What explains international interest rate co-movement?[*]

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Abstract

This paper decomposes international interest rate co-movement into contributions of domestic, foreign, and global supply, demand, and monetary policy shocks across seven advanced economies. We develop a Bayesian structural panel vector autoregression, integrating informative priors and homogeneity restrictions on contemporaneous relations, a hierarchical Minnesota prior with cross-sectional shrinkage, and a factor structure for structural shocks. We show that interest rate co-movements are driven by monetary policy responses to synchronized business cycle fluctuations, caused by demand shocks. In the short run, domestic shocks dominate interest rate variation. Cross-country spillovers arise gradually over time and have a smaller impact in larger economies.

Keywords: Spillovers, Structural vector autoregressions, Panel vector autoregressions, Informative priors

JEL-Codes: C11, C30, E52, F42

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1 Introduction

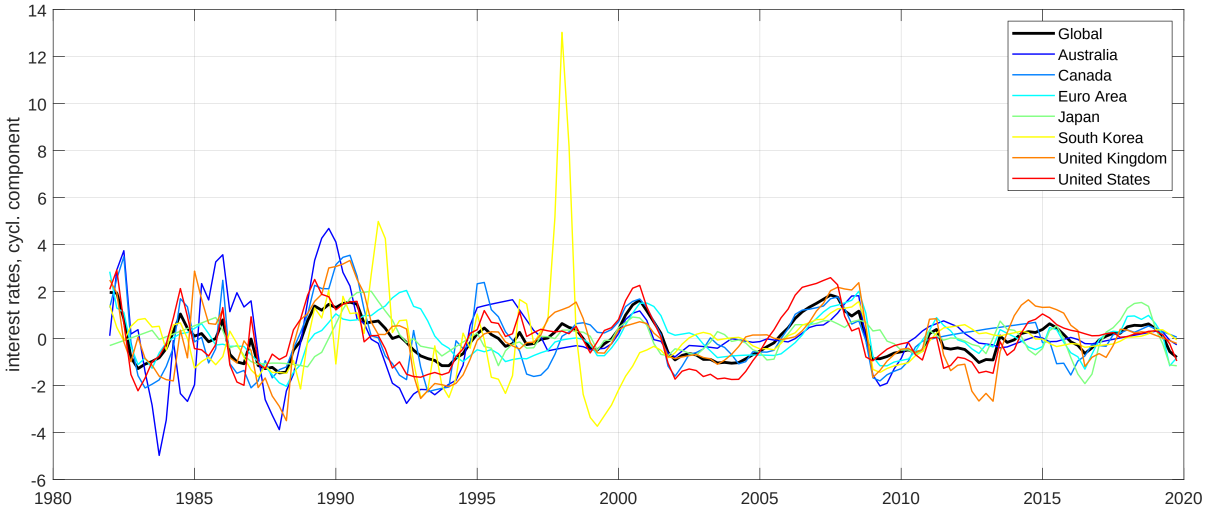
Over the past four decades, fluctuations in central bank policy rates exhibit strong international co-movement. Correlations between national interest rates (or shadow rates during zero-lower-bound periods) from seven major economics and a global GDP-weighted average are substantial. They range from around 80% (South Korea) to 95% (United States). These patterns are not only driven by a common downward trend, but also reflected in common cyclical fluctuations of interest rates displayed in Figure 1.

These high correlations raise a central policy question: Why do interest rates move together? Do monetary policy shocks spill over across borders, or do central banks systematically respond to common global economic conditions? These competing explanations imply different monetary policy rules. Spillovers suggest that central banks respond to foreign conditions, potentially constraining monetary autonomy. In contrast, common global supply or demand shocks can produce similar macroeconomic conditions across countries, generating synchronized policy responses even under purely domestic mandates. This is relevant because central bank behavior shapes agents' expectations: the effectiveness of policy depends on whether agents believe the central bank follows a domestic-oriented rule or one influenced by foreign variables.

Previous research has studied alternative explanations for correlated interest rates in isolation.[1] We jointly quantify the contribution of a broad set of structural shocks behind international interest rate co-movements using a Bayesian structural panel vector autoregressive (PVAR) model. These shocks are domestic, foreign, and global shocks to country-specific supply, demand, monetary policy, and exchange rate equations. In our framework, co-movement across countries can arise from foreign shocks, which capture country-specific events that spill over to other economies, or from global shocks, which reflect unexpected worldwide developments or correlated common shocks. Comparing the relative strength of

[1] One part of the literature focuses on spillovers of structural shocks from a single large country like the US (see Kim, 2001; Maćkowiak, 2007; Dedola et al., 2017), another on monetary policy shock spillovers between a group of countries (see Gambacorta et al., 2014; De Santis and Zimic, 2022), while yet another investigates the role of global factors (see Mumtaz and Surico, 2009; Charnavoki and Dolado, 2014; Forbes et al., 2024).

Figure 1: Co-movement of the cyclical component of interest rates



NOTES: The figure shows HP-filtered nominal policy interest rates, spliced with shadow rates (Krippner, 2013) during zero lower bound periods. We add a HP-filtered global interest rate (black line), calculated as the GDP-weighted average of the individual interest rates.

alternative explanations allows us to identify their quantitative importance. If monetary policy shocks contribute more to the correlation than other structural shocks, this would be evidence of direct cross-country spillovers. Our model also yields the relevant components of monetary policy rules, showing whether systematic interest rate responses focus exclusively on domestic variables or also include foreign ones.

Our model combines data at business-cycle frequencies from seven advanced economies: Australia (AU), Canada (CA), the Euro area (EA), Japan (JP), South Korea (KO), the United Kingdom (UK), and the United States (US). We use data on output gaps, inflation, shadow rates, and real effective exchange rates between 1980:Q3 and 2019:Q4.[2] We measure the drivers of interest rate co-movements by estimating the contribution of the structural shocks to (a) the forecast error variance of interest rates and (b) the forecast error correlation between domestic and global interest rates over different horizons.

We provide four main results. First, domestic shocks are the main driver of interest rates variations in the short run but become less important at longer horizons. The importance of foreign shocks increases with forecast horizons. The impact contribution of global

[2]We select these advanced economies because they jointly account for a large share of global GDP. For simplicity, we use *currency area* and *country* interchangeably in the following.

shocks varies strongly across countries, before becoming more similar in the medium run. Domestic shocks explain between 45% (Canada) and 95% (Japan and South Korea) of the forecast error variance decomposition (FEVD) of interest rates at short horizons. After five years, foreign and global shocks are about equally important. Their combined contribution reaches between 55% (United States) and 85% (Canada). We see the increasing long-run importance as evidence that cross-border spillovers take some time to materialize.

Second, the influence of global and foreign shocks is mostly restricted to demand shocks. Monetary policy shocks play a minor role for the FEVD with an average share below 10%. This indicates that monetary policy in all countries of our sample is conducted to a large extent autonomously, rather than being influenced by monetary policy shocks spilling over from a dominant economy.

Third, both global and foreign shocks contribute strongly to the forecast error *correlation* decomposition (FECorD) between domestic and global interest rates. Global shocks matter most at short horizons due to their synchronized impact across countries, while foreign shocks gain relevance over time as they gradually spill over across borders. The importance of domestic shocks depends on the openness and size of an economy: they dominate only in the US, and are much less important in small, open economies such as Australia and Canada, where global shocks are the main drivers. Thus, we document that monetary policy in small open economies adjusts largely due to foreign and global shocks.

Fourth, demand shocks are the most important cause of forecast error correlations of interest rates over all horizons. The three different types of demand shocks (domestic, foreign, global) have a joint share between 50% and 60% to the FECorD after five years. This is because demand shocks cause quasi-synchronous impulse-response functions of output gaps and inflation across countries. Interest rates are correlated because inflation-targeting central banks react similarly to these synchronous business cycle fluctuations. The identified national monetary policy rules confirm that central banks in our countries have a strong policy focus on domestic variables.

To summarize, we find little evidence of direct international monetary policy spillovers, supporting the view that observed interest rate co-movements are compatible with monetary autonomy. This finding aligns with theoretical results such as Kulish and Rees (2011).

Our analysis contributes to different strands of the literature on international interest rate co-movements. We complement work on the long-run decline in nominal and real interest rates (Holston et al., 2017; Del Negro et al., 2019), by adding insights on cyclical co-movements. We extend earlier studies on cross-country transmission, many of which examine only two countries or isolate shocks from a single large economy (e.g., Kim, 2001; Maćkowiak, 2007; Dees et al., 2007; Georgiadis, 2015). We build on recent papers that estimate multi-country spillovers of monetary policy shocks (Gerko and Rey, 2017; Rogers et al., 2018; Liu et al., 2022), but differ by also accounting for the role of supply and demand shocks, which we find to be more important. Our framework is also related to structural factor models and augmented factor VAR models (Charnavoki and Dolado, 2014; Mumtaz and Surico, 2009; Forbes et al., 2024), which document similar global effects, but do not allow for spillovers of local shocks as we do. Finally, we contribute to the literature on the international effects of US monetary policy shocks. Our results show that foreign interest rates do not react significantly to a US shock, as in Gerko and Rey (2017); Liu et al. (2022).[3]

Our Bayesian structural PVAR model introduces a coherent framework for structural multi-country models which integrates (1) informative priors, (2) homogeneity restrictions on the contemporaneous relations inspired by global VAR models, (3) hierarchical Minnesota prior with extra shrinkage on the cross-sectional dimension, and (4) a factor structure for structural shocks.

Informative priors on structural contemporaneous relationships, i.e. short-run elasticities and semi-elasticities (Baumeister and Hamilton, 2015, 2018), incorporate identification uncertainty around restrictions. Thereby, they account for a lack of conclusive theoretical evidence which is especially appealing in multi-country models.[4] We derive priors from

[3]Some studies do find significant spillovers but disagree on the sign (Feldkircher and Huber, 2016; Dedola et al., 2017; Rogers et al., 2018; Crespo Cuaresma et al., 2019; De Santis and Zimic, 2022).

