

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 actions taken by the Federal Reserve play a growing role in shaping global economic activity. The adverse international effects of a U.S. tightening have more than tripled 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 time-variation is limited.

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

JEL Codes: C32, E32, E44, E71, G41.

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1. Introduction

Globalization has led to unprecedented levels of worldwide economic integration, encompassing both trade and cross-border financial flows. As depicted in Figure 1, conventional measures of global trade and financial integration exhibited a substantial upward trend over the last forty years, slowing down only with the onset of the Great Recession. More importantly, the U.S. dollar is the dominant currency in trade invoicing and issuance of financial assets (Gopinath, 2015; Rey, 2016). In this context, the pivotal role of the dollar as an international currency brought many observers to view the Federal Reserve (Fed) as a world banker.¹ This implies that U.S. monetary policy decisions lead to international spillovers. In this paper, we examine whether globalization has changed the international transmission of U.S. monetary policy.

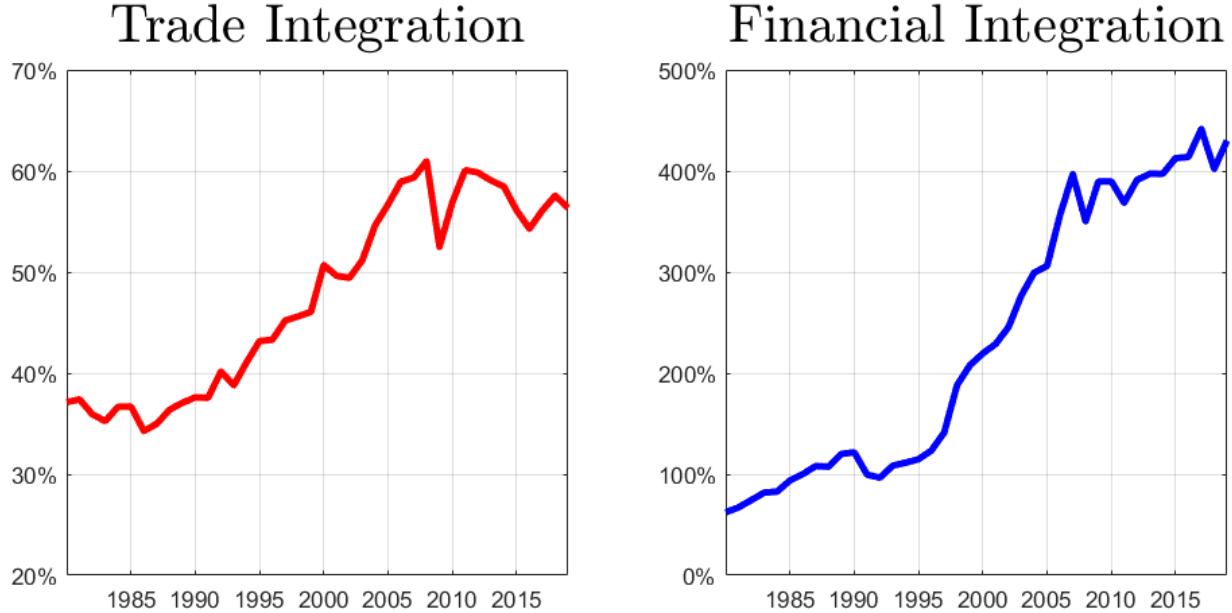
There is substantial evidence showing that a monetary tightening engineered by the Fed, which is transmitted primarily through trade and financial frictions, generates global recessionary effects (e.g., Dedola et al., 2017, Iacoviello and Navarro, 2019, Miranda-Agrippino and Rey, 2020; De-gasperi et al., 2020, Gopinath et al., 2020, or Breitenlechner et al., 2022). Yet, it is less clear how globalization has affected the international transmission of U.S. policy shocks. Due to the dominant role of the dollar, the large and increasing interconnectedness of global real and financial markets could have significantly exacerbated international spillovers. However, the existing literature has largely overlooked this time dimension. This paper aims to fill this gap by examining the evolving interaction between trade and financial integration and global spillovers to monetary policy shocks.

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 that is mediated through the reaction of U.S. industrial production, U.S. prices, as well as global economic, trade, and financial indicators. To gauge the effects on the global business cycle and trade, we rely on the global production and trade indices published by the Federal Reserve Bank of Dallas (Grossman et al., 2014), which exclude U.S. economic activity and allow us to focus on the transmission of policy shocks to the rest of the world (RoW henceforth). We use the global factor in international risky asset prices by Miranda-Agrippino and Rey (2020) as a proxy for the global financial cycle. Identification of U.S. monetary policy is achieved through the high-frequency monetary policy instrument proposed by Miranda-Agrippino and Ricco (2021), which directly controls for the information channel of monetary policy. The estimation of the TVP-VAR is done in a Bayesian fashion following Primiceri (2005) and Paul (2020).

The international transmission channels of monetary policy are manifold, potentially offsetting, and hard to identify. They are summarized into a trade and financial channel. The former refers

¹ See Gourinchas and Rey (2007), Gopinath (2015), Rey (2016), Gopinath and Stein (2018), Gourinchas et al. (2019), Ilzetzki et al. (2019) and Boz et al. (2022).

Figure 1: Motivating Evidence: Global Trade and Financial Integration as Share of GDP.



Notes: Global trade integration is defined as the sum of global exports and imports as a percentage of GDP (World Bank national accounts data). Global financial integration is defined as the sum of global external assets and liabilities as a percentage of GDP (data from the External Wealth of Nations database; Lane and Milesi-Ferretti (2018)). See Ca'Zorzi et al. (2023) for similar definitions.

to the traditional Mundell-Fleming effects (e.g., Obstfeld and Rogoff; Obstfeld and Rogoff, 2005; 2007), 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 depress demand for imports (RoW export). This demand effect should be partly counteracted through a revaluation of the currency, which leads to expenditure-switching from home goods to foreign goods. However, since U.S. imports are mostly 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. itself. 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 output spillovers in response to the appreciation of the dollar (Gopinath et al., 2020; Georgiadis and Schumann, 2021).² Given that

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.

trade integration has significantly increased over time, spillovers originating from macroeconomic linkages may have increased. Furthermore, as discussed in Farhi and Werning (2014), Bruno and Shin (2015b), and Rey (2016), 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. This, in turn, 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 (Bruno and Shin, 2015a; Miranda-Agricoppino and Rey, 2020). Particularly, the existence of the global financial cycle (GFC, see Rey, 2015, 2016) acts as a transmitter for the financial channel. Again, the currency plays an important role. The dollar dominates the global financial system and is an important transmitter of global risk shocks (Georgiadis et al., 2021, 2023). Against the myriad of 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 the course of 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 stabilize with the onset of the Great Recession, which is consistent with the slow-down in trade and financial integration that we see in the data (Ca'Zorzi et al., 2023). The observed time-variation is substantial and statistically significant. While a 1% hike in monetary policy leads to about a -0.7% contraction in RoW industrial production in 1993, a same-sized shock in 2008 results in a downturn of about -3.4% . 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, when looking at the time-variation, we find a disconnect between the two channels. While the effects on RoW trade are significantly time-varying and track well the pattern that we find for RoW industrial production, the ones on global financial conditions are relatively constant over time. 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 the need of incorporating spillback effects in the calibration of U.S. policy decisions today more than in the past (see also Yellen, 2019; Lane, 2019; Obstfeld, 2020; and Brainard, 2022).

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 on this aspect by performing a heterogeneity analysis and counterfactual simulations. Regarding the heterogeneity analysis, we generally find stronger spillover effects for emerging market (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 16 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*, as shown by a broad array of papers (Kalemlı-Özcan, 2019; Degasperi et al., 2020; Miranda-Agrippino et al., 2020). However, when it comes to explaining the *variation* in effects over time, 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 produce two results. On the one hand, the financial channel is important *on average*, accounting for about 30% of the overall response of RoW industrial production. When it comes to explaining time-variation in the RoW industrial production response, on the other hand, its role is again found to be limited.

The remainder of the paper proceeds as follows. Section 2 provides an overview of the literature and highlights the contribution of our paper. Section 3 discusses the empirical strategy and specification. In Section 4, we present the empirical results, including various extensions, the heterogeneity analysis, the counterfactual exercises, and a battery of sensitivity checks. Finally, Section 5 concludes.

2. Contribution to the Literature

Our main contribution is to study the time-varying nature of international spillovers of U.S. monetary policy and its drivers. Therefore, we relate to three strands of the literature.

First, we connect to the vast literature documenting the negative effects of U.S. monetary policy tightenings on the global and real financial cycles.³ Specifically, we are 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 (particularly, conditional on different income levels, degree of financial and trade openness, exchange rate regimes, dollar trade invoicing, and gross dollar exposure). On the contrary, we use a time-varying model instead of a linear model and document the increasing international spillovers of U.S. policy shocks.

Second, this paper contributes to the literature exploring the time-varying nature of monetary policy shocks. Cogley and Sargent (2005) and Primiceri (2005) are among the first papers to use time-varying parameter VARs to examine potential time-dependency. A number of contributions use the methodology to explore monetary policy effectiveness (Boivin and Giannoni, 2006; Canova and Gambetti, 2009) or the response of the stock market to monetary policy shocks (Galí and Gambetti, 2015; Paul, 2020; Aastveit et al., 2021). We focus on the time-varying effects of U.S. monetary policy shocks on global cycles, extending the analysis beyond U.S. domestic variables. Our findings underscore the relevance of time-variation in understanding the international ramifications of such disturbances.

Third, we contribute most closely to the strand of literature examining the time-varying dimension of international monetary policy. Specifically, this 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. They find i) an increased connectedness between and within countries in periods of financial distress (see also Canova et al., 2007 and Billio et al., 2016), ii) policy coordination between the three central banks (especially in the 1980s), and iii) larger and more persistent monetary policy disturbances in the Great Moderation than in any subsequent periods. 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. Another contribution by Ilzetzki and Jin (2021) compares the international transmission of U.S. monetary policy shocks prior and after

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.

1990s. They find that, prior to the 1990s, world industrial production declines in response to a U.S. monetary tightening (as predicted by textbook open economy models), 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 favour of a growing (negative) role of U.S. policy tightenings for global economic activity. Cuaresma et al. (2019) estimate a GVAR model identified via sign restrictions and report heterogeneous effects of U.S. policy shocks effects across countries and over time, without finding a clear time pattern in such spillovers.

3. 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 (Database of Global Economic Indicators).⁴ 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. 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 2015M12, which we split in 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 2015M12. 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 2008M12-2015M12.⁵ Unless otherwise noted, we estimate all model specifications with $p = 3$ lags due to the curse of dimensionality inherent in TVP-VAR estimation. Since the state-space of the model grows quadratically with the number of variables

⁴ 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.

⁵ 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 4.6 we show that the our results are similar when excluding the zero lower bound period from the estimation sample.

in the model, we are limited to a small set of variables and/or a small number of lags. 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 state-of-the-art 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).⁶ As shown in Bauer and Swanson (2023a), the construction of this instrument is also robust to the “Fed response to news” channel of Bauer and Swanson (2023b).

3.1 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).⁷

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

$$\mathbf{y}_t = \mathbf{c}_t + \sum_{j=1}^p \mathbf{A}_{jt} \mathbf{y}_{t-j} + \mathbf{B}_t z_t + \mathbf{u}_t, \quad \mathbf{u}_t \sim \mathcal{N}_M(\mathbf{0}, \Sigma), \quad (3.1)$$

where \mathbf{c}_t is a $n \times 1$ vector of time-varying intercepts and \mathbf{A}_{jt} ($j = 1, \dots, p$) are $n \times n$ time-varying coefficient matrices of the lagged endogenous variables. Reduced-form innovations are given by the $n \times 1$ vector \mathbf{u}_t , which follow a multivariate Gaussian distribution with zero mean and covariance matrix Σ . Additionally, the model includes an exogenous variable z_t (the monetary policy instrument) with its respective $n \times 1$ time-varying coefficients \mathbf{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.

