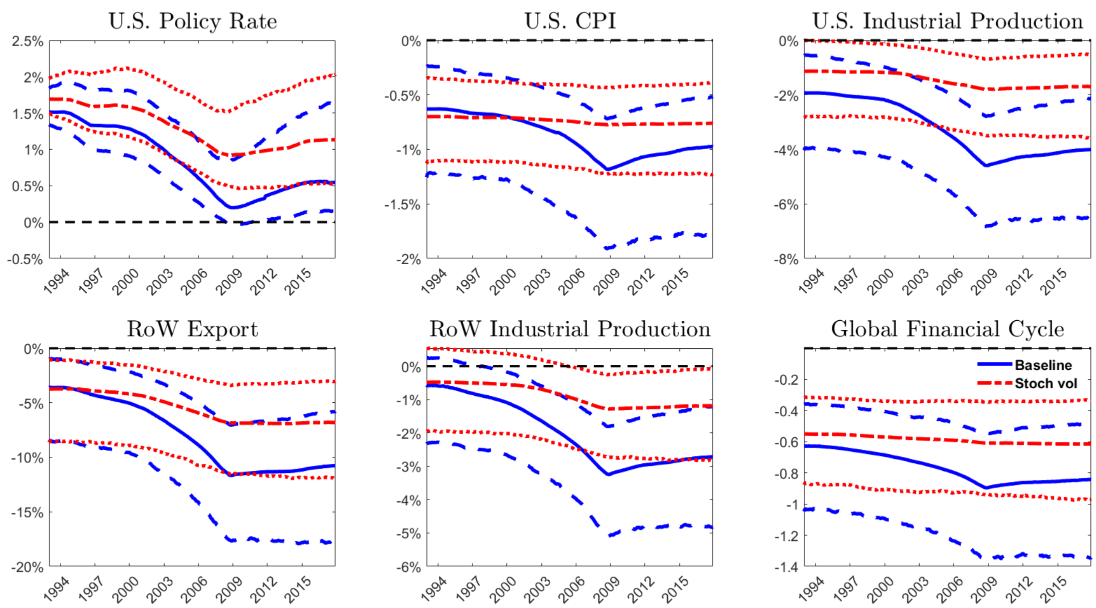
## F. Additional Results of the Sensitivity Analysis

**Figure F1:** Robustness: Zero Lower Bound.




*Notes*: Responses to a contractionary U.S. monetary policy shock that induces a one percentage point increase in the federal fund rate in 1993M1. Baseline model in blue; red lines correspond to the model excluding the ZLB period (ending in 2008M11).

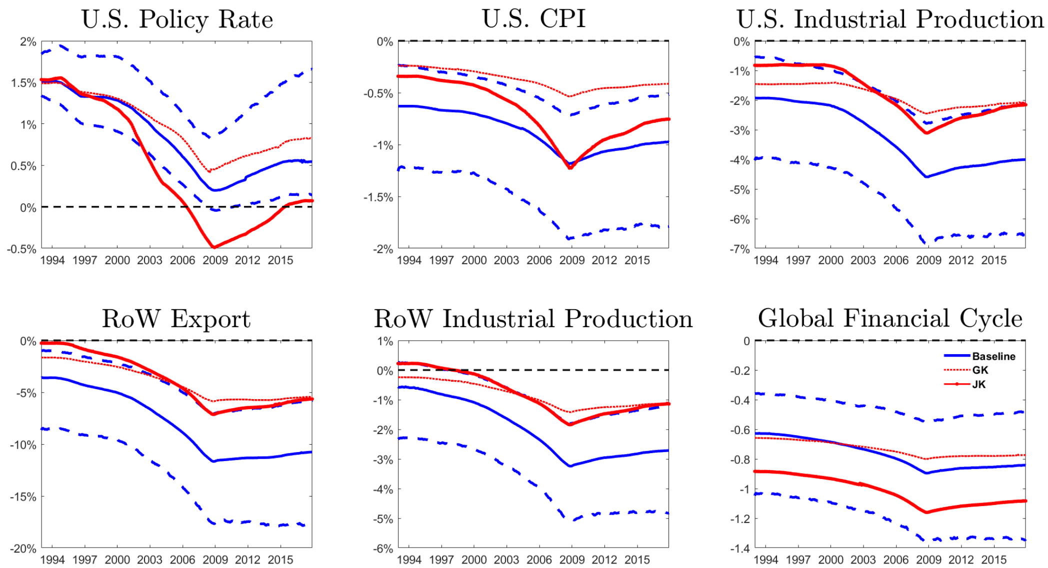
**Figure F2:** Robustness: Stochastic Volatility.