[4]Informative priors avoid recursive structures (as in Chen et al., 2016; Bluwstein and Canova, 2016, where the order of countries matters) or block-exogeneity (as in Kim and Roubini, 2000; Kim, 2001; Maćkowiak, 2007) that are hard to justify. Compared to studies applying sign and/or magnitude restrictions in multi-country models (Gambacorta et al., 2014; Liu et al., 2022; De Santis and Zimic, 2022), we clearly acknowledge the uncertainty around restrictions and avoid controversial restrictions on foreign responses.

the open-economy model of Lubik and Schorfheide (2007). This theoretical model allows us additionally to impose economically meaningful exclusion and homogeneity restrictions on structural contemporaneous coefficients associated with foreign terms. Homogeneity restrictions discipline cross-country spillovers. They are akin to, but not equivalent, to those in global VARs, and arguably have stronger theoretical justification.

We adapt the prior setup of Baumeister and Hamilton (2015) to account for potential differences across countries in our model. To do so, we rely on a hierarchical prior set-up which adds flexibility in estimating prior variances as in Giannone et al. (2015), and impose extra shrinkage by letting the prior variance of autoregressive parameter decrease with county-specific GDP weights. We extend the model further and impose a factor structure in the spirit of Stock and Watson (2005). This allows us to differentiate between country-specific and latent global supply, demand, and monetary policy shocks in our model. We identify factors through zero and sign restrictions on the factor loadings as in Korobilis (2022). Global supply shocks capture both commodity-driven disruptions, such as oil price fluctuations, and broader structural changes, such as China's accession to the WTO. Global demand shocks are related to global disturbances in the financial sector. We include global monetary policy shocks to allow for coordinated policy actions.

We present our model and discuss identification challenges in detail in section 2. The results section 3 provides additional details on the main results mentioned above, discusses the nature of global shocks and investigates the monetary policy rules identified by our model. Section 4 concludes the paper.

2 Bayesian structural PVAR model

The Bayesian structural PVAR model combines J equations per country, for C countries. In matrix form, it is given by

$$\mathbf{A}\mathbf{y}_t = \mathbf{B}\mathbf{x}_{t-1} + \chi\mathbf{u}_{gt} + \mathbf{u}_t \tag{1}$$

$$\mathbf{u}_t \sim \mathcal{N}(\mathbf{0}, \mathbf{D}), \qquad \mathbf{u}_{gt} \sim \mathcal{N}(\mathbf{0}, \mathbf{I}_G).$$

The endogenous variables are captured in the $(n \times 1)$ vector $\mathbf{y}_t = (\mathbf{y}_{1t}, \ldots, \mathbf{y}_{Ct})'$ with $\mathbf{y}_{ct}$ containing the J endogenous variables of country c, and $n = CJ$. The right-hand-side variables include p lags and a constant collected in the $(k \times 1)$ vector $\mathbf{x}_{t-1}$, with $k = CJp+1$. The main results hold for models with alternative lag structures and with a deterministic trend, Appendix Figure E.2.

The $(n \times n)$-matrix $\mathbf{A}$ contains structural contemporaneous parameters, which can be interpreted as semi-elasticities and elasticities. We impose three types of restrictions on the contemporaneous coefficients: (1) trade-weighted homogeneity restrictions on some foreign terms, (2) exclusion restrictions mostly on the remaining foreign terms, and (3) informative prior distributions (Baumeister and Hamilton, 2018). We discuss the specific model equation and identification restrictions in subsection 2.1. The $(n \times k)$-matrix $\mathbf{B}$ contains the structural lag coefficients. We use hierarchical shrinkage priors (Giannone et al., 2015), where the degree of shrinkage depends on GDP weights, to deal with the large numbers of autoregressive parameters. The prior setting is discussed in subsection 2.2.1.

We allow for "combined" structural shocks, $\chi \mathbf{u}_{gt} + \mathbf{u}_t$, which consist of structural country-specific shocks and global shocks. The $(n \times 1)$ vector of structural country-specific shocks $\mathbf{u}_t$ jointly follows a normal distribution with mean zero and diagonal variance matrix $\mathbf{D}$. These shocks are uncorrelated across equations and countries. We capture potential correlation across countries through G structural global shocks $\mathbf{u}_{gt}$. The global shocks independently follow a standard-normal distribution and load onto each country according to the $(n \times G)$-dimensional loading matrix χ, which collects the country- and shock-specific loadings χ_c^j. The loading matrix is identified through economically motivated zero and sign restrictions (Korobilis, 2022), as described in subsection 2.2.

Our specification ensures that contemporaneous co-movement across countries can materialize via the effect of domestic shocks on foreign countries or through global shocks. Switching perspectives, we refer to the former as foreign country-specific events spilling over to other countries. Global shocks induce economic fluctuations that are common across countries. These shocks reflect either unpredictable changes in global developments outside our model, or shocks that are common for all the countries in our sample. For example, they can be driven by so-called "primitive" shocks (as defined in e.g., Ramey, 2016) such as natural disasters or geopolitical events having a worldwide effect, or by co-ordinated monetary policy shocks leading to all central banks jointly changing their policy rates (Georgiadis and Jančoková, 2020).

2.1 Structural contemporaneous relations

We use the structural PVAR model to investigate the correlation of interest rates for Australia, Canada, the Euro Area, Japan, South Korea, the United Kingdom and the United States; this set of countries represents around 54% of the world's economic activity as of 2019. For each country $c \in \{1, \ldots, C\}$, we include the output gap (y_{ct}) as a measure of economic activity, year-on-year inflation rates (π_{ct}), Krippner-shadow interest rates (r_{ct}) – which capture both conventional and unconventional monetary policy actions (Krippner, 2013) – and year-on-year growth rates of the real effective exchange rates (q_{ct}), all collected in $\mathbf{y}_{ct} = (y_{ct}, \pi_{ct}, r_{ct}, q_{ct})'$. The real effective exchange rates are defined such that an increase in q_{ct} indicates an increase in competitiveness. Our sample spans quarterly data from 1980:Q3 to 2019:Q4.[5] Our main results are qualitatively robust to variations in the data sample and choice of variables.

The structural relations between endogenous variables are derived from theoretical open-economy models such as Lubik and Schorfheide (2007). Specifically, we formulate an empirical open economy Phillips curve (labeled "s"), IS curve ("d"), monetary policy rule ("m"), and an exchange rate equation ("er"). We index the type of structural equation

[5] Online Appendix A explains the data in more detail. Two country selections deserve note. First, we rely on constructed data (provided by Eurostat, the ECB and Oxford Economics) for a counterfactual Euro Area between 1980 and 1999. Second, we exclude China because, especially for the first part of our sample, there are issues with the availability and quality of Chinese data.

with superscript $j \in \{s, d, m, er\}$ for each country c in our sample:

$$y_{ct} = \alpha^{c,\pi}\pi_{ct} + \alpha^{c,q}q_{ct} + \text{lag terms} + \chi_c^s u_{gt}^s + u_{ct}^s \tag{s}$$

$$y_{ct} = \beta^{c,r}r_{ct} + \beta^{c,\pi}\pi_{ct} + \beta^{c,q}q_{ct} + \beta^{c,y^*}y_{ct}^* + \text{lag terms} + \chi_c^d u_{gt}^d + u_{ct}^d \tag{d}$$

$$r_{ct} = (1 - \rho^c)\left(\psi^{c,\pi}\pi_{ct} + \psi^{c,y}y_{ct} + \psi^{c,q}(q_{ct} - \pi_{ct}^*)\right) + \text{lag terms} + \chi_c^m u_{gt}^m + u_{ct}^m \tag{m}$$

$$q_{ct} = \theta^{c,y}(y_{ct} - y_{ct}^*) + \text{lag terms} + u_{ct}^q. \tag{er}$$

The "lag terms" contain $p = 4$ lags and a constant. We model common economic fluctuations through $G = 3$ global shocks in the supply, demand and monetary policy equation. We exclude a global exchange rate shock in our baseline model, since the currencies in our analysis are the overwhelmingly dominant currencies in the world during the time of our analysis.

The $\alpha, \beta, \psi, \theta$-coefficients are the (semi-)elasticities of the structural equations (s) to (er). The structural coefficients $\beta^{c,y^*}, \psi^{c,q}, \theta^{c,y}$ capture spillovers from foreign endogenous variables. Based on theoretical open-economy models, we add real effective exchange rate growth to both the supply and demand curve, and include foreign output gaps in the demand curve. The monetary policy rule follows an open-economy version of the standard Taylor rule. Here, the monetary policy authority can also set interest rates in relation to changes in nominal exchange rate fluctuations. Finally, the exchange rate equation is introduced under the assumption of purchasing power parity and relates changes in the real exchange rate to differences in foreign and domestic output. Using the theoretical model, this implies a dependence of real exchange rates on changes in the terms of trade.

With these structural relations, we impose three different types of restrictions on the contemporaneous relations in $\mathbf{A}$: homogeneity restrictions on the foreign terms, exclusion restrictions, and informative prior distributions.

2.1.1 Homogeneity and exclusion restrictions

Homogeneity restrictions on the foreign terms allow us to aggregate foreign terms into a single variable and coefficient. All aggregated contemporaneous foreign variables (y_{ct}^*, π_{ct}^*)

are trade-weighted averages of country-specific terms.[6] As an example, β^{c,y^*} is the demand elasticity of country c with respect to foreign output gaps. The homogeneity restriction implies that the demand elasticity with respect to output gaps from country $c', c \neq c'$ is $\beta^{c,y_{c'}} = w_{cc'}\beta^{c,y^*}$. That is, the elasticities scale with the importance of the foreign country c' in the trade basket of country c. The restriction allows us to identify the $C-1$ coefficients $\beta^{c,y_{c'}}$ through a single parameter β^{c,y^*}. This effectively removes the curse of dimensionality for $\mathbf{A}$ arising from the panel dimension of the model. We do not impose homogeneity restrictions on lag coefficients.

While the homogeneity restrictions are comparable to those used in GVAR models (see, e.g., Dees et al., 2007; Feldkircher and Huber, 2016; Crespo Cuaresma et al., 2019), we set them on the structural contemporaneous coefficients instead of the reduced from parameters and covariance, since the former matches theoretical considerations. The two alternatives are not equivalent: homogeneity restrictions on the reduced form place the restrictions on the inverse of $\mathbf{A}$. That is, GVAR models usually assume that the impact of *foreign shocks*, instead of *elasticities* as in our case, scales with the amount of trade with the country of origin.