Let $\boldsymbol{\theta}_t = \text{vec}(\mathbf{c}_t, \mathbf{A}_{1t}, \dots, \mathbf{A}_{pt}, \mathbf{B}_t)$ be a vector that stacks all coefficients on the right-hand side of Equation 3.1. The dynamics of the model’s time-varying parameter is specified as follows

$$\boldsymbol{\theta}_t = \boldsymbol{\theta}_{t-1} + \boldsymbol{\nu}_t, \quad \boldsymbol{\nu}_t \sim \mathcal{N}(\mathbf{0}, \mathbf{Q}), \quad (3.2)$$

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 3.1) and the state equation (Equation 3.2) 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

⁶ The instrument of Miranda-Agrippino and Ricco (2021) is constructed exploiting Greenbook forecasts, which are released to the public with a significant lag. The instrument is available up until 2015M12 (Degasperi and Ricco, 2021). This is the reason why our sample ends in that period.

⁷ Our sample starts in the so-called Great Moderation. Hence, contrary to Primiceri (2005), we do not allow for stochastic volatility in the residuals in our benchmark estimations. We show in the robustness checks that the results are not sensitive to this choice.

inverse-Wishart for Σ and \mathbf{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), we plug in zeros for the observations prior to 1991M1 (i.e., during the pre-sample period from 1980M1 to 1991M1, we always observe the monetary instrument during the estimation sample). The hyperparameter governing the prior belief on the amount of time-variation in θ 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.

3.2 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 \mathbf{u}_t via a linear mapping $\mathbf{u}_t = \mathbf{S}_t \boldsymbol{\varepsilon}_t$, where \mathbf{S}_t collects the contemporaneous impulse matrix in t . We are interested in the effects of U.S. monetary policy and thus in identifying one particular column of the matrix \mathbf{S}_t , which we denote by s_t . Without loss of generality, we assume that the impulse vector 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 the shock via

$$z_t = \varphi \varepsilon_{1t} + \zeta_t, \quad \zeta_t \sim \mathcal{N}(0, \sigma_\zeta^2), \quad (3.3)$$

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.⁸ In an additional exercise, we also investigate the ratios of two impulse response functions. In this case, we only assume that in each time period t the system is hit by the same shock.

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

$$r_{t,i1} = \frac{s_{t,i}}{s_{t,1}} = \frac{B_{t,i}}{B_{t,1}}. \quad (3.4)$$

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.⁹ The posterior

⁸ This assumption has been recently tested by Amir-Ahmadi et al. (2023). They show that a few distinct periods are informative about monetary policy shocks. However, using a time-varying specification they estimate even larger effects, pointing to a downward bias in the linear specification. We view our estimates thus as a conservative benchmark.

⁹ 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

quantities of A_{jt} ($j = 1, \dots, p$; $t = 1, \dots, T$) 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 one percentage point 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 3.3, see Paul, 2020).

4. 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.

4.1 Evidence from the Linear Specification

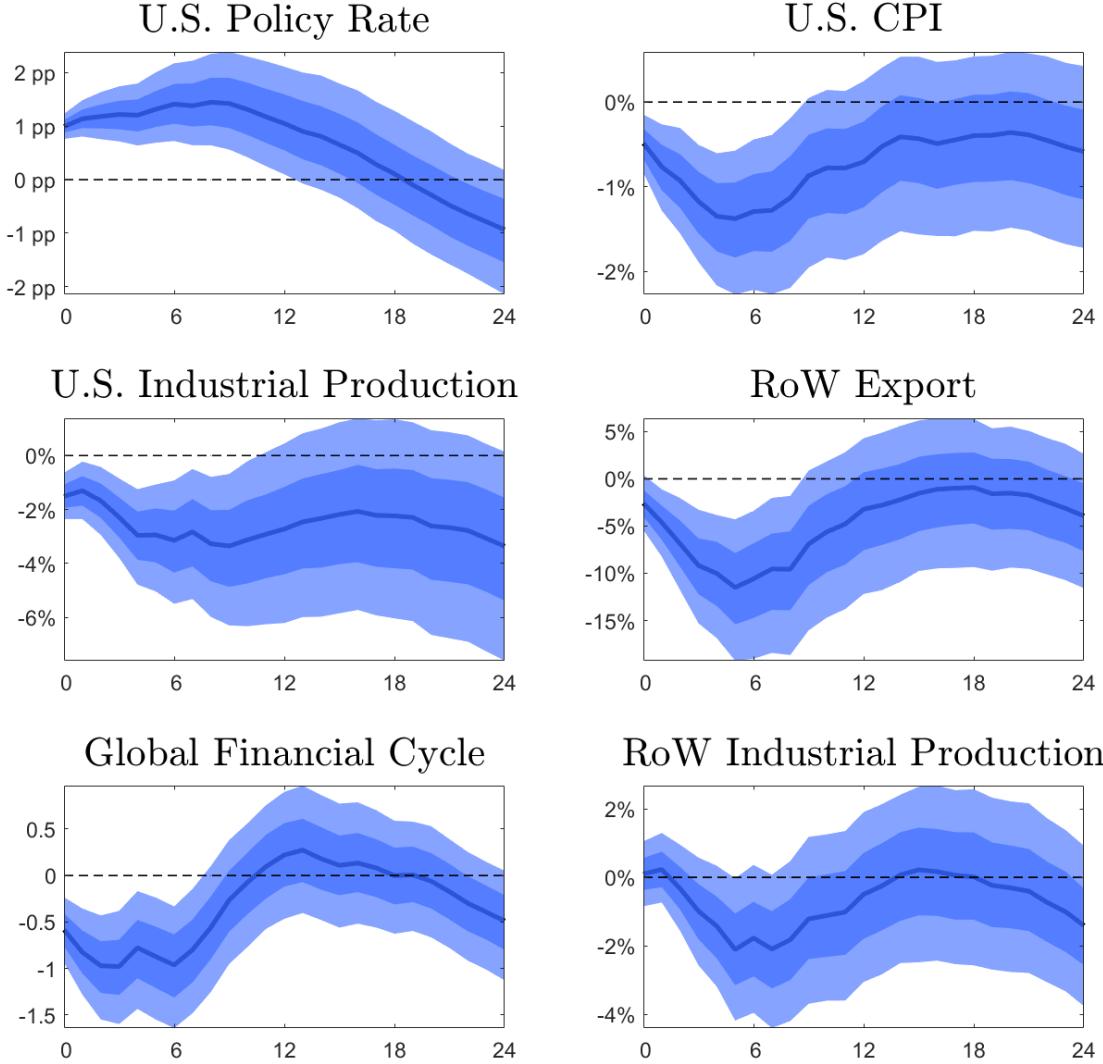
We first estimate a linear VAR considering the variables described before.¹⁰ We test the relevance of the policy instrument in the linear VAR and we find a first-stage F-statistic of 23.9, which is safely above the standard threshold of 10 (see for instance Montiel Olea et al., 2021).

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 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. For instance, Miranda-Agrrippino and Ricco (2021) (we) find a peak contraction of -2.6% (-3.3%) for industrial production and -1% (-1.4%) for consumer prices to a one percentage point shock. 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 one year (see e.g. Miranda-Agrrippino 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

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.

¹⁰ 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 as in (Paul, 2020). Since a pre-sample period is not needed, we constrain the sample by the availability of the monetary policy instrument (1991M1-2015M12).

Figure 2: Linear Impulse Responses Functions.



Notes: Responses to a one percentage point (1 pp = 100 basis point) contractionary monetary policy shock. Median response and 68% and 95% confidence intervals are reported (wild bootstrap; 2,000 samples).

and heightened financial stress give rise to a worldwide economic downturn, quantified by a -2% 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.

4.2 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.

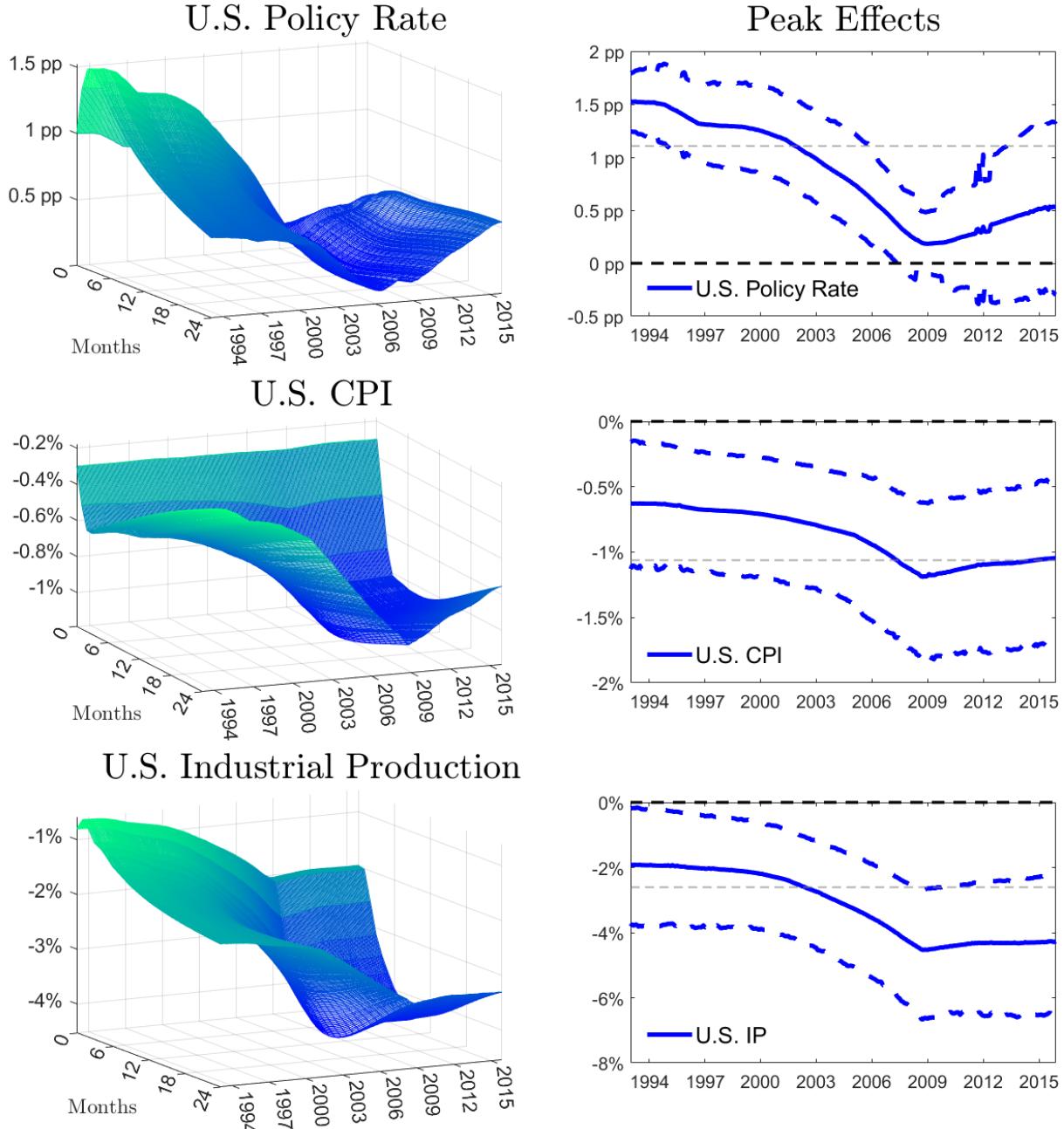
To facilitate interpretation, we show the results in two ways. First, we report the evolution over time of the median impulse-response functions in the left column. Second, we display the peak response of the variables (together with its 68% credibility sets) in the right column. For comparison, we also 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). We then follow Paul (2020) and obtain same-size shocks for all subsequent periods.¹¹

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 that take into account agents' expectations about the path of the policy rate, 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).¹² Second, time-dependent patterns arise in the responses of U.S. aggregates. Peak contractions in domestic industrial production aggravate over time (rising from -2.0% in 1993 to -4.1% 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 -0.7% , which grows in magnitude to -1.2% until the last financial crisis.

¹¹ Consistently, we use the same variation in z_t underlying the TVP-VAR shocks to normalize the linear VAR's responses, which delivers a 0.75 percentage point increase in the policy rate within the linear specification.