*Notes*: Peak Responses to a contractionary U.S. monetary policy shock that induces a one percentage point increase in the federal fund rate in 1993M1. Baseline model vs. extended model that allows for stochastic volatility specification.

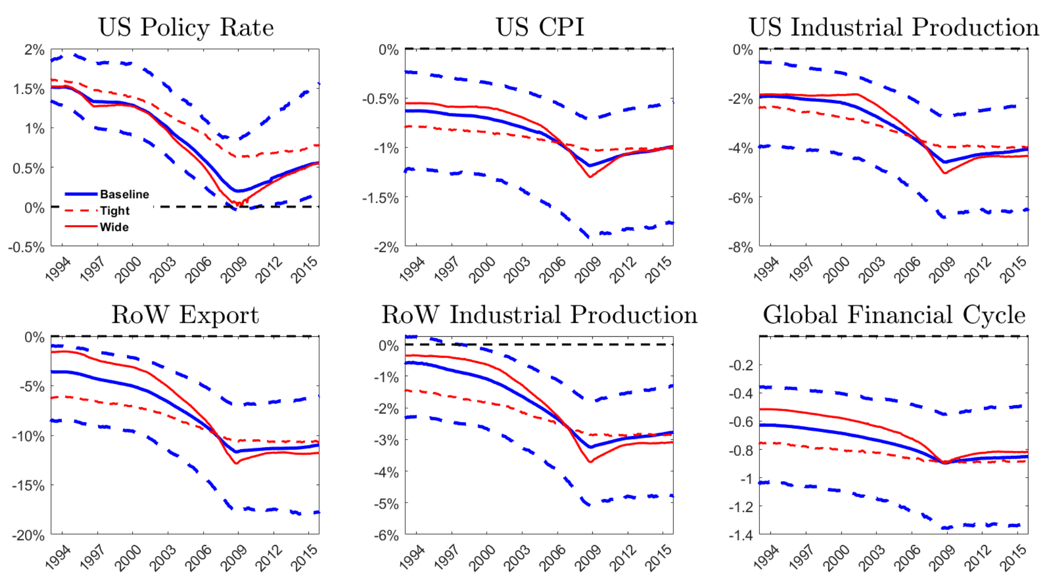
Details on stochastic volatility: As in our benchmark analysis and in Primiceri (2005), we use OLS estimates on a pre-sample (1980M1-1992M12) to calibrate the prior distributions. In addition, a prior belief on the extent of time-variation in $A_t^{-1}$ and $h_{it}$ $(i = 1, \ldots, n)$ must be specified. Using the notation of Primiceri (2005), this boils down to a selection choice on three parameters: $\kappa_Q$ (governing time-variation of autoregressive coefficients), $\kappa_W$ (variance of the residuals), $\kappa_S$ (covariance of the residuals). We set $\kappa_Q = 0.015$ (as before), $\kappa_W = 0.001$, and $\kappa_S = 0.001$. The value of $\kappa_W$ is among the ones considered by Primiceri (2005), while our $\kappa_S$ 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 F3:** Robustness: Different Monetary Policy Instruments.




*Notes*: Responses to a contractionary U.S. monetary policy shock that induces a one percentage point increase in the federal fund rate in 1993M1. Baseline model in blue (instrument of Miranda-Agrippino and Ricco, 2021); red lines correspond to the estimates using the high-frequency monetary policy instruments of Gertler and Karadi (2015) (GK) and Jarociński and Karadi (2020) (JK).

**Figure F4:** Robustness: Prior Calibration.




*Notes*: Responses to a contractionary U.S. monetary policy shock that induces a one percentage point increase in the federal fund rate in 1993M1. Baseline model in blue; red lines corresponds to the estimates using a tighter ($\kappa_Q^2$ = 0.01) and wider ($\kappa_Q^2$ = 0.02) prior against the baseline prior ($\kappa_Q^2$ = 0.015).