As a second restriction, the structural relations formulated in equations (s) to (er) exclude contemporaneous relations among some variables, directly reducing the number of coefficients to be identified. For example, the Phillips curve does not contain an interest rate term, or any foreign terms. The exclusion restrictions are uncontroversial, as they apply mostly to the role of foreign variables. Yet, we show that our results are robust if we deviate from some of them, for example, by including interest rate differentials and inflation differentials in the exchange rate equation to allow for uncovered interest rate parity.

2.1.2 Informative priors on contemporaneous parameters

We implement prior beliefs on the structural contemporaneous coefficients as in Baumeister and Hamilton (2018). We set a t-distribution with a prior scale of 0.4 and three degrees

[6] We derive bilateral trade weights $w_{c,c'}$ from the BIS. In our sample period, the BIS reports 3-year averages of bilateral trade weights for their "narrow basket" real effective exchange rates. We average these weights and normalize them such that they sum up to 1.

Table 1: Prior on contemporaneous parameters for country c

Parameter	Prior mode	Prior scale	Restrictions
Parameters set as in Baumeister and Hamilton (2018)			
Student-t distribution with 3 degrees of freedom			
$\alpha^{c,\pi}$	2	0.4	≥ 0
$\beta^{c,r}$	-1	0.4	≤ 0
$\beta^{c,\pi}$	0.75	0.4	
$\psi^{c,\pi}$	1.5	0.4	≥ 0
$\psi^{c,y}$	0.5	0.4	≥ 0
Beta(2.6,2.6)			
ρ^c	0.5	0.2	$0 \leq \rho^c \leq 1$
Additional parameters in open-economy model			
Student-t distribution with 3 degrees of freedom			
$\alpha^{c,q}$	-0.5	0.4	
$\beta^{c,q}$	0.2	0.4	
β^{c,y^*}	0.5	0.4	≥ 0
$\psi^{c,q}$	0	0.4	
$\theta^{c,y}$	1	0.4	≥ 0

Notes: We choose prior distribution families (incl. degrees of freedom and scales) as in Baumeister and Hamilton (2018). The mode and sign restrictions for parameters referring to domestic coefficients (upper panel) are set as in Baumeister and Hamilton (2018). The mode and restrictions for the remaining parameters (lower panel) are derived from the small open economy model of Lubik and Schorfheide (2007).

of freedom for the majority of parameters, allowing for heavier tails compared to a normal distribution. We follow Baumeister and Hamilton (2018) in the prior specifications for the parameters of their closed-economy model, $(\alpha^{c,\pi}, \beta^{c,r}, \beta^{c,\pi}, \psi^{c,\pi}, \psi^{c,y}, \rho^c)_{c=1}^C$, shown in the upper block of Table 1. We refer the reader to their paper for a detailed discussion on these prior beliefs.

The prior specifications of the the remaining structural contemporaneous coefficients, $(\alpha^{c,q}, \beta^{c,q}, \beta^{c,y^*}, \psi^{c,q}, \theta^{c,y})_{c=1}^C$, are given in the lower block of Table 1. We derive our informative prior beliefs based on theoretical insights, mainly from the New-Keynesian small open economy model of Lubik and Schorfheide (2007). To do this, we express each parameter as a function of deep structural parameters from a theoretical model, and determine the prior mode based on common values of these parameters from the literature.[7] Note that we do not differentiate priors across different countries. However, we provide robustness checks where we check the sensitivity of results to our prior choices, see Appendix Figure E.4.

In equation (s), the parameter $\alpha^{c,q}$ measures the elasticity of supply to real effective exchange growth rates. In Lubik and Schorfheide (2007), this elasticity depends on the import share, the intertemporal substitution elasticity, discount rate and the slope coefficient of the Phillips curve. We set the import share to 0.2. To avoid singularities in the theoretical model, intertemporal substitution elasticities are commonly restricted to be larger than zero and smaller than one (Lubik and Schorfheide, 2007; Justiniano and Preston, 2010). We follow Lubik and Schorfheide (2007) and settle for a mean of 0.5. As in Baumeister and Hamilton (2018) we assume a discount rate of zero and a slope coefficient of 0.25. Similarly, Lubik and Schorfheide (2004) set a prior for the slope coefficients allowing for a wide rage between 0 and 1. Last, as in Baumeister and Hamilton (2018), we replace expected values in the theoretical model with autoregressive forecasts from an AR(1) process with autoregressive parameter of 0.75. Based on these values we choose a prior mode for $\alpha^{c,q}$ of -0.5.

In the open economy IS curve, the parameter $\beta^{c,q}$ measures the dependence of aggregate demand on competitiveness. It combines the weight of the forward-looking component of the IS curve (for which we assume a value of 0.67) with the import share and the intertemporal substitution elasticity. Our choices for these three parameters follow Baumeister and Hamilton (2018); Lubik and Schorfheide (2007) and imply a prior mode of 0.2 for $\beta^{c,q}$.

We set a positive prior mode of 0.5 for β^{c,y^*}. The value results from the relation of the impact of foreign output on domestic output to the forward-looking component of output in the IS curve, import share, and the intertemporal substitution elasticity. By allowing for global shocks, we intentionally weaken shock transmission via the aggregate demand channel described by β^{c,y^*}. To limit this weakening effect, we introduce a sign restriction on $\beta^{c,y*}$.

[7] Online Appendix B provides the exact relations of the coefficients in the SVAR model and the theoretical model of Lubik and Schorfheide (2007).

We assume that the monetary policy authority can set interest rates according to a generalized Taylor rule in line with the specification in Lubik and Schorfheide (2007), Adolfson et al. (2007), and Justiniano and Preston (2010), which lets the central bank react to nominal exchange rates. This implies that the coefficient on foreign inflation is $-\psi^{c,q}$, the negative of the coefficient on domestic real effective exchange rate growth. We set the prior mode of this coefficient to zero (as in Adolfson et al., 2007), as the countries in our sample are characterized by flexible exchange rate regimes.

Equation (er) determines exchange rates as a function of contemporaneous domestic and foreign output as well as foreign exchange rates. We assume that purchasing power parity holds, which implies that q_{ct} directly relates to the terms of trade. As in Lubik and Schorfheide (2007), the difference in domestic and foreign demand growth thus determines the terms of trades endogenously, such that growth in domestic and foreign demand balances out. We hence apply $\theta^{c,y}$ to the difference between domestic and foreign output gaps. A prior mode of one is based on the relation of import shares and intertemporal substitution elasticity.

Following Baumeister and Hamilton (2018), we impose three additional sets of priors related to impact responses to economic shocks. For these priors, we use non-dogmatic asymmetric t-distributions with location parameter μ, scale parameter σ, degrees of freedom ν, and shape parameter λ, the latter one controlling the degree of asymmetry. First, we assume for every country that the output response to a contractionary monetary policy shock is smaller than the interest rate response ($\mu = -0.3; \sigma = 0.5; \nu = 3; \lambda = -2$). This prior leaves 12.5% chance that the output reaction is (in absolute terms) larger than the interest rate response, and a 4.5% chance that it is positive. Second, we assume a positive output response to aggregate supply shocks ($\mu = 0.3; \sigma = 0.5; \nu = 3; \lambda = 2$), which leaves a 4.5% chance of a negative output reaction. Third, we set as prior distribution on $\beta^{d,c\pi} - \beta^{d,cr}(1-\rho^c)\psi^{c\pi}$ as in Baumeister and Hamilton (2018). In a New-Keynesian model, this element describes the slope of the IS curve after replacing nominal interest rates by the monetary policy rule. Because we assume a negative slope, we use an asymmetric t-distribution with mode $\mu = -0.1$, scale $\sigma = 1$, degrees of freedom $\eta = 3$ and asymmetry parameter $\lambda = -4$, allowing for a 6.5% chance of a positive slope.

2.2 Remaining priors, and posterior inference

2.2.1 A flexible Minnesota prior for structural variances and lag coefficients

Baumeister and Hamilton (2018) condition a Minnesota prior for $\mathbf{B}$ and an inverse gamma prior for $\mathbf{D}$ on structural contemporaneous coefficients $\mathbf{A}$. We follow their strategy, with two main differences, which are explained in more detail in Appendix C.3.

First, we use hierarchical priors for the overall scaling parameter of the Minnesota prior, λ_0, and the variance component related to individual variables i in country c, $\mathbf{s} = \{s_c^i\}_{\substack{i\in\{y,\pi,r,q\} \\ c\in\{1,\ldots,C\}}}$. Following Giannone et al. (2015), we set a gamma distribution as prior for λ_0 and an inverse gamma distribution for s_c^i:

$$p(\lambda_0) = \gamma(\lambda_0; \kappa_{\lambda_0}, \tau_{\lambda_0}), \quad \text{and} \quad p(s_c^i) = \gamma((s_c^i)^{-1}; \kappa_{s_c^i}, \tau_{s_c^i}).$$

For λ_0 we choose a mode of 0.2 and standard deviation of 0.4. For s_c^i, we set a shape of 0.1, and a scale of 0.05 for $i \in \{y, \pi, r\}$, and a scale of 2 for $i = q$. This distribution has a mode close to the variance of an AR(4) process of the corresponding variable. At the same time, it is very dispersed, without a finite variance or mean.

The hierarchical priors give more flexibility to the prior of structural variances $\mathbf{D}$, where each element d_c^j follows an inverse gamma distribution with shape $\kappa = 2$ and a scale τ_c^j that depends on draws of $\mathbf{A}$, $\mathbf{s}$ and the (fixed) correlation of residuals from AR(4) processes estimated for each endogenous variable:

$$p\left(d_c^j | \mathbf{A}, \mathbf{s}\right) = \gamma\left((d_c^j)^{-1}; \kappa, \tau_c^j(\mathbf{A}, \mathbf{s})\right).$$

As a second difference to Baumeister and Hamilton (2018), we set tighter priors for lag coefficients on variables from countries with a smaller GDP. This approach embodies our belief that past developments from smaller countries may be less important, while simultaneously addressing the curse of dimensionality in the structural lag coefficients. A model without size-dependent shrinkage (i.e., with a standard Minnesota prior) produces nearly identical median results, see Figure E.2 in the Appendix, but with higher uncertainty.[8]

The conditional distribution of $p(\mathbf{B}|\mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{D})$ is a normal distribution for each $\mathbf{b}_c^j$:

$$p\left(\mathbf{b}_c^j|\mathbf{A}, \lambda_0, \mathbf{s}, d_c^j\right) = \phi\left(\mathbf{b}; \mathbf{m}_c^j(\mathbf{A}), d_c^j\mathbf{M}_c^j(\lambda_0, \mathbf{s})\right).$$

The prior mean, $\mathbf{m}_c^j(\mathbf{A})$, multiplies the belief that each data series follows an AR(1)-process with coefficient 0.75 with $\mathbf{A}$. The prior variance of lag l of variable i in country c' for structural equation j in country c is

$$d_c^j\lambda_0^2\frac{\omega_{c'}^2}{l^2 s_{c'}^i},$$

where $\omega_{c'}$ is the GDP weight of country c'. Hence, the prior precision depends on the variance ratio $d_c^j/s_{c'}^i$, scales with λ_0^2, and is increasing with the lag length and inverse GDP weights. We use fairly uninformative priors on the constant terms, with a variance of $100\lambda_0^2$. We account for interest rate smoothing in the monetary policy equation through an additional prior on the lagged coefficients of the monetary policy rate of country c (Baumeister and Hamilton, 2018). This prior is a multivariate normal distribution with mean zero with the exception of the coefficient on $r_{c,t-1}$, for which we set a mean of ρ^c, and variance $0.1d_c^m$.