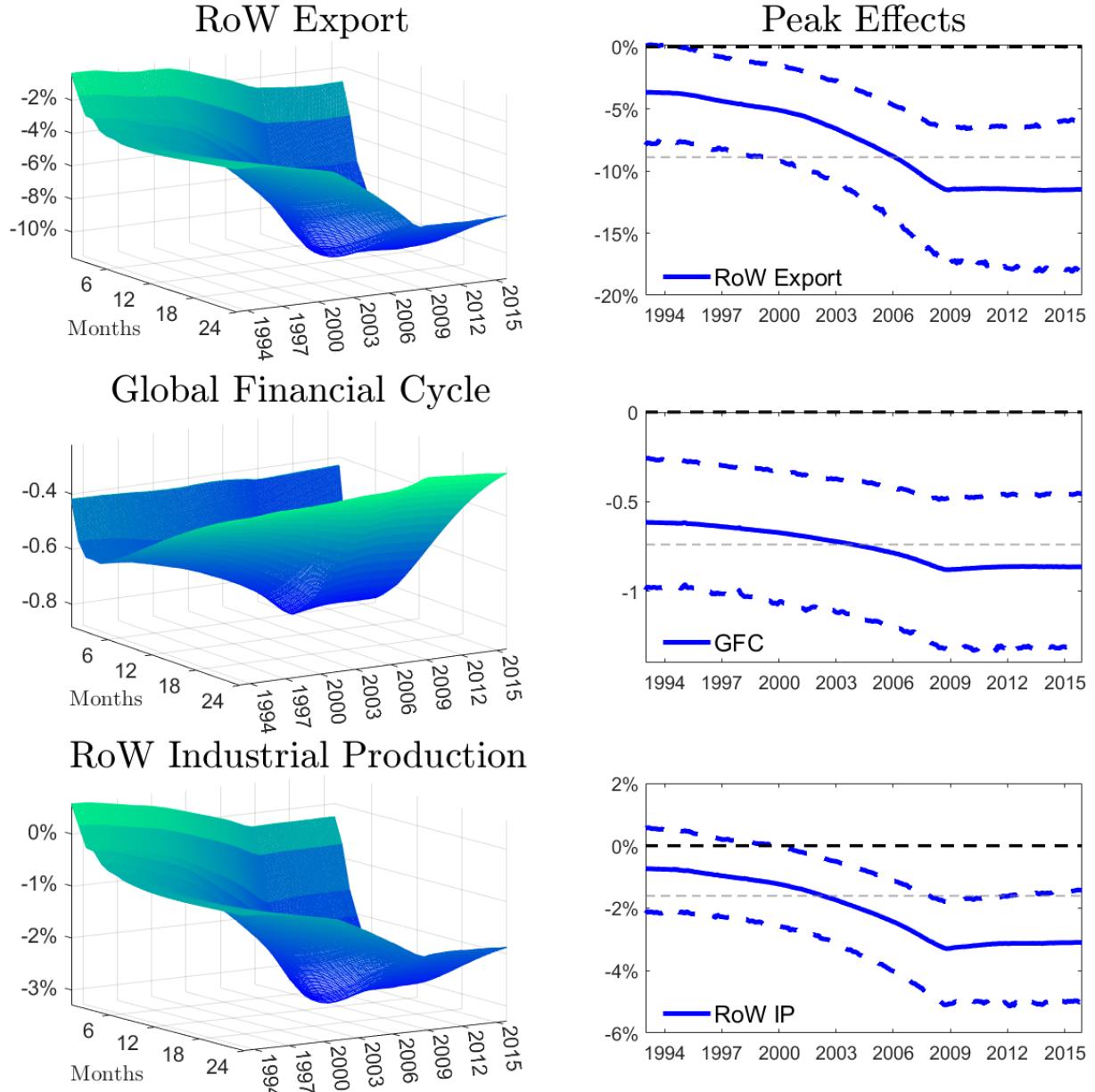
¹² Figure C1 reports the evolution of the long-run trend along with the on impact response of the U.S. policy rate underlying our specification. Accordingly, the long-run trend is relatively stable in 1990s, but starts decreasing in the 2000s, and becomes even negative during the zero lower bound period. See Appendix C for more details.

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 point) 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 0.75 pp increase in the policy rate in the linear specification).

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 point) 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 0.75 pp increase in the policy rate in the linear specification).

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 (negative) 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 -0.7% in 1993 to -3.4% in 2008M9 (Lehman Brothers' bankruptcy).¹³ 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.¹⁴ 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 (in relative terms) of approximately 40% (from -0.64 to -0.89 in terms of peak effects).

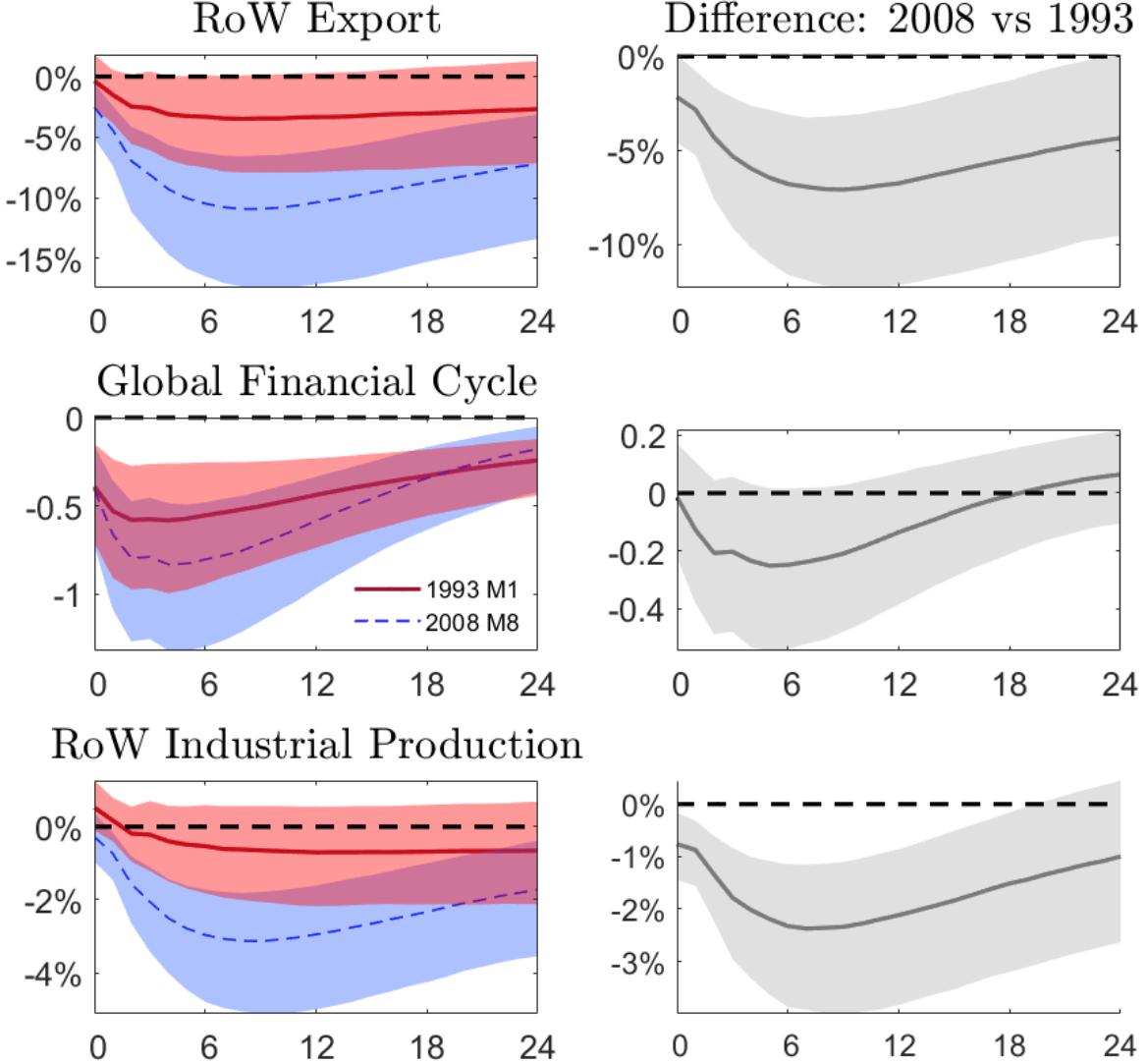
Although the magnitude of the global effects has been growing since the beginning of the sample, the pace of this growth notably accelerates in the early 2000s. This intensification of international spillovers coincides precisely with a period characterized by factors such as the trade boom, relaxed financial regulation and supervision of banks (Shin, 2012), and a sharp rise in 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 2007-09 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 domestic effects due to U.S. monetary policy shocks can also result from spillover-spillback loop effects. This means that international spillovers spillback to the domestic economy and affect economic aggregates. Evidence for an active spillback mechanism is provided by Breitenlechner et al. (2022), who find that spillbacks account for nearly 50% of the overall effect of U.S. monetary policy on domestic real activity using counterfactual simulations. If we apply this “rule of thumb” in our analysis, we find that the increasing effects on RoW industrial production

¹³ Given that the shocks are normalized to have the same magnitude, the effects 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.

¹⁴ 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 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. 2008M9. 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).

(from -0.7% to -3.4% , i.e., +390% in relative terms, implying a predicted increase in domestic effects of approximately +195%) fully explain the variation in the response of U.S. industrial production (from -2.0% in 1993 to -4.1% , i.e., +200% in relative terms).

Finally, we investigate whether time-variation is statistically significant by focusing on particular episodes in the sample. We consider the periods 1993M1 and 2008M9. 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). Due to the proliferation of parameters and latent states, uncertainty bounds are quite sizable. Nonetheless, some noteworthy insights emerge. The primary distinction

between the two examined periods lies in the effects on RoW industrial production. On the one hand we find rather weak 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.

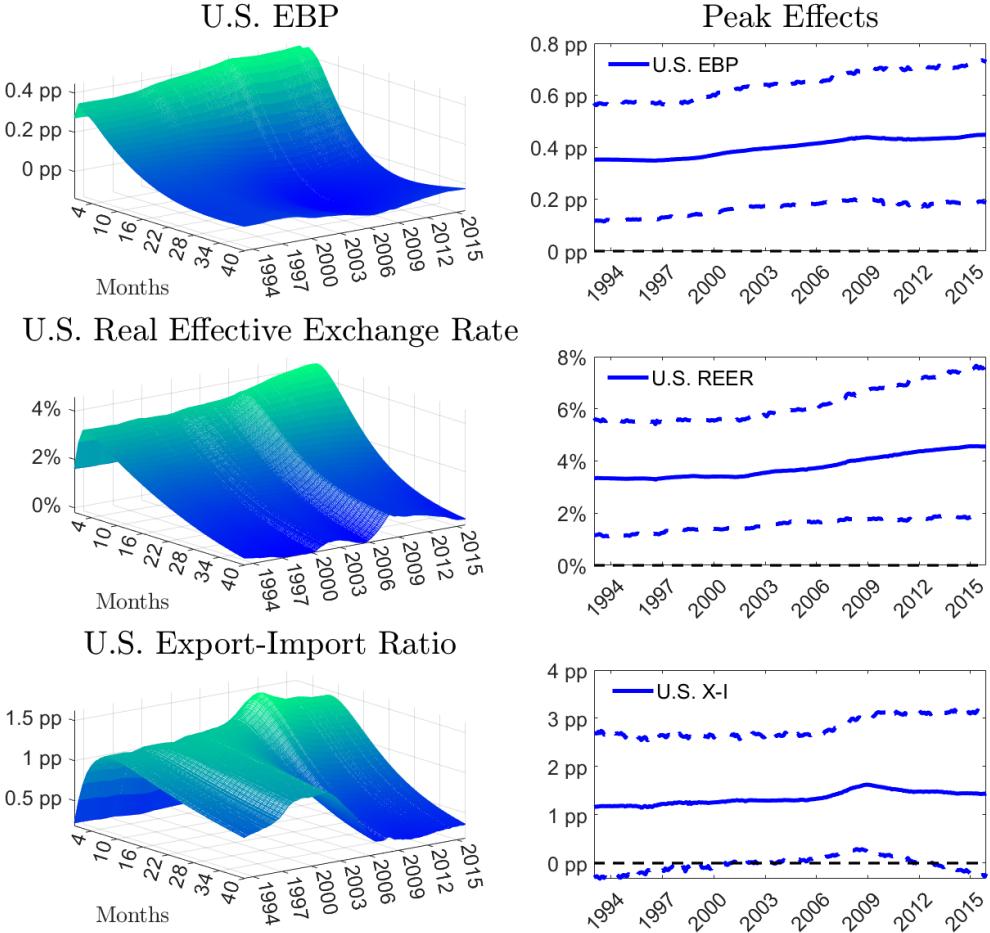
4.3 Wider Propagation Channels

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 $n = 7$ variables. Since the state-space would become too large to estimate sensible results, we reduce the number of lags to $p = 2$.¹⁵ Estimation and prior specification are kept unchanged (we refer to Appendix A for the exact variable definitions and transformations).