2.2.2 A structural prior for global shocks and their loadings

We assume that our combined structural shocks, $\chi\mathbf{u}_{gt} + \mathbf{u}_t$, follow a static factor structure. To obtain global shocks $\mathbf{u}_{gt}$ and their loadings χ_c^j we follow Korobilis (2022), who assumes a similar structure for the reduced form residuals and imposes sign restrictions on the factor loadings. Because global shocks only load onto one type of structural equation, we only have a single loading per equation. We assume that global supply shocks and demand

[8] Canova and Ciccarelli (2004); Koop and Korobilis (2016); Korobilis (2016); Koop and Korobilis (2019); Camehl (2023), for example, develop alternative ways to estimate *reduced form* PVAR models albeit without addressing structural identification.

shocks are expansionary, and that global monetary policy shocks are contractionary in all countries of our sample, hence, $\chi_c^j \geq 0, \forall c \in \{1, \ldots, C\}, j \in \{s, d, m\}$. We set normal prior distributions truncated at zero for the loadings χ_c^j with a large prior variance, and standard-normally distributed global shocks u_{gt}^j:

$$\begin{aligned} p(u_{gt}^j) &= \mathcal{N}(0,1), && j \in \{s, d, m\} \\ p(\chi_c^j|\mathbf{D}) &= \mathcal{TN}_{\chi_c^j \geq 0}(0, d_c^j V_\chi), && V_\chi = 100. \end{aligned}$$

2.2.3 Drawing from the posterior distribution

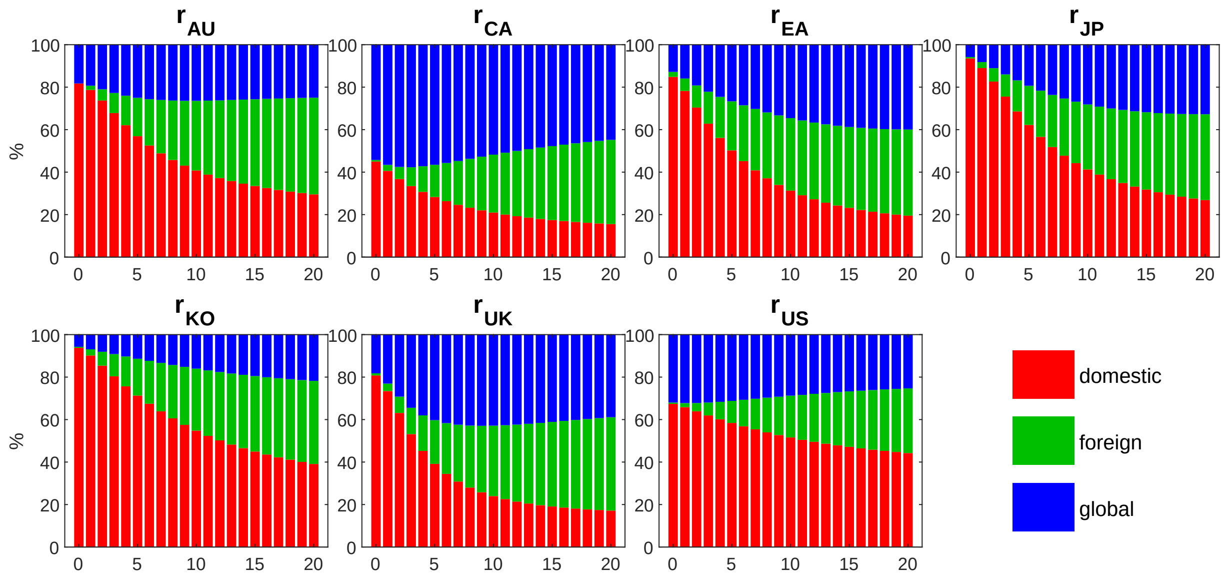
In summary, the joint prior distribution of the model is

$$p(\mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{B}, \mathbf{D}, \chi, \mathbf{U}_{gT}) = p(\mathbf{A})p(\lambda_0)p(\mathbf{s})p(\mathbf{D}|\mathbf{A}, \mathbf{s})p(\mathbf{B}|\mathbf{A}, \mathbf{D}, \lambda_0, \mathbf{s})p(\chi|\mathbf{D})p(\mathbf{U}_{gT}), \quad (2)$$

where the $(T \times G)$ matrix $\mathbf{U}_{gT} = \left[\mathbf{u}_{g1}', \ldots, \mathbf{u}_{gT}'\right]'$ combines all structural global shocks. Combining the prior distributions in Equation (2) with the data likelihood gives the posterior distribution. We use a Gibbs sampler with three main parts to sample from the posterior. In the first part, we use a Metropolis-Hastings step to sample from the conditional distribution $p(\mathbf{A}, \lambda_0, \mathbf{s}|\mathbf{Y}_T, \chi, \mathbf{U}_{gT})$. In the Appendix, we show that Baumeister and Hamilton (2015) and Giannone et al. (2015) use the same type of marginal likelihood, allowing us to combine $(\mathbf{A}, \lambda_0, \mathbf{s})$ in a single step, where global shocks and loadings can be treated as part of the data. In the second part, we draw structural variances $\mathbf{D}$ and lag coefficients $\mathbf{B}$ from the usual normal-inverse gamma model. In the third step, we draw loadings from truncated normal distributions, and normally distributed global shocks as in Korobilis (2022). Details are described in Appendix C.4.

The posterior distributions of the structural contemporaneous coefficients $\mathbf{A}$ converge, see Appendix C.5. They clearly update the informative priors, see Appendix Figures D.5 to D.7. The only exception are the slope of the Phillips curve $\alpha^{c,\pi}$ and the elasticity of growth rates of real effective exchange rates with respect to output growth differentials, $\theta^{c,y}$. However, a robustness check with larger prior scales (1 instead of 0.4) on these two

Figure 2: Forecast error variance decomposition of interest rates



NOTES: The figure shows the forecast error variance decomposition of country-specific interest rates (in subplots) to domestic, foreign and global shocks over 20 quarters.

parameters results in posteriors that are very close to our baseline model. This indicates that the information in the data merely confirms the prior, see Figure E.4 in the Appendix.

3 Results

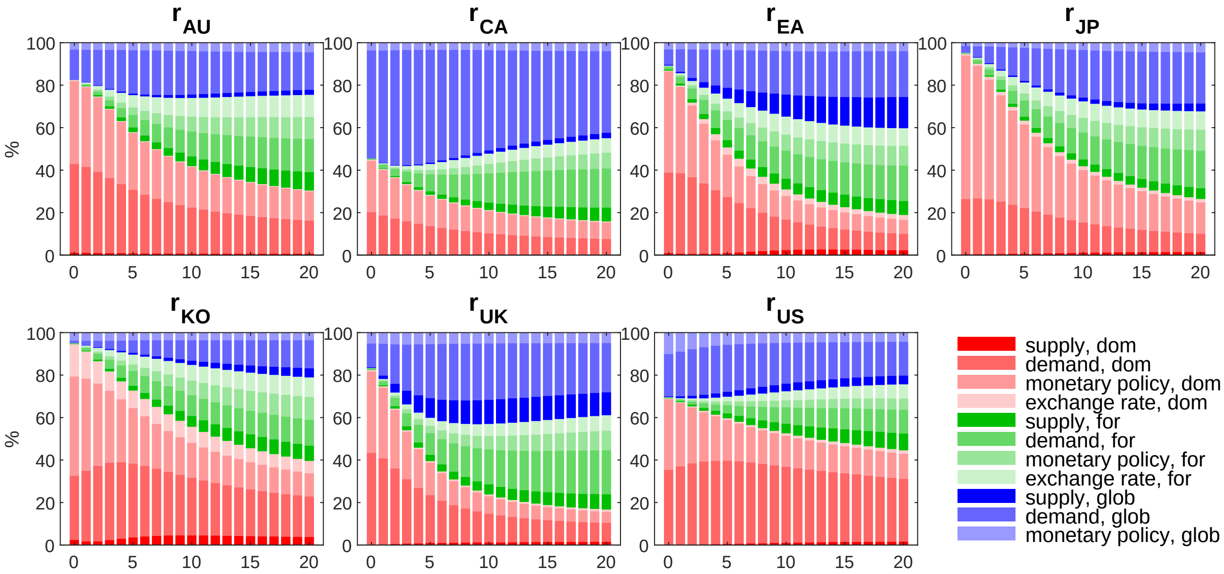
3.1 Drivers of interest rates

First, we quantify the importance of international factors in explaining interest rate movements. To that end, we compute median forecast error variance decomposition for interest rates to domestic, foreign, and global shocks over 20 quarters, shown in Figure 2.[9]

Our first key finding is that domestic shocks are the primary drivers of interest rates variations at short horizons. Their influence is initially strong, explaining between 45% (Canada) and 95% (Japan and South Korea) of the forecast error variance decomposition on impact, but declines rapidly over time. Domestic shocks tend to be more influential in countries that are either large (e.g., the United States) or have interest rates that are relatively uncorrelated with global interest rates (e.g., Australia and Korea). Foreign shocks

[9]Note that, while contributions sum up to 100 for every draw of the posterior distribution, this does not necessarily apply for the median. We normalize contributions to 100 even though differences are negligible.