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 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 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., 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.

¹⁵ We show in Figure E1 that, in these alternative specifications, the responses of the variables considered so far are extremely similar to the ones in our benchmark estimations.

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



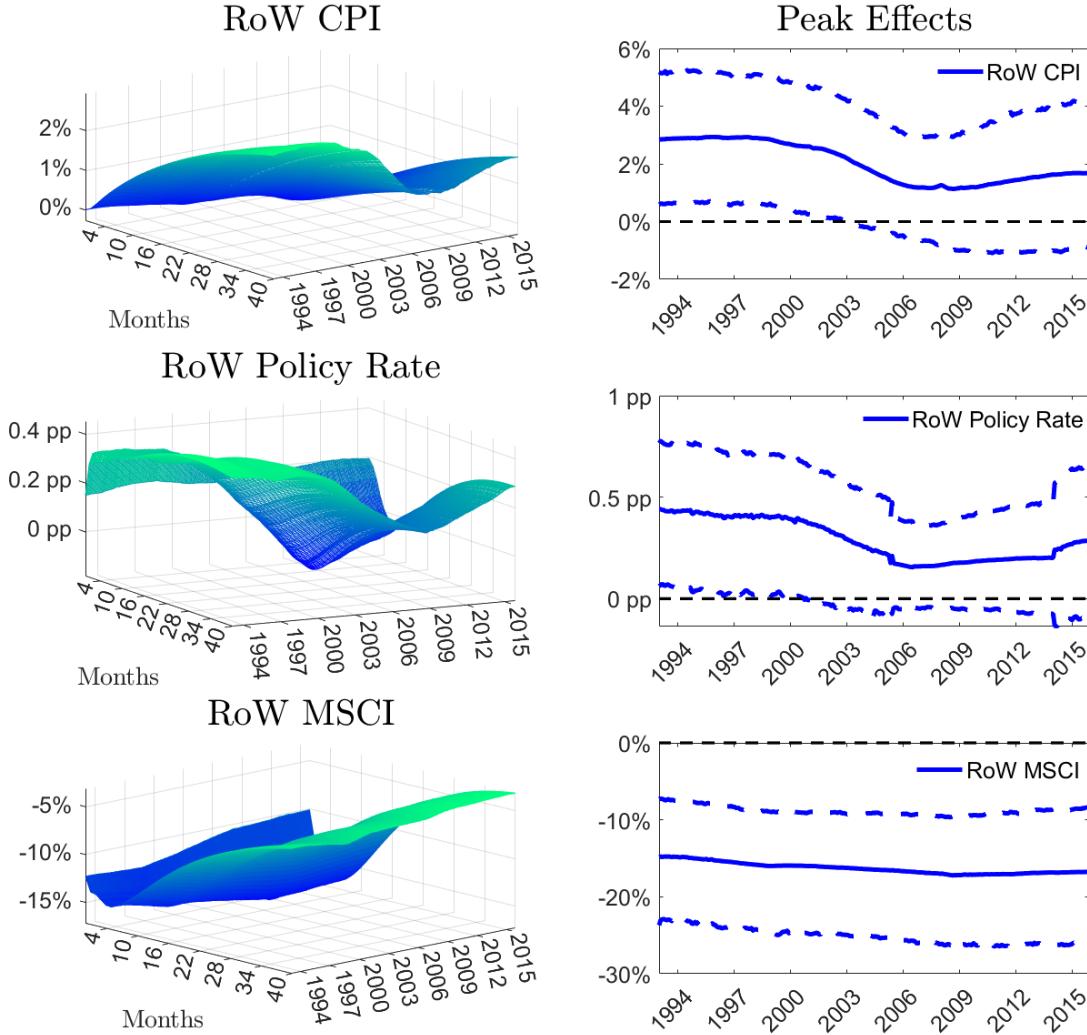
Notes: Responses of additional variables to a contractionary U.S. monetary policy shock that induces a one percentage point (1 pp, 100 basis point) increase in the federal fund rate in 1993M1. Left column reports the evolution over time of the median impulse-response functions. The right column reports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

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 persistent 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 +2.8% to +1.3% at peak), which can be attributed to the rise of global value chain participation (Georgiadis et al., 2019).¹⁶ 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 advanced economies.¹⁷ The gradual decrease in the policy rate

¹⁶ Georgiadis et al. (2019) argue that there is a causal link between the rise of global value chain participation and the decline of the exchange rate pass-through in the last decades.

¹⁷ 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 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 point) 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.

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.

4.4 Heterogeneity Analysis

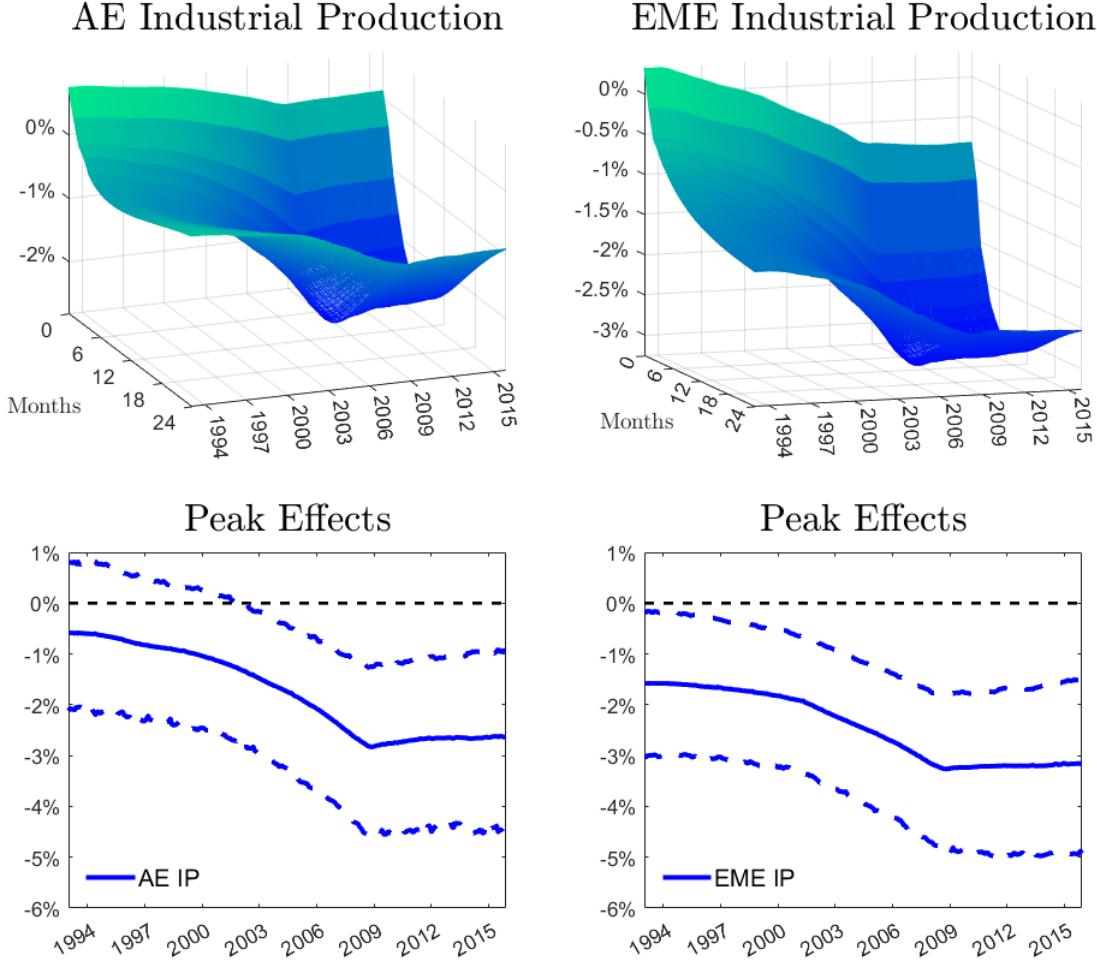
We have documented a significant time-variation in the international spillovers of U.S. monetary policy. In this and the following section, we conduct additional exercises to explain the findings and link it to different channels. As highlighted by Kalemli-Özcan, 2019, De Leo et al. (2022), and Ca'Zorzi 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.¹⁸ The results, shown in Figure 8, reveal more pronounced recessionary effects in emerging economies - consistent with existing studies. The peak effect on emerging (advanced) countries' industrial production is -1.8% (-0.6%) in 1993 and -3.3% (-2.8%) in 2008. In those years, the peak response of RoW industrial production, which combines both emerging and advanced economies, is -0.7% and -3.4% , 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. However, uncertainty bands are quite wide and encompass possible estimation outcomes. 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

¹⁸ 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.



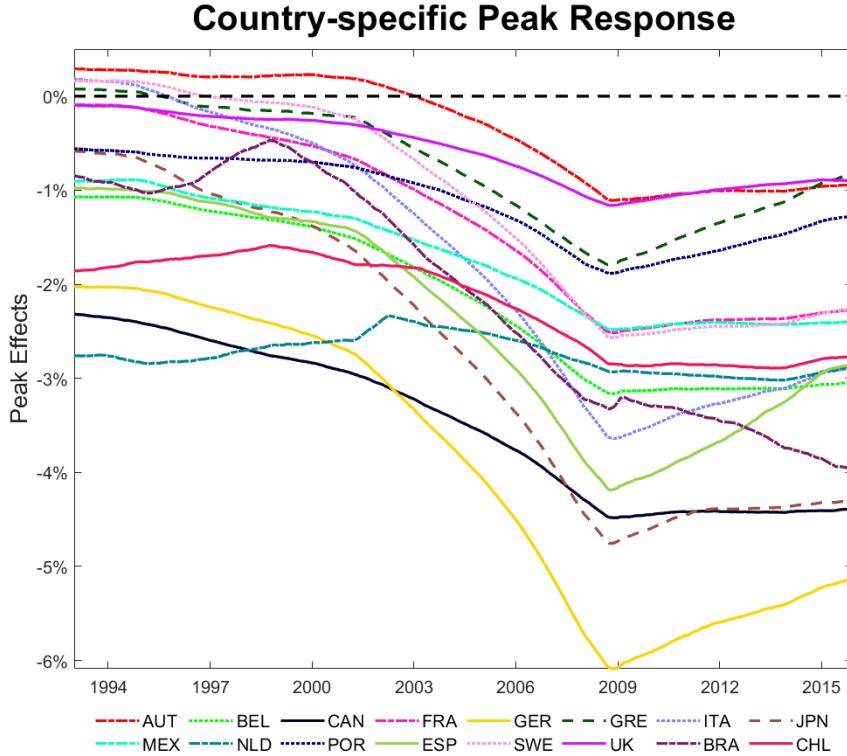
Notes: Responses to a contractionary U.S. monetary policy shock that induces a one percentage point (pp) increase in the federal fund rate in 1993M1. Upper panel reports the evolution over time of the median impulse-response functions. The lower panel reports the peak effects over time for each variable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

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 16 countries in our analysis. These countries were selected based on two criteria: i) they were included in the aggregate RoW production measure of Federal Reserve Bank of Dallas and, ii) for which monthly data is available from 1980 onwards. Our sample includes: Austria, Belgium, Brazil, Canada, Chile, France, Germany, Greece, Italy, Japan, Mexico, Netherlands, Portugal, Spain, Sweden, UK.¹⁹

We report the evolution of country-specific peak responses in Figure 9, while the median time-varying impulse response functions for each country can be found in Appendix E (Figure E2-

¹⁹ Data for most countries are obtained from the OECD's website, except for Brazil and Chile, where data is retrieved from the IMF. Consistently with the aggregate index, we utilize seasonally adjusted industrial production measures that exclude the construction sector.

Figure 9: Peak Effects of Country-Specific Industrial Production.



Notes: Peak 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 1993M.

Figure E4). 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 all countries considered. These effects tend to intensify over time until the occurrence of 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 Germany, historically exhibit a greater susceptibility to U.S. shocks, whereas others like the UK and Austria are less affected. This dimension can be summarized by taking the average peak effect in country i over time. 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 Italy, while it is more limited in Chile, Austria, and the Netherlands.

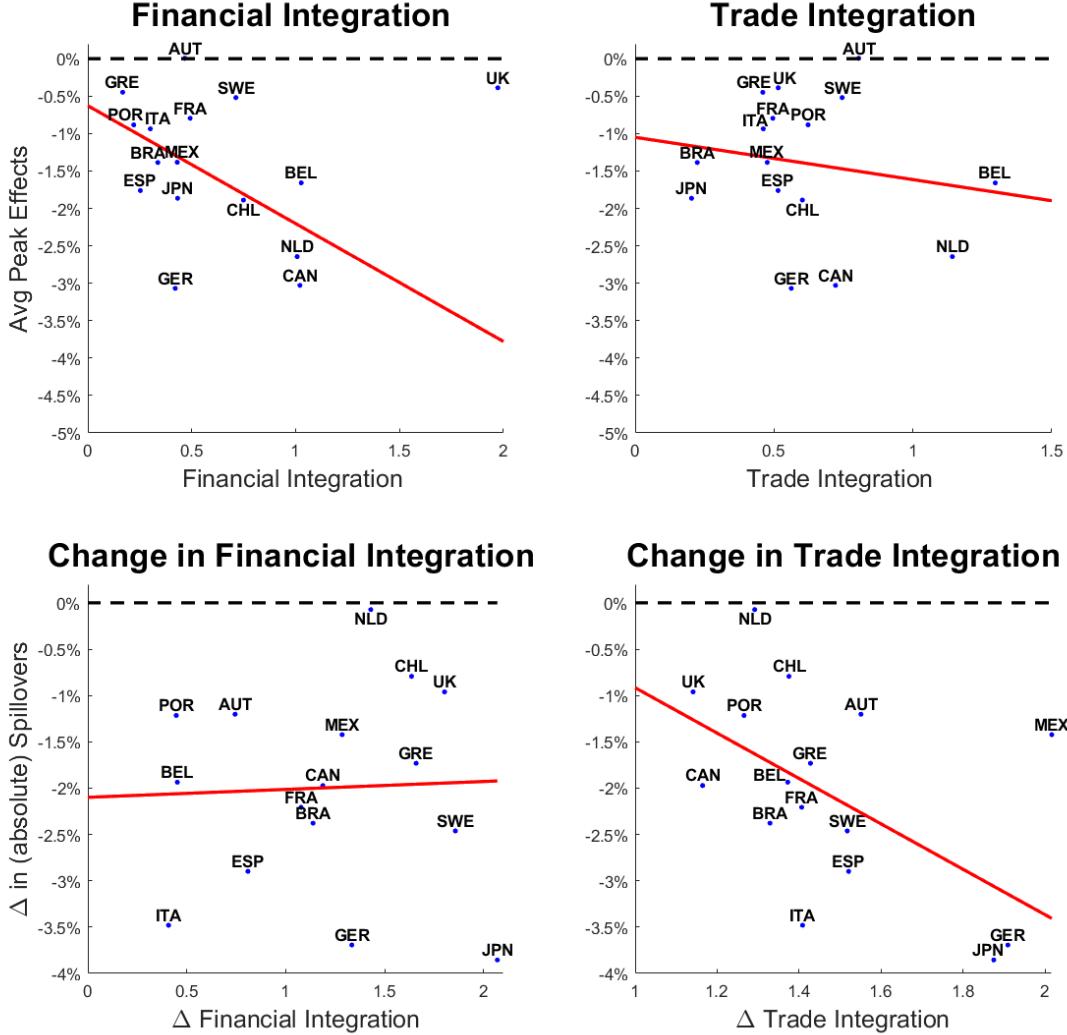
4.5 Evaluation of the Channels

Our empirical analysis again points towards significant time patterns in global spillovers. But what are the main drivers of such dynamics? We try to 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 similar choices in this regard, see Ca'Zorzi et al., 2023).²⁰ Financial and trade integration are then computed 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 ratio between integration in 2007 (right before the Great Recession) and integration in 1993 (first year in our sample), which encompasses the time period where time-variation manifests.

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 *variation over time*). In the upper panel, we report the correlation of average peak effects with average integration, while in the lower panels we show the correlation of variation in peak 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 correlation coefficient for this negative relationship is -0.48 , and it proves statistically significant at conventional confidence levels when conducting one-sided correlation tests ($p\text{-value} = 0.03$). Conversely, there is no relationship between changes in financial integration and changes in recessionary effects (the correlation is equal to 0.02 ; $p\text{-value} = 0.47$). 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. While we acknowledge the inherent reduced-form nature of this analysis, we believe

²⁰We retrieve the data for financial integration from Bénétrix et al. (2015). Data for trade integration is instead taken 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 consider the overall trade to GDP ratio (irrespective of currency) as a proxy for overall trade integration.

Figure 10: Correlations of Country-Specific Peak Responses, Financial and Trade Integration.



Notes: Upper panels: Relationship between the average peak 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 peak response (the difference between the peak response in 2008 and 1993) and the corresponding change in financial and trade integration (the ratio of integration in 2007 to 1993; we consider 2007 here not to consider the massive movements in trade and financial integration occurred in 2008). Red lines: univariate regressions interpolating the points (weighted least squares robust to outliers).

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 is in line 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 Zakrašek (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 affected with a lag), but before the U.S. policy rate (fast-moving variable, which is allowed to react contemporaneously).²¹ By doing so, we assume that the *financial multiplier* is zero on impact, but we allow it to change over time in all subsequent horizons. We keep everything else as in the baseline estimations.

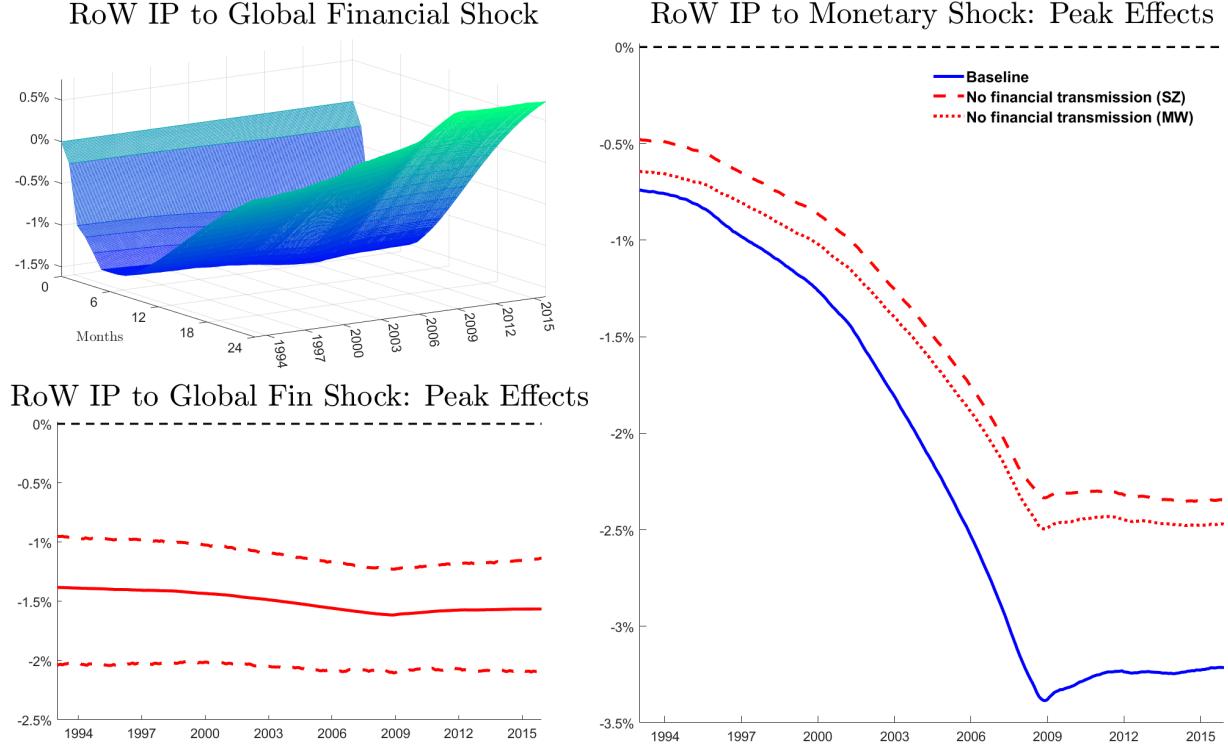
Figure 11 illustrates that the recessionary effects of this shock on RoW industrial production have remained constant over time (plots on the left). Digging deeper, we conduct counterfactual simulations aimed at shutting down the financial transmission of U.S. monetary policy. We eliminate the impact of U.S. disturbances on the global financial cycle by generating hypothetical *global financial shocks* (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 using both the approach outlined in Sims and Zha (2006) and the recent refinement of McKay and Wolf (2023), which deliver similar results.²² Our findings reveal that, on average, the financial channel is relevant, contributing to about 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. This is supported by a battery of different results. First, the increasing response of RoW export (as opposed to a constant response of global financial conditions), that signals a more pronounced real transmission through trade channels. Second, we show the potential

²¹ 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.

²² In the Sims and Zha (2006) (SZ) counterfactual, we generate a sequence of global financial shocks that completely offset the impact of the U.S. policy shocks on global financial cycle (one for each t in the sample). The McKay and Wolf (2023) (MW) approach involves subjecting the economy to financial shocks only at horizon $t = 0$ in order to closely approximate the counterfactual scenario. Notice that the approach of McKay and Wolf (2023) would ideally require the identification of multiple global financial shocks. This proves difficult in our application. We hence see our exercise with one shock only as a further robustness check.

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 (SZ) and McKay and Wolf, 2023 (MW)).

relevance for spillovers of the economic rise of emerging economies via composition effects. Third, cross-sectional correlation analysis points toward 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.

4.6 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. While the reduced sample generates a slightly more unstable inference, 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: $\Sigma_t = \mathbf{A}_t^{-1} \text{diag}(\exp(h_{1t}), \dots, \exp(h_{nt})) \mathbf{A}_t^{-1'}$. The matrix \mathbf{A}^{-1} is non-triangular and the free elements follow a random-walk. Similarly, the log-volatilities in h_{it} ($i = 1, \dots, n$) also follow a random walk. We estimate the stochastic volatility model as in Kim et al. (1998) but with the refinement of Omori et al. (2007). 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 \mathbf{A}_t^{-1} and h_{it} ($i = 1, \dots, n$) must be specified. Using the notation of Primiceri (2005), this boils down to a selection choice on three parameters: κ_Q (governing time variation of autoregressive coefficients), κ_W (variance of the residuals), κ_S (covariance of the residuals). We set $\kappa_Q = 0.01$, $\kappa_W = 0.001$, and $\kappa_S = 0.001$. The values of κ_Q and κ_W are among the ones considered by Primiceri (2005), while our κ_S is relatively tighter (to avoid ill behaviors in our relatively shorter sample). The results of this exercise are presented in Figure F2. If we compare the global responses to the baseline estimates, the qualitative pattern is the same. 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 κ_Q^2 , which regulates the time-variation in the reduced-form VAR coefficients (see the exact prior specification in Appendix B). 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 ($\kappa_Q^2 = 0.02$).

5. Concluding Remarks

We study whether and how the international effects of U.S. monetary policy shocks have changed over the last decades. While existing studies have largely abstracted from this time dimension by employing linear models, we provide evidence in support of growing global spillovers of U.S. policy disturbances.

To do so, we estimate a TVP-VAR which features U.S. and global indicators. Identification is achieved by means of state-of-the-art methods which exploit high-frequency external instruments. We integrate the policy instrument of Miranda-Agricappino and Ricco (2021) in the TVP-VAR as an exogenous variable, following the methodology proposed by Paul (2020). 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 as well as 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.