Figure 3: Detailed forecast error variance decomposition of interest rates

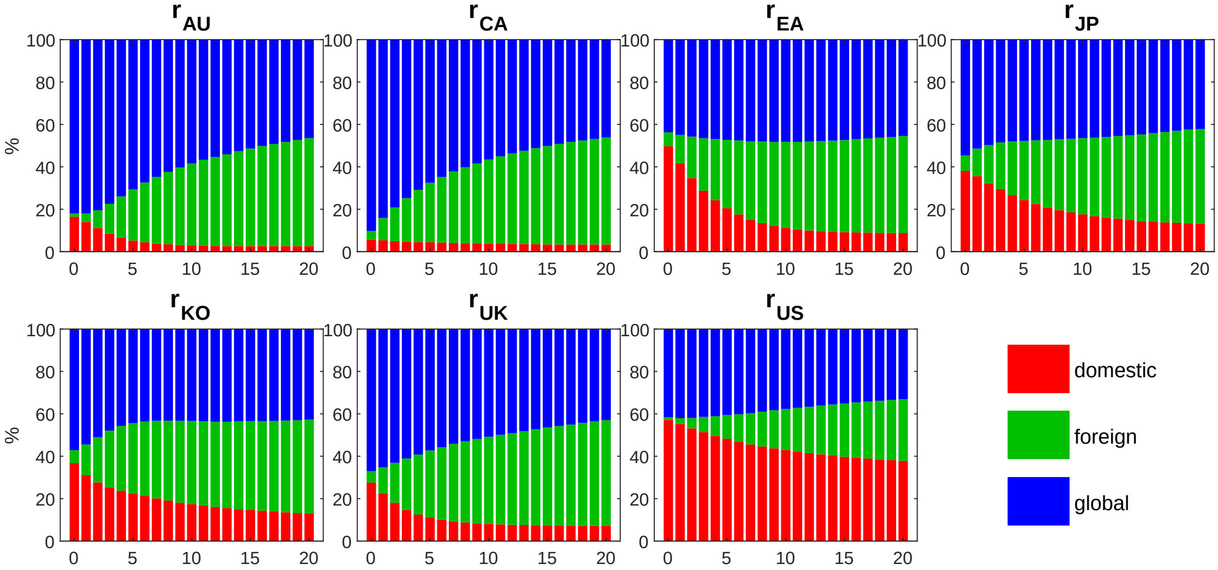


NOTES: The figure shows the forecast error variance decomposition of country-specific interest rates (in subplots) to domestic (dom) and foreign (for) supply, demand, monetary policy and exchange rate shocks and global (glob) supply, demand and monetary policy shocks over 20 quarters.

have no immediate effect but gain importance at longer horizons. That is, spillovers from other countries become more important over time. The reason for this is that foreign shocks change economic conditions in their origin-country directly but transmit to other countries only indirectly through their influence on endogenous foreign variables. Because the structural contemporaneous coefficients on these variables (in $\mathbf{A}$) are relatively small, the short-run impact of foreign shocks is limited. Meanwhile, global shocks enter the structural equations directly. On impact, contributions vary between below 5% (Korea) and 55% (Canada). They then converge to a range of 20% to 45% after 5 years for all countries. Overall, the non-negligible share of foreign and global shocks in explaining variation in interest rates rationalizes the positive correlation of interest rates across countries.

Our second key finding is that within both global and foreign shocks, demand shocks are the dominant contributors to interest rate fluctuations. Together they explain between 25% (Korea and United States) and 55% (Canada) of the overall variation after five years. Global and foreign monetary policy shocks contribute less to the FEVD, below 10% over all horizons, see Figure 3. This pattern is consistent across all forecast horizons and all countries in our sample. While domestic monetary policy shocks do play a larger role, their

Figure 4: Forecast error correlation decomposition between global and country-specific interest rates



NOTES: The figure shows median forecast error correlation decomposition between global (GDP weighted) and country-specific interest rates (in subplots) for domestic, foreign and global shocks over 20 quarters.

influence declines steadily over time across all countries, in line with previous findings by De Santis and Zimic (2022) and Baumeister and Hamilton (2018). Taken together, these findings suggest that international monetary policy spillovers are limited. We interpret the small role of global and foreign monetary policy shocks as evidence of monetary policy autonomy, i.e. central banks are largely shielded from international policy influences. Instead, the primary drivers of common interest rate variations are demand shocks.

Next, we quantify in Figure 4 interest rate co-movement by decomposing the correlation of forecast errors of country-specific and global interest rates. We compute the latter as the GDP-weighted sum of the country-specific interest rates. Adapting the usual formula for the forecast error variance decomposition, the contribution of shock k to the covariance of forecast errors between the interest rates of country c and the global interest rate at horizon H is

$$100 \times \frac{\sum_{h=0}^{H-1} \theta_{ck,h}\theta_{gk,h}d_k}{\sum_{k=1}^{n}\sum_{h=0}^{H-1} \theta_{ck,h}\theta_{gk,h}d_k}$$

where $\theta_{ck,h}$ is the impulse-response of interest rates of country c to shock k at horizons $h = \{0, \dots, H-1\}$ and $\theta_{gk,h}$ the response of the global interest rate, which we again calculate as the GDP-weighted average of country-specific responses.

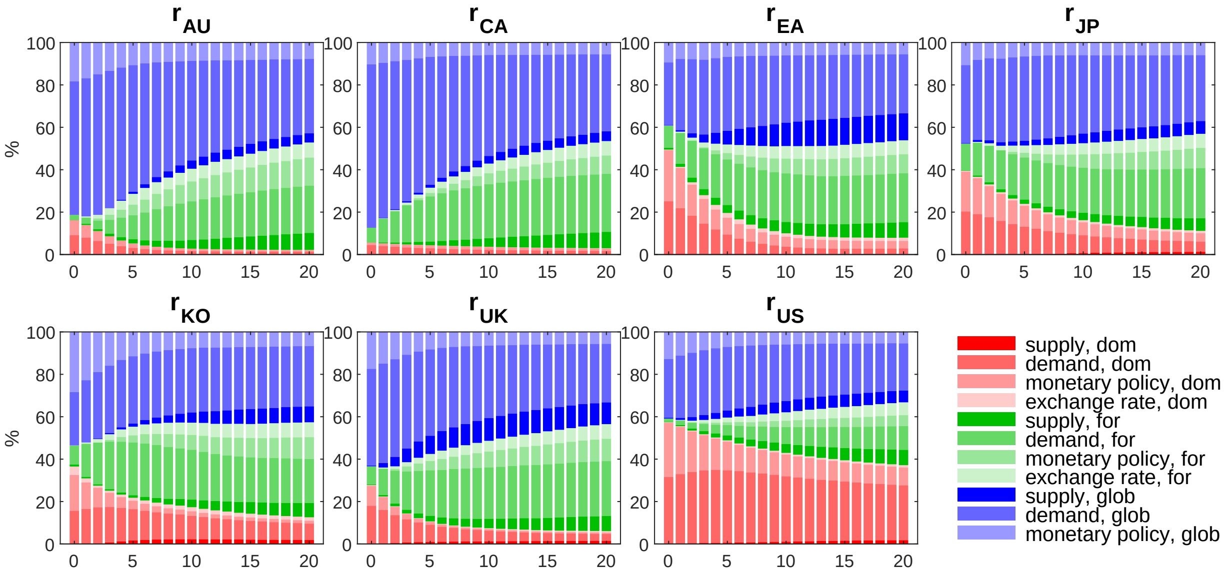
As the third key result, we show that both global and foreign shocks contribute to forecast error *correlations* between country-specific and global interest rates. In the short run, global shocks account for a substantial share, between 40% and 90% of these correlations. This reflects that global shocks cause synchronized economic fluctuations, and result in immediate co-movements in interest rates across countries. In contrast, foreign shocks originate in a single country and need time to propagate internationally. Their influence grows over time, while the contribution of global shocks declines. After 20 horizons, each explain between 30% and 50% of the forecast error correlation across countries.

The size and openness of an economy influence the importance of domestic shocks. Larger and more closed economies are less exposed to global and foreign shocks, and external forces play a smaller role in shaping their monetary policy. In the case of the US, domestic shocks are the main drivers of the interest rate correlation, explaining 35% at long and 60% at short horizons. For small and open countries (Australia and Canada), instead, global and foreign shocks are the most influential (around 95% at long horizons, with demand shocks alone contributing more than 50%).[10] The construction of the global interest rates as GDP-weighted averages further amplifies the role of domestic shocks in large economies. Indeed, when we replace the global interest rates with US interest rates, domestic shocks only explain a large part of the correlation for the US, see Appendix Figure D.4. They contribute to a small extent for other countries (as in JP and KO) because impulse response functions in the domestic economy and the US move in the same direction. This happens mostly in reaction to the demand shocks. Importantly, the FECorDs based on US rates still support our main conclusions: the relative importance of foreign versus global shocks shifts similarly over different horizons, and the cross-country patterns are consistent with those found using the global interest rate.

The weight in the global interest rate gives also an explanation for the the difference

[10]This channel might exist in particular for commodity exporters (Australia and Canada, in our sample). In a robustness check, we test whether global supply and demand shocks create additional spillovers in these countries through exchange rate movements. Specifically, we let these two shocks load positively (supply) and negatively (demand) on the exchange rate equation. The new loadings are significantly different from zero. The importance of global shocks goes down at the expense of foreign shocks. Other basic results remain qualitatively unchanged, see Appendix Figure E.1.

Figure 5: Detailed forecast error correlation decomposition between global and country-specific interest rates



NOTES: The figure shows median forecast error correlation decomposition between global (GDP weighted) and country-specific interest rates (in subplots) to domestic (dom) and foreign (for) supply, demand, monetary policy and exchange rate shocks and global (glob) supply, demand and monetary policy shocks over 20 quarters.

between forecast error *variance* and *correlation* decompositions. In the former, domestic shocks are the most important, in the latter, global shocks contribute most for small economies. Moreover, economies with a low overall correlation to the global rate (South Korea and Japan) also have a larger contribution of domestic shocks.