Specifically, our results reveal that the impact of a U.S. tightening on global industrial production has more than tripled over the last three 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 on 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 both 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 is likely to entail growing spillbacks. We indeed find that domestic effects of U.S. policy shocks have strengthened over time. Drawing on existing estimations by Breitenlechner et al. (2022), we argue that such dynamics could be a consequence of the growing spillovers abroad, which, in turn, translate into amplified spillbacks domestically.

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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 2015M12. 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	FFR_t	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}	$100 \times [\ln(CPI_t^{US}) - \ln(CPI_{t-1}^{US})]$	CPI_t^{US} is U.S. consumer price index (Fred).
ip_t^{US}	$100 \times [\ln(IP_t^{US}) - \ln(IP_{t-1}^{US})]$	IP_t^{US} is U. S. industrial production (Fred).
xim_t^{US}	XIM_t^{US}	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}	EBP_t^{US}	EBP_t^{US} is the U.S. excess bond premium by Gilchrist and Zakrajsek (2012).
gfc_t	GFC_t	GFC_t is the global financial cycle indicator by Miranda-Agrippino and Rey (2020).
$reer_t$	$100 \times \ln(REER_t^{US})$	$REER_t^{US}$ is the U.S. real effective exchange rate from the BIS.
cpi_t^{RoW}	$100 \times [\ln(CPI_t^{RoW}) - \ln(CPI_{t-1}^{RoW})]$	CPI_t^{RoW} is RoW consumer prices (excluding the U.S.) Grossman et al. (2014).
ip_t^{RoW}	$100 \times [\ln(IP_t^{RoW}) - \ln(IP_{t-1}^{RoW})]$	IP_t^{RoW} is the RoW industrial production (excluding the U.S.) by Grossman et al. (2014).
$msci_t^{RoW}$	$100 \times [\ln(MSCI_t^{RoW}) - \ln(MSCI_{t-1}^{RoW})]$	$MSCI_t^{RoW}$ is RoW MSCI (excluding the U.S.).
ex_t^{RoW}	$100 \times [\ln(EX_t^{RoW}) - \ln(EX_{t-1}^{RoW})]$	EX_t^{RoW} is RoW exports (excluding the U.S.) by Grossman et al. (2014).
ip_t^{AE}	$100 \times [\ln(IP_t^{AE}) - \ln(IP_{t-1}^{AE})]$	IP_t^{AE} is the advanced economies' industrial production (excluding the U.S.) by Grossman et al. (2014).
ex_t^{AE}	$100 \times [\ln(EX_t^{AE}) - \ln(EX_{t-1}^{AE})]$	EX_t^{AE} is AE exports (excluding the U.S.) by Grossman et al. (2014).
ip_t^{EME}	$100 \times [\ln(IP_t^{EME}) - \ln(IP_{t-1}^{EME})]$	IP_t^{EME} is the emerging market economies' industrial production by Grossman et al. (2014).*
ex_t^{EME}	$100 \times [\ln(EX_t^{EME}) - \ln(EX_{t-1}^{EME})]$	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 2015M12. Exceptions is: EME IP_t 1987M1 to 2015M12.

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 θ , which we denote with θ_0 . For all these initial values, we specify Gaussian distributions. We also have to specify prior densities for the covariance matrices Q , where we use an inverse-Wishart prior density.²³ Last, we also need a prior distribution for the covariance matrix of the VAR Σ , 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

$$\begin{aligned}\theta_0 &\sim \mathcal{N}(\hat{\theta}_{OLS}, 4 * V(\hat{\theta}_{OLS})), \\ Q &\sim iW(\kappa_Q^2 * \tau * V(\hat{\theta}_{OLS}, \tau)),\end{aligned}\tag{B.1}$$

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 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. 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 Σ , which is defined as follows

$$\Sigma \sim iW(\mathbf{I}_M, M + 1),$$

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

²³We 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 3.3, 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 3.1 in its companion form

$$z_t = \mu_t + \alpha_t z_{t-1} + \eta_t, \\ z_t := \begin{bmatrix} y_t \\ y_{t-1} \\ \vdots \\ y_{t-p+1} \end{bmatrix}, \quad \mu_t := \begin{bmatrix} c_t \\ \mathbf{0} \\ \vdots \\ \mathbf{0} \end{bmatrix}, \quad \alpha_t := \begin{bmatrix} A_{1t} & A_{2t} & \dots & A_{pt} \\ I_n & \mathbf{0} & \dots & \mathbf{0} \\ \vdots & \ddots & \ddots & \vdots \\ \mathbf{0} & \dots & I_n & \mathbf{0} \end{bmatrix}, \quad \eta_t := \begin{bmatrix} u_t \\ \mathbf{0} \\ \vdots \\ \mathbf{0} \end{bmatrix}. \quad (\text{C.1})$$

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

$$y_t = (I_n - \alpha_t)^{-1} \mu_t + \sum_{h=0}^{\infty} \alpha_t^h \eta_{t-j} + o_p(1). \quad (\text{C.2})$$

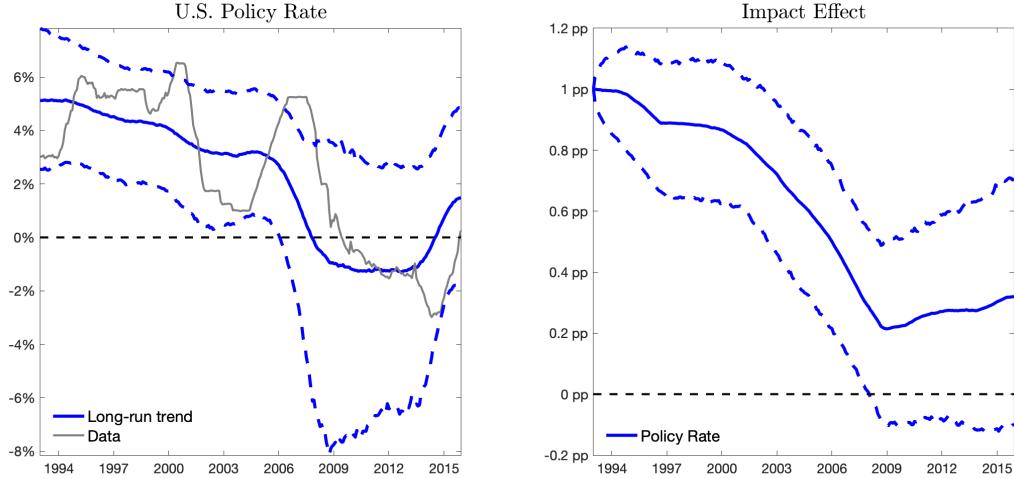
We use this approximation to compute the implied trends of the model's variables:

$$\tau_t = (I_n - \alpha_t)^{-1} \mu_t, \quad (\text{C.3})$$

where the elements of τ_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.

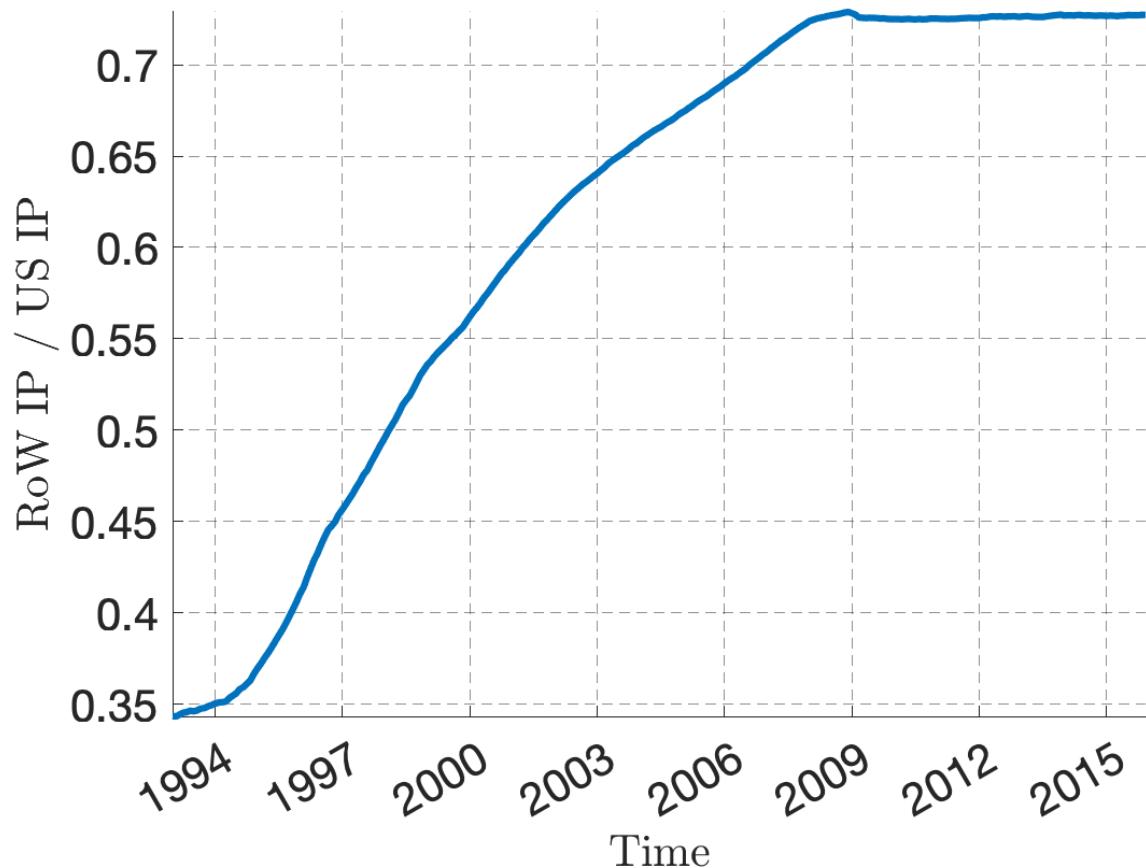


Notes: The left plot shows the actual U.S. policy rate along with the estimate of its long-run trend (unconditional mean). Bands depict the 68% credible intervals. The right plot shows the on impact response of the U.S. policy rate along with its 68% credible sets.

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.

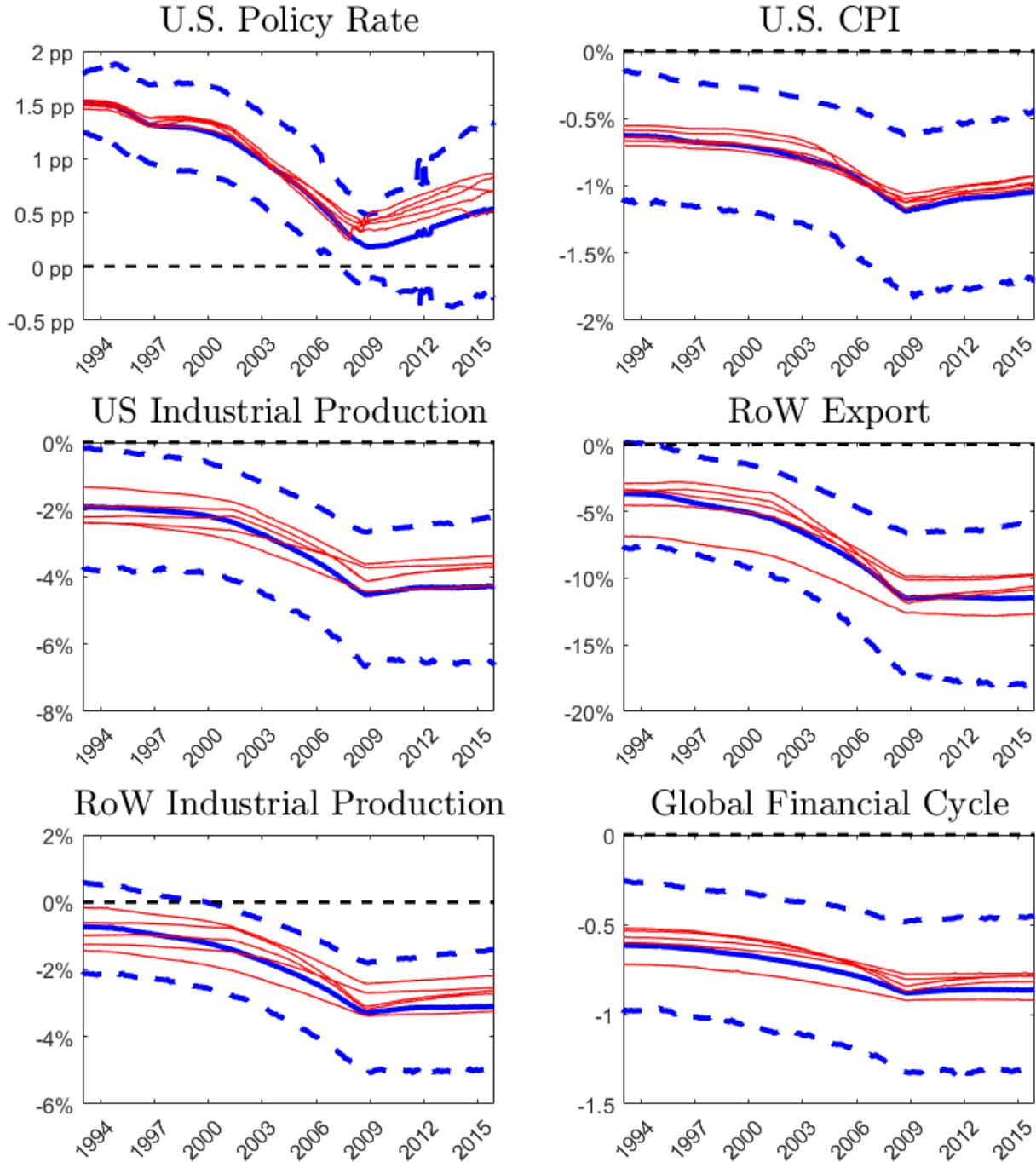
D. Additional Results of the Baseline Specification

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



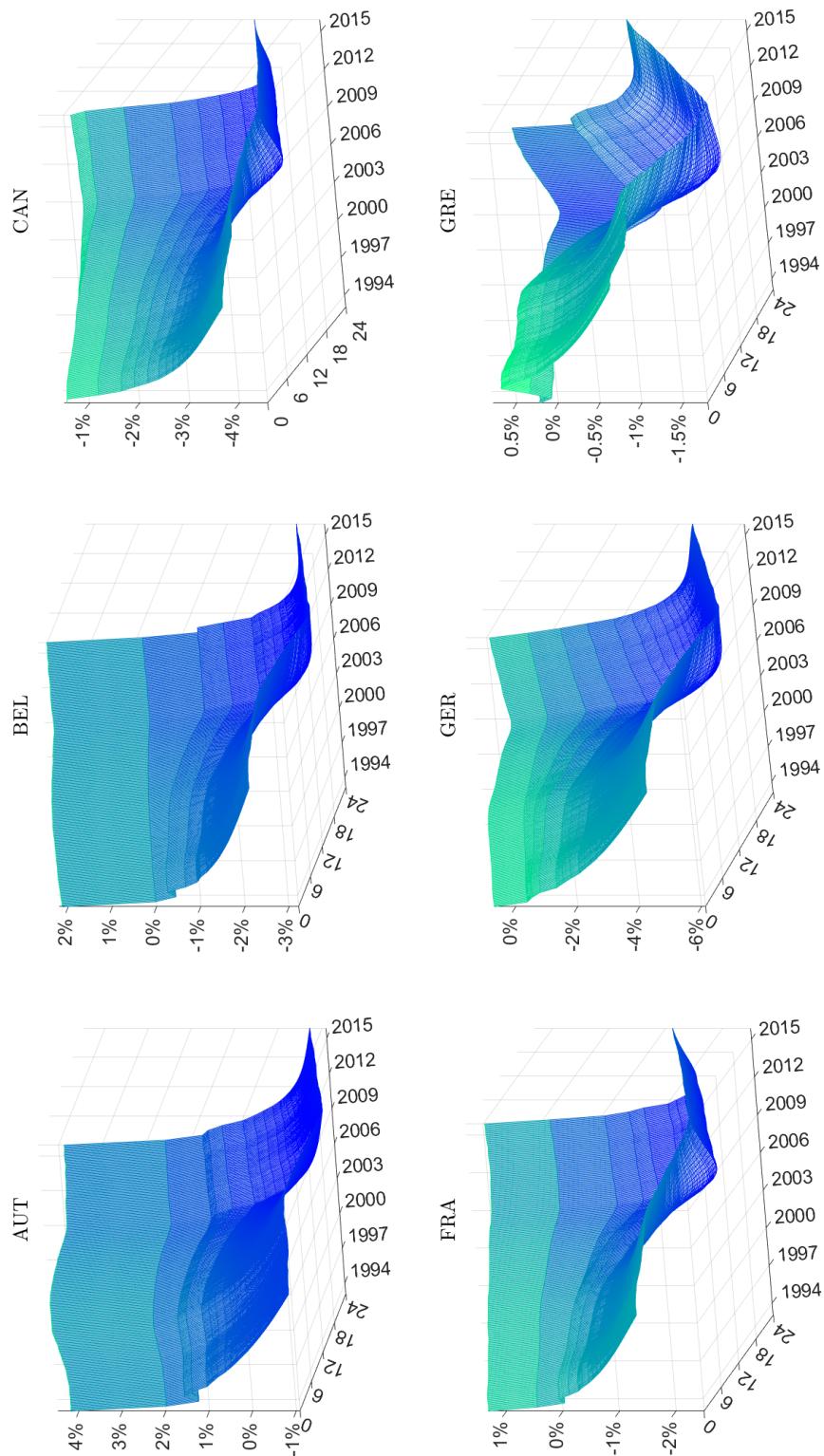
E. Additional Results of the Extended Specification

Figure E1: Time-Varying Impulse Responses Functions of Baseline Variables in Extended Specification.



Notes: Responses of variables in the baseline model to a contractionary U.S. monetary policy shock that induces a one percentage point (pp) increase in the federal fund rate in 1993M1. The plots report peak effects over time for each variable with 68% posterior credible sets in blue. In red are the mean responses of the baseline variables in the five extended specifications.

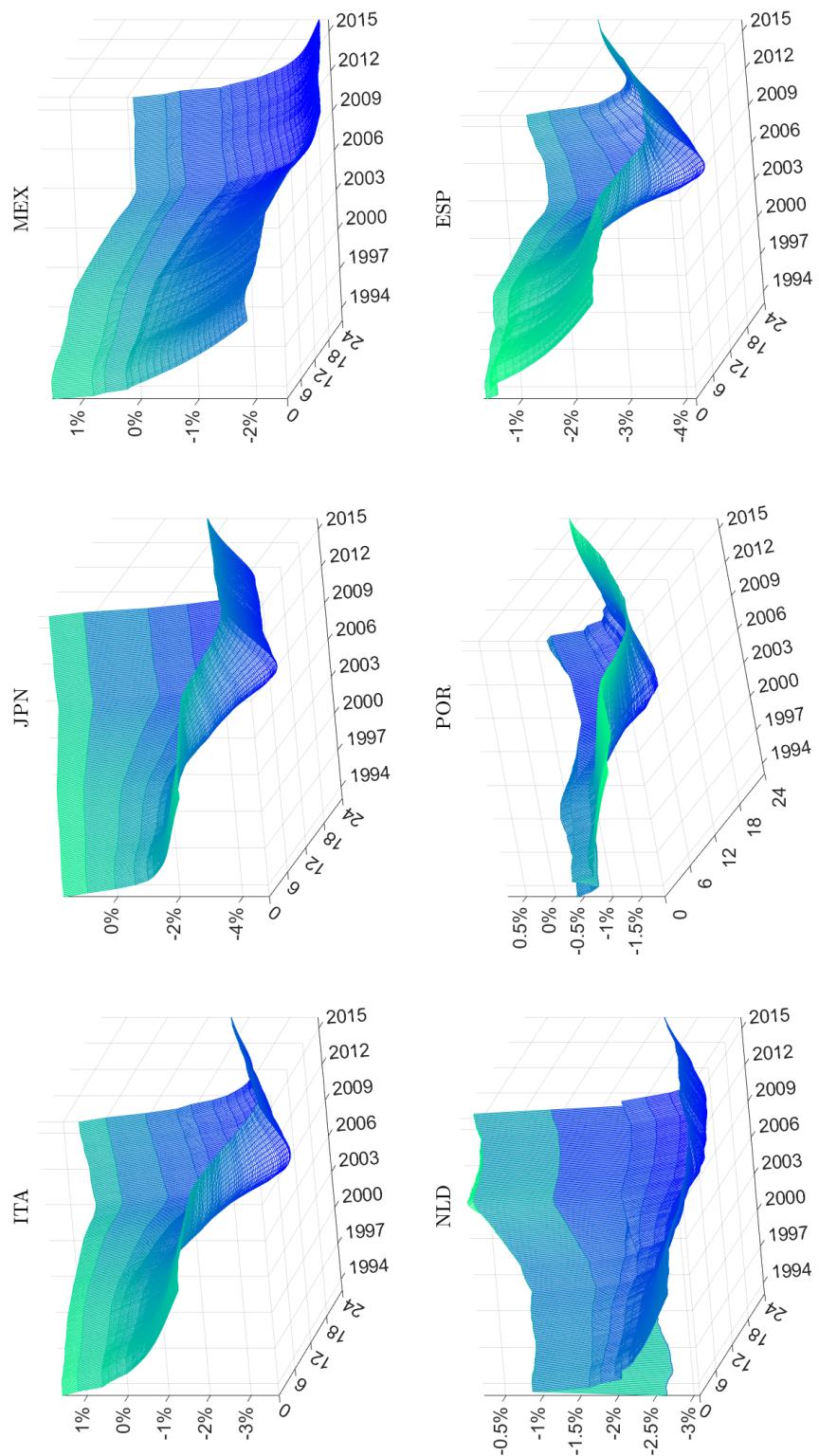
Figure E2: Country-Specific Industrial Production Impulse Responses.



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: Austria, Belgium, Canada, France, Germany, and Greece.

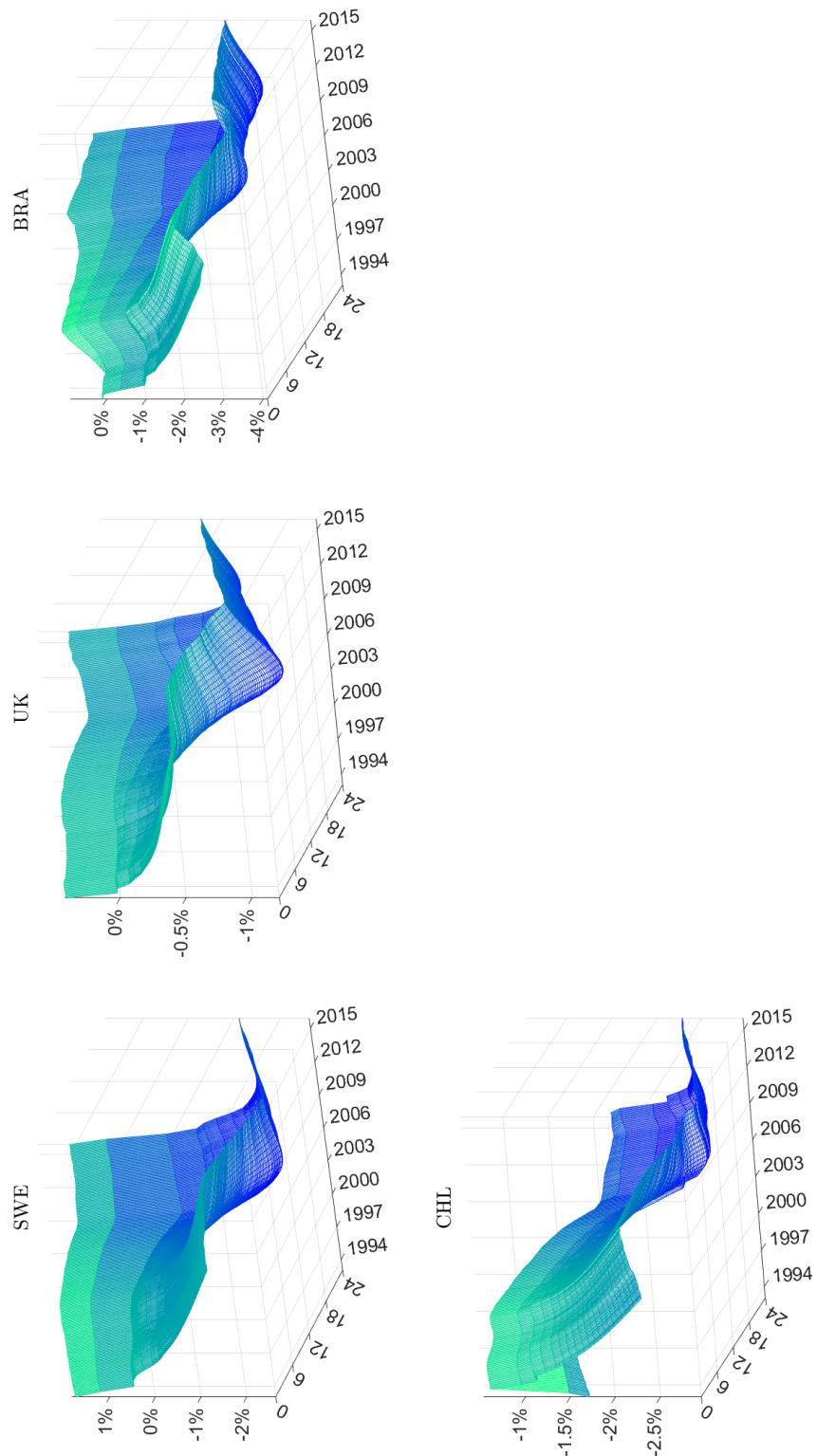
Figure E3: Country-Specific Industrial Production Impulse Responses.