As the fourth key result, we find that the most important shocks in the forecast error correlation decomposition of interest rates are demand shocks, see Figure 5. This shows that interest rates are correlated because inflation-targeting central banks react similarly to fluctuations in prices and output – a fact we investigate further in the impulse response analysis. The three different types of demand shocks jointly contribute between 45% and 85% to the correlation on impact and between 50% and 60% after 20 quarters. They are the only types of shocks where the 95% credibility sets for foreign and global shocks regularly do not include zero, see Appendix Table D.1. Monetary policy shocks, instead, account for a lower share, between 10% and 40% on impact and between 15% and 20% at long horizons.

The two largest economies in our sample also have the largest individual contributions among the foreign shocks. US demand shocks account for 50% to 60% of the aggregate contribution of foreign demand shocks after 20 quarters, while EA demand shocks account for only 5% to 15%. These two countries together also provide the most important foreign monetary policy shocks. In section 3.2.3, we document that this is not purely mechanical due to their large weight in the global interest rate, but also related to spill-overs to the other countries in the sample.

In summary, the relative importance of the drivers of the forecast error variance and forecast error correlation indicate that monetary policy is largely conducted autonomously. Central banks react synchronously to demand shocks. Additionally, we find limited evidence for direct international spillovers from foreign monetary policy shocks. We investigate these channels further in the impulse response function analysis.

3.2 Dynamic effects of structural shocks

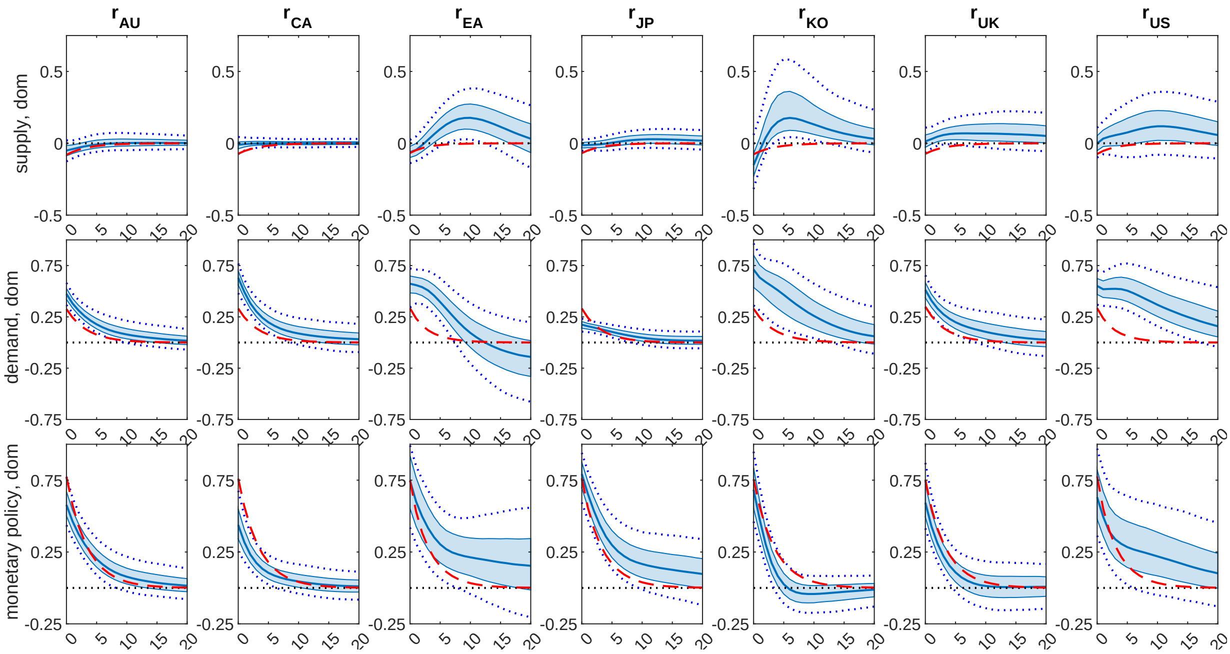
In the following, we discuss the impulse response functions to those shocks that contribute strongest to the long-run correlation of domestic interest rates to the global interest rate. These are domestic and global shocks, as well as US and EA demand shocks. In addition, we show and discuss impulse-response functions to US and EA monetary policy shocks.

3.2.1 Dynamic effects of domestic shocks

Figure 6 shows median impulse responses of country-specific interest rates (in columns) to domestic demand, supply, and monetary policy shocks (in rows) for 20 quarters together with the 68% and 95% posterior credibility sets (shaded area and dotted lines, respectively). The figure includes median responses from our informative priors (dashed red lines). The impact effects are qualitatively similar to those of the posterior IRFs. The prior median responses then die out quite quickly. Posterior dynamics, especially in the Euro Area, Korea and US, are much more persistent, speaking to the informativeness of the data.

Overall, domestic responses are in line with findings in the literature. The dynamics differ across shocks, but they are rather homogeneous across countries: Domestic demand

Figure 6: Impulse responses of interest rates to domestic shocks



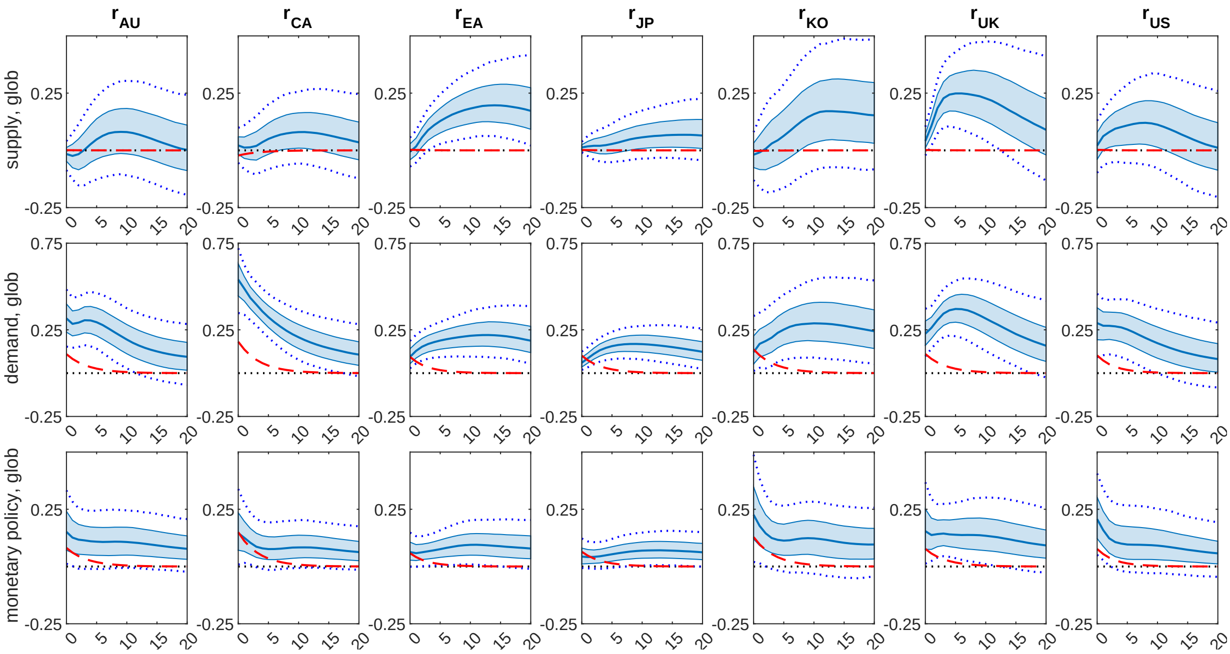
NOTES: The solid lines in the figure show median impulse responses of country-specific interest rates (in columns) to country-specific demand, supply and monetary policy shocks (in rows) over 20 quarters. The shaded areas (dotted lines) show the 68% (95%) posterior credibility sets. Dashed red lines show prior median response. To make them economically comparable, the shocks are normalized a size of one unit. Differences across shock variances are shown in Application Figure D.20.

shocks increase interest rates on impact with the peak effect for the countries within one year. Unexpected tightening in domestic monetary policy increases interest rates sharply on impact. The immediate response is smaller than the initial shock size because of the direct endogenous response of monetary policy to lower output gaps and inflation caused by the contractionary monetary policy shock. The posterior credibility bands associated with the impulse responses do not contain zero for several quarters, in part because of the significant degree of interest rate smoothing in all countries. Domestic supply shocks initially cause a small decrease or no effect in interest rates. For most economies a re-bouncing effect is visible after a few quarters.

3.2.2 Dynamic effects of global shocks

Figure 7 shows impulse responses of country-specific interest rates (in columns) to global shocks. The posterior IRFs are qualitatively similar to those generated by the prior (red dashed line) and the data are informative about the dynamics. Global demand cause persistent positive co-movement in interest rates. Global monetary policy shocks also induce positive co-movement, but credibility sets include zero immediately or after very few horizons. IRFs to global supply shocks are mixed, and credibility sets mostly include zero.

Figure 7: Impulse responses of interest rates to global shocks



NOTES: The solid lines in the figure show median impulse responses of country-specific interest rates (in columns) to global demand, supply and monetary policy shocks (in rows) over 20 quarters. The shaded areas (dotted lines) show the 68% (95%) posterior credibility sets. Dashed red lines show prior median response. The posterior shocks have size of one unit. Prior IRFs are scaled for comparison.

The impulse response functions explain the differences in contributions to interest rate correlations. The importance of global demand shocks is due to both strong and persistent homogeneous interest rate responses. These co-movements are caused by endogenous monetary policy reactions to common changes in the domestic output gap and inflation, which lead to quasi-synchronous changes in interest rates. Appendix Figures D.12 and D.13 show that output gaps and inflation increase on impact after a global demand shock, and that the increase lasts for at least five quarters. In reaction to the stimulated economy and increased price levels, central banks increase interest rates for five to 15 quarters. Our findings are in line with Mumtaz and Surico (2009) and Charnavoki and Dolado (2014) who use structural factor models to study the effect of global supply and demand shocks (identified via sign or exclusion restrictions) on global activity, inflation, and commodity prices.

Movements caused by global monetary policy shocks are also comparable across the countries in our sample:[11] Interest rates increase on impact. Therefore, we find that global monetary policy shocks explain a small share of the forecast error correlations in the short run. The effects are not persistent, since credibility sets quickly include zero – not only for interest rates, but also for output gaps and inflation, see Appendix Figures D.12 and D.13. Consistent with the low persistence of impulse response functions, the contribution of global monetary policy shocks to forecast error correlations drops over the horizons.