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: Italy, Japan, Mexico, Netherlands, Portugal, Spain.

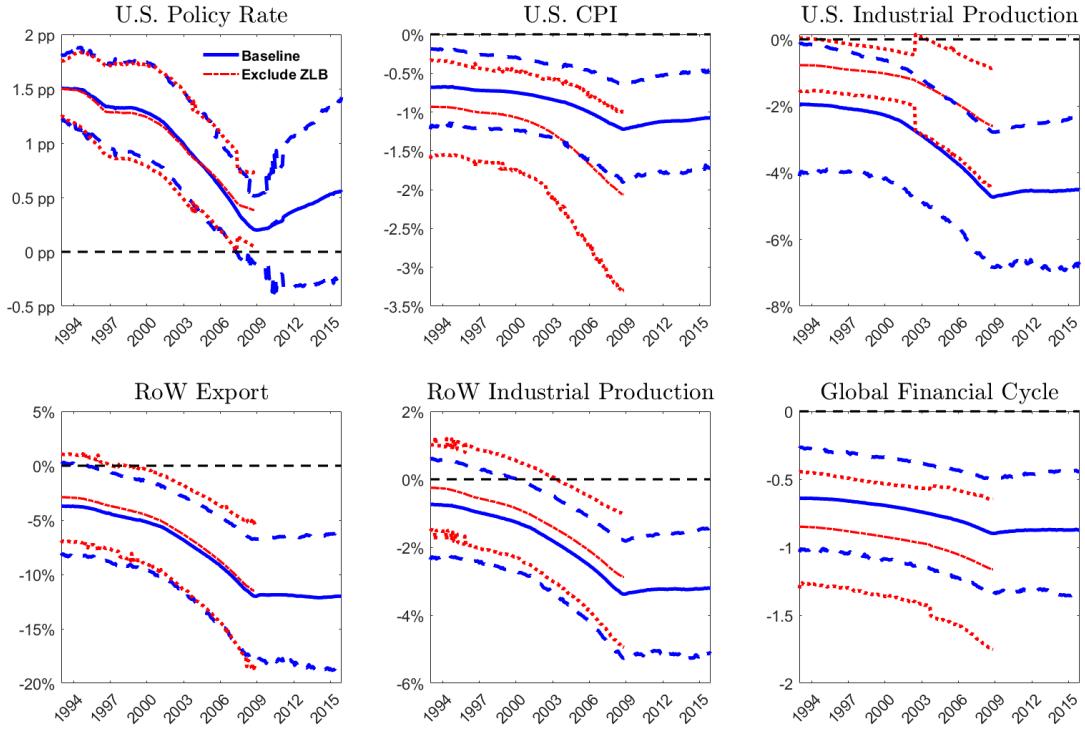
Figure E4: Country-Specific Industrial Production Impulse Responses.



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: Sweden, UK, Brazil, Chile.

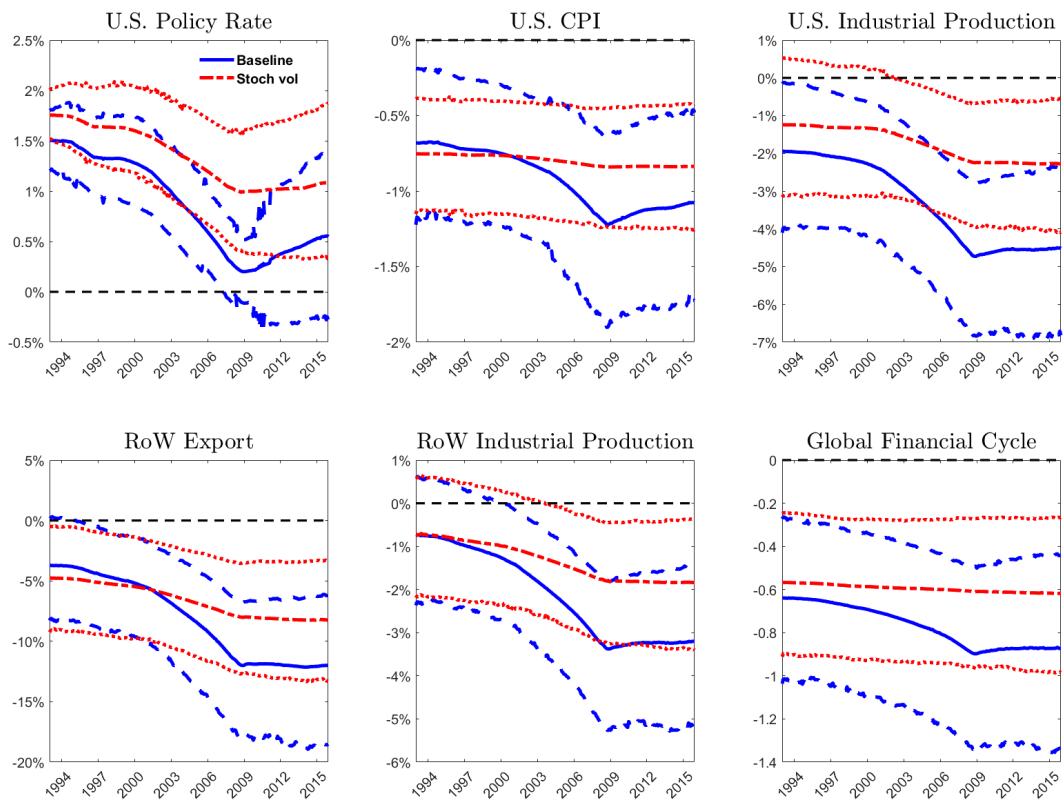
F. Additional Results of the Sensitivity Analysis

Figure F1: Robustness: Zero Lower Bound.



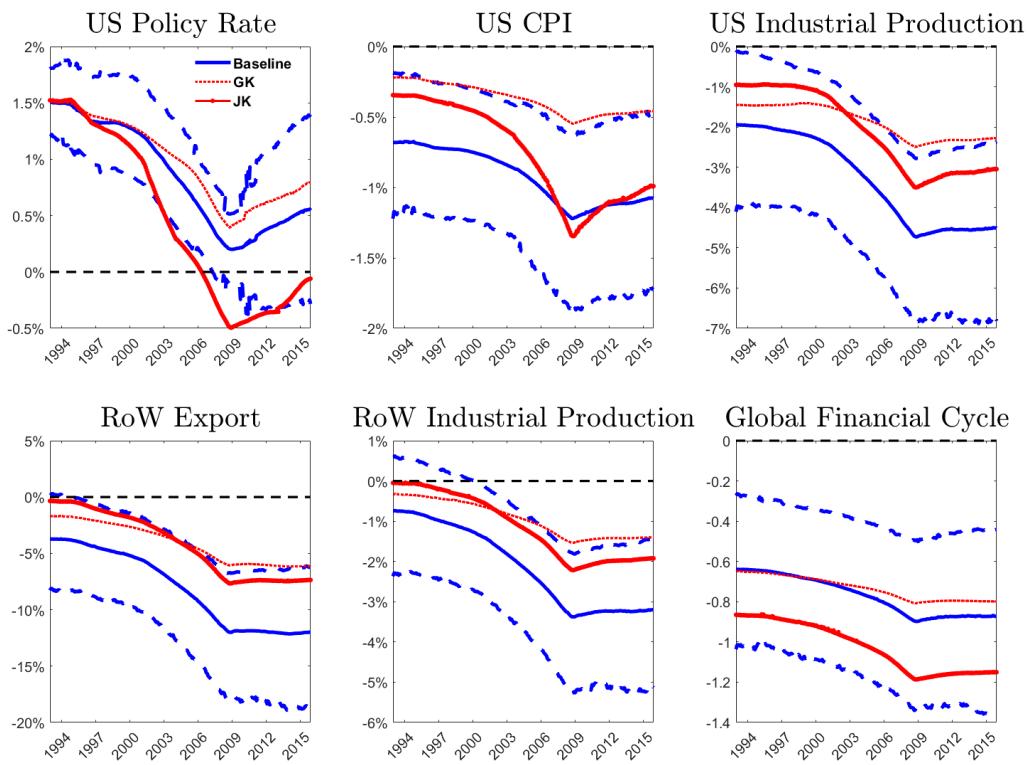
Notes: Responses to a contractionary U.S. monetary policy shock that induces a one percentage point increase in the federal fund rate in 1993M1.
Baseline model in blue; red lines correspond to the model excluding the ZLB period (ending in 2008M11).

Figure F2: Robustness: Stochastic Volatility – Global Responses.



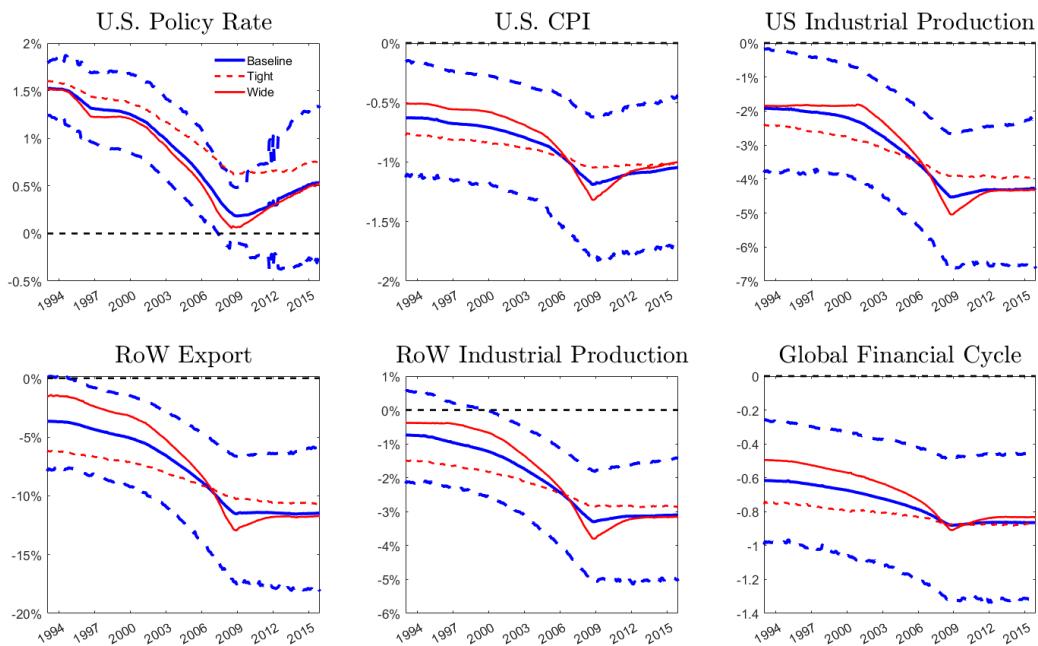
Notes: Peak Responses to a contractionary U.S. monetary policy shock that induces a one percentage point increase in the federal fund rate in 1993M1. Baseline model vs. extended model that allows for stochastic volatility specification.

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 ($\kappa_Q^2 = 0.015$).