Compared to global monetary policy and demand shocks, global supply shocks cause limited responses of interest rates. Short-run responses are mixed and statistically not different from zero. Medium-term responses to global supply shocks are positive due to spillovers via aggregate demand, in particular for the Euro Area and UK: as output increases in all countries after a global supply shock, so does *foreign* output, which shifts domestic aggregate demand endogenously outward and causes monetary tightening, see the posterior distribution of structural contemporaneous coefficients β^{dc,y^*} and $\psi^{c,y}$, $\psi^{y,\pi}$ in Figure 12 and Appendix Figure D.6. However, on impact the exogenous shock to supply curves dominates, which explains the visible delay in interest rate responses.

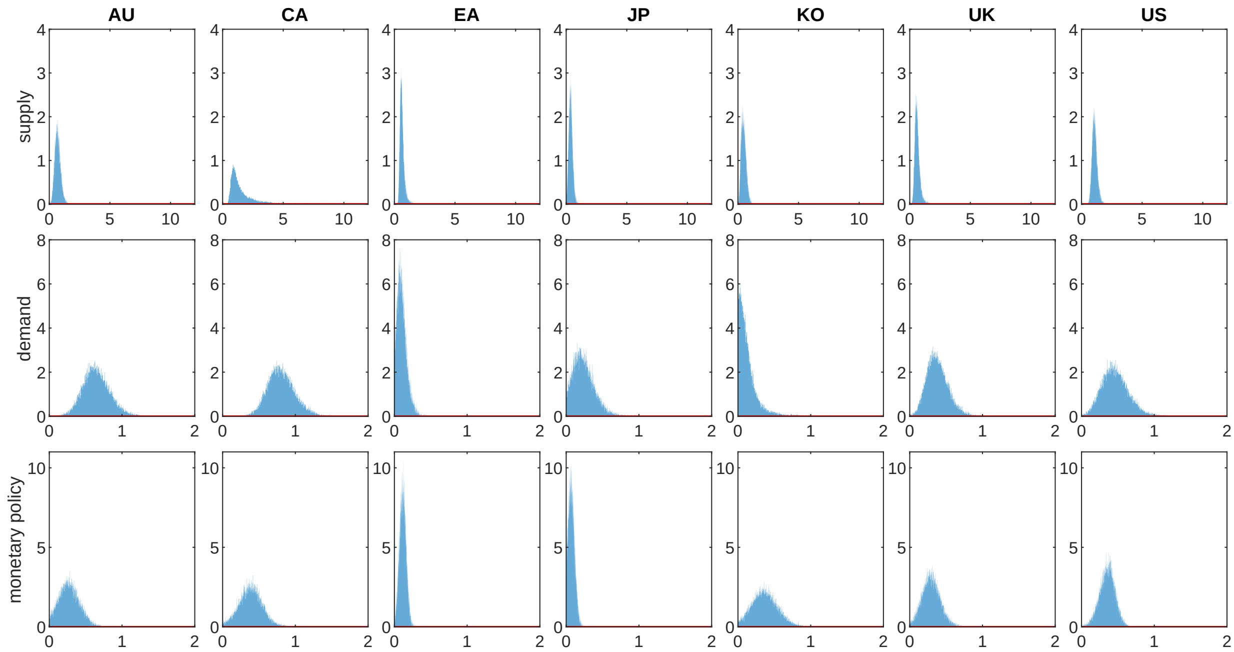
For all global shocks, the impact response to global shocks is muted compared to domestic shocks. The estimated loadings, shown in Figure 8, can explain differences in the magnitude of the response across countries. The data are very informative for loadings, and the sign restrictions are not binding in most cases.

3.2.3 Dynamic effects of foreign shocks

In this section, we focus on discussing the effects of demand and monetary policy shocks originating in the US and EA, shown in Figure 9. The demand shocks create in general the largest international reactions, and monetary policy shocks are included because of

[11] This result complements empirically Georgiadis and Jančoková (2020)'s finding that monetary policy shocks identified in empirical and theoretical single-country models are positively correlated across borders.

Figure 8: Posterior distributions of loadings



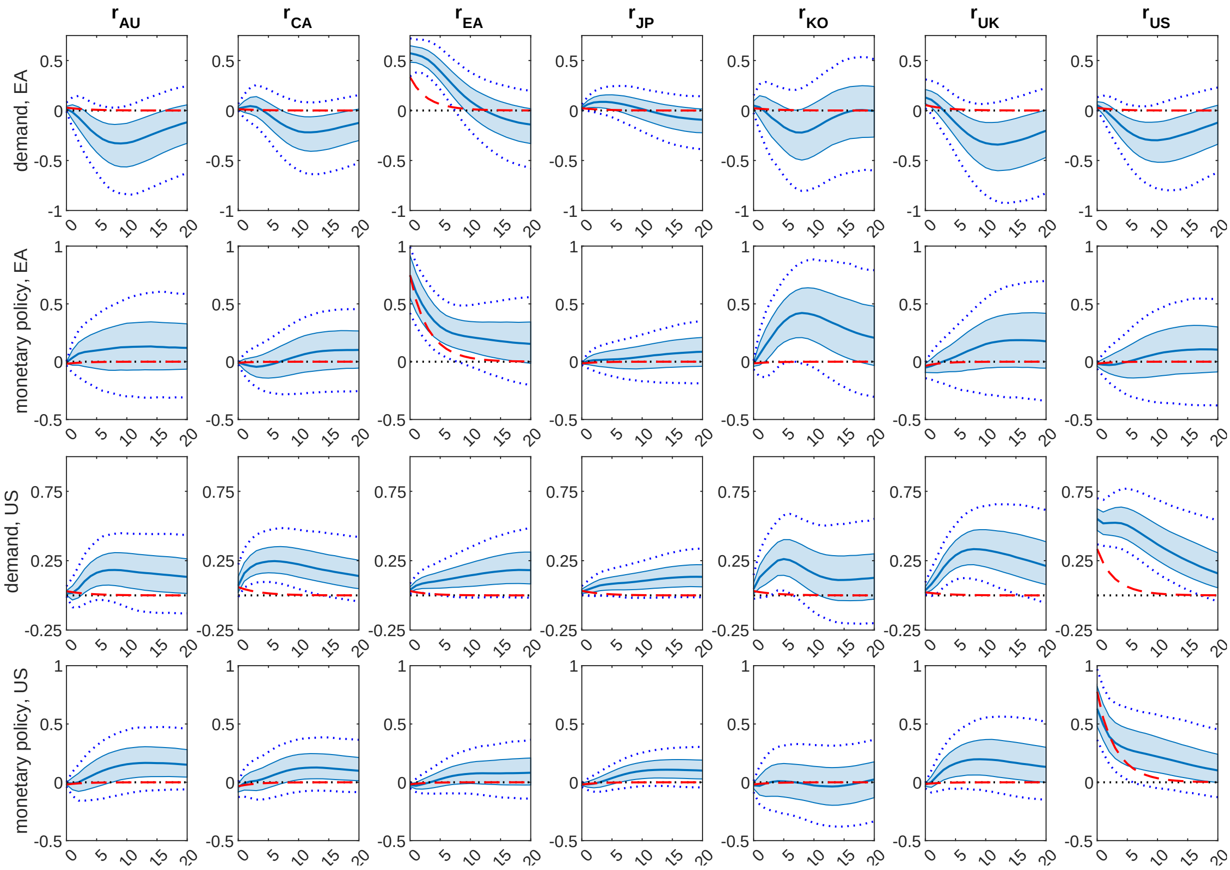
NOTES: The histograms show the posterior distribution of global shock loadings together with the uninformative prior distribution (red line).

the extensive interest in the existing literature. Prior impulse responses (dashed red lines) imply negligible spillovers. The shape of the posterior responses differs clearly from the prior responses.

Similar to the effects of global shocks, US demand shocks increase interest rates in other countries with the peak effect occurring after around five horizons. Euro Area shocks, on the other hand, cause a medium-run drop in interest rates in other countries. The reason for these differences stems from different responses of inflation and output: US demand shocks cause inflation to increase in almost all countries, and output to increase in most, while EA demand shocks instead induce medium run drops in other countries output.

Monetary policy shocks from the EA and US do not cause credibly identified international reactions. Response are positive but not different from zero (with higher uncertainty for responses to EA shocks). The main reason is because the US and EA output gaps recover relatively quickly after a domestic shock, reducing the negative spillover effects substantially (indeed, some countries even experience a small boom after 1-3 years). Another reason could be attempts by other central banks to counteract potential capital

Figure 9: Impulse responses of interest rates to US and EA demand and monetary policy shocks



NOTES: The third (seventh) column in the first (last) two rows show *domestic* responses, the other ones *foreign* responses. The solid lines in the figure show median impulse responses of country-specific interest rates (in columns) to US and EA demand and monetary policy shocks (in rows) over 20 quarters. The shaded areas (dotted lines) show the 68% (95%) posterior credibility sets. Dashed red lines show prior median response. The shocks have size of one unit.

outflows (Miranda-Agrippino and Rey, 2020).

Notably, Australian and Canadian interest rates do not react much to US monetary policy shocks but their exchange rates do, see Appendix Figure D.17. The two countries experience a peak decrease in competitiveness after around two quarters.

While a dominant transmitter role of the US is stressed in the literature (such as Miranda-Agrippino and Rey, 2020; Rey, 2016; De Santis and Zimic, 2022), our results give a more diverse picture, as they highlight the role of global shocks as well as the differences of responses to shocks originating in other countries. These reactions should caution against block-exogeneity assumptions in VAR models. Moreover, our results relate to studies reporting similarities in spillover effects and mutual reactions caused by unconventional monetary policy shocks of the Fed and ECB (such as Curcuru et al., 2018; Miranda-Agrippino and Nenova, 2022; Jarociński, 2022). However, it should be noted that a generalization of our findings towards other currency areas might not be possible. Especially for the case of emerging economies, one would need to pay extra attention to the role of exchange rate arrangements and capital controls, which could imply very different dynamics to the ones we document.

3.3 What are global shocks?

The global shocks are essential drivers of economic developments. To underline their crucial role, we evaluate whether the data support including global shocks. To that end, we compare the model fit of the model with global shocks to an alternative excluding global shocks via the Savage-Dickey density ratio (Verdinelli and Wasserman, 1995). A model without global shocks is a restricted version of the model with global shocks. Integrating out the global shocks, the restrictions imply that $\chi_c^j = 0$ for all countries and equations. As in Chan (2018) the Bayes factor in favor of the model with global shocks (unrestricted model) is calculated using the log Savage-Dickey density ratio as:

$$\log(SDDR) = \log\left(\frac{p\left(\chi = \mathbf{0}\right)}{p\left(\chi = \mathbf{0}|\mathbf{Y}_T\right)}\right),$$

defined as the marginal prior density relative to the marginal posterior density evaluated at the restriction. The majority of loadings is clearly different from zero, see also Figure 8. Accordingly, we find $\log(SDDR) = 638$, strongly supporting the inclusion of global shocks.

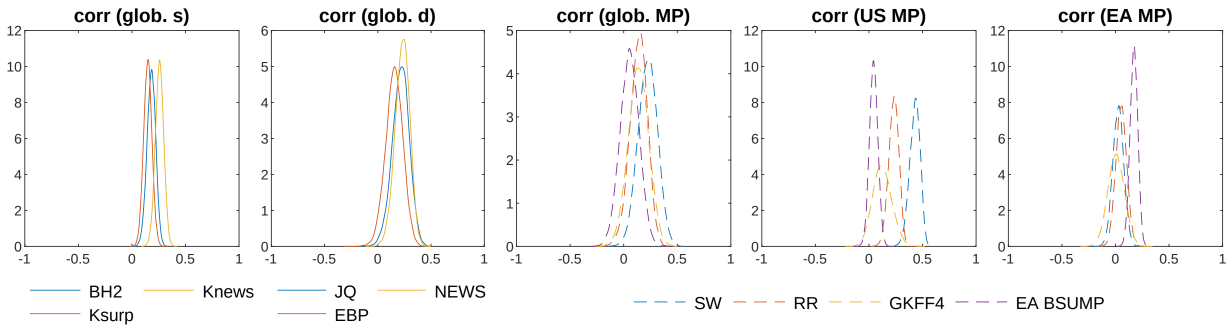
But how can we interpret global shocks? Our global supply shocks are linked to oil supply shocks but also capture global supply changes unrelated to oil. For example, the Venezuela oil strike (December 2002), the start of the Iraq War (March 2003), and the Libyan Civil War (February 2011) are associated with negative global supply shocks, see the identified global shock series in Appendix Figure D.21. Moreover, the accession of China to the WTO and the Fukushima nuclear disaster with the associated disruption of global supply chains coincide with supply shocks that are clearly different from zero.

The nuclear disaster in Chernobyl (April 1986) and the terrorist attacks on 9 September 2001 coincide with strong negative demand shocks. A series of large negative global demand shocks start with the collapse of Lehman Brothers (September 2008). The largest negative global demand shock is timed at the peak of the global financial crisis in 2009. Hence, we think that large shocks to uncertainty and credit supply shocks causing disruptions to worldwide financial markets may be interpreted in our model as a negative global demand shock. This finding remains across all robustness checks.

We include global monetary policy shocks because monetary policy shocks might be internationally correlated (Georgiadis and Jančoková, 2020). These shocks significantly improve the statistical fit of the model with a log SDDR of roughly 50. However, 95% credibility sets of the global monetary policy shocks seldom exclude zero. Thus, our data do not provide strong evidence of unexpected joint monetary policy action. Two events deserve mentioning: First, the largely unexpected stock market crash on October 19, 1987, causing worldwide losses in stock prices, coincides with negative global monetary policy shocks in that quarter. Second, on October 8, 2008, central banks (Bank of Canada, the Bank of England, the European Central Bank, the Federal Reserve, Sveriges Riksbank, and the Swiss National Bank, with support of Bank of Japan) issued a joint policy statement and cut collectively interest rates by 50 basis points surprising financial markets (Badinger and Schiman, 2023). At the quarterly frequency, interest rates in all countries decreased by much more (between 1 and 2 percentage points, with the exception of Japan). However, this quarter coincides with the height of the financial crisis, which is best captured by a global demand shock.

To investigate these discussed links further, we calculate the correlation (at each posterior draw) of global supply and demand shocks to prominent shock and instrument series for oil supply shocks and credit supply shocks, respectively, see the first two subplots in Figure 10. We find a positive correlation of our global supply shocks to exogenous expansions of oil supply. Likewise, our global demand shock is positively correlated to proxies of expansionary credit supply shocks. Overall, the correlations with our global shocks do not exceed 0.5 for all measures, which indicates that our global shocks also capture additional

Figure 10: Correlation of global shocks, and US and EA monetary policy shocks with common shock proxies from the literature



NOTES: The histograms show the correlation of the posterior draws of the global shocks with various shock proxies from the literature, always multiplied such that the proxy is expansionary. *BH2*: oil supply shocks of Baumeister and Hamilton (2019); *Ksurp*: oil supply expectation shocks of Känzig (2021); *Knews*: oil supply news instrument of Känzig (2021) based on high-frequency changes in oil futures prices around OPEC production announcements; *JQ*: innovations to the financial conditions index of Jermann and Quadrini (2012); *EBP*: excess bond premium of Gilchrist and Zakrajšek (2012); *NEWS*: textual proxy series of Mumtaz et al. (2018) counting the words "credit crunch" and "tight credit"in nine US newspapers; *SW*: monetary policy shocks from the DSGE model of Smets and Wouters (2007); *RR*: Romer-Romer-type instrument provided by Coibion et al. (2017); *GKFF4*: high-frequency instrument from Gertler and Karadi (2015); *EA BSUMP*: EA monetary policy shocks identified through narrative sign restrictions in Badinger and Schiman (2023).

worldwide developments.

The last three subplots in Figure 10 show the correlation of global, US and EA monetary policy shocks to three common instruments for US and one instrument for EA monetary policy shocks. The domestic shocks show the expected correlation. Our global monetary policy shocks are only positively correlated to US monetary policy shocks from the DSGE-model by Smets and Wouters (2007) (blue dashed line). However, the correlation is much smaller than the one to US shocks. Moreover, Smets and Wouters (2007) model a closed economy, that is, they do not differentiate between global and local shocks like we do (Georgiadis and Jančoková, 2020).

3.3.1 Differences to the global financial cycle

Our global demand shocks are related to worldwide disruptions in the financial sector. Rey (2015), Miranda-Agrippino and Rey (2020), and Miranda-Agrippino and Rey (2022), among others, stress that these disturbances are reasons for co-movements in risky assets, called the "global financial cycle" (GFC). This might imply that co-movements in financial variables are shifted in our model to the global demand shock. However, residuals from an AR(4) model of Miranda-Agrippino and Rey (2020)'s GFC factor are not correlated to any of the global shocks. Yet, there is a negative correlation to US monetary policy shocks.

Augmenting our baseline model with the GFC factor of Miranda-Agrippino and Rey (2020) as an additional endogenous variable isolates the influence of the global financial cycle on country-specific developments.[12] Two pieces of evidence indicate that while both our global demand shocks and the GFC are related to financial disturbances, the two are not equivalent as the global demand shock captures demand disruptions more broadly. First, a residual shock to the GFC factor has hardly any contribution to the forecast error correlations of interest rates to the global interest rate after five years. Moreover, the country-specific loadings of the global demand shock are of similar size compared to the model without the GFC factor, Appendix Figure E.6. Second, the impulse response functions to a residual GFC shock are not consistent with a global demand shock, see Appendix Figure E.7. Growth rates of real effective exchange rates fall strongly on impact in almost all countries. Most other responses are zero or contain zero in their credibility sets. The exception is the US, where output increases and inflation falls on impact.

3.3.2 Differences to US shocks

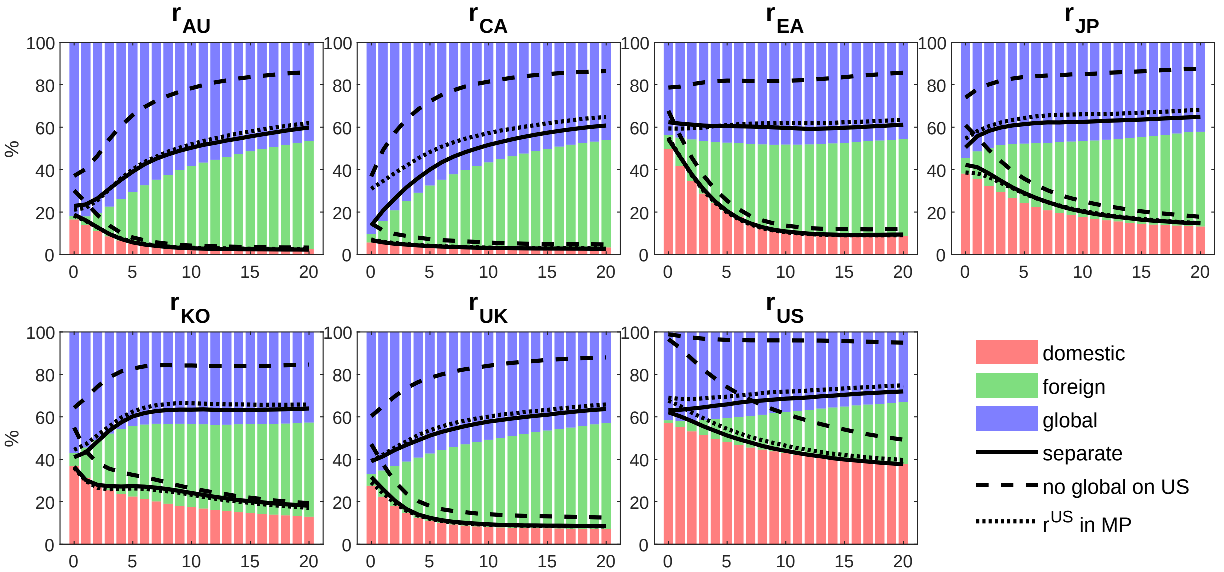
The economic size and importance of the US begs the question whether our global shocks are simply observationally indistinguishable US shocks. To investigate this, we consider three alternative specifications where we attribute a special role to the US.

First, we relax the homogeneity assumption and allow structural coefficients $\mathbf{A}$ on US variables to be different from other foreign variables. We find that allowing for a unique role of direct spillovers from the US to other countries does not change the importance of global shocks, see solid lines labeled *separate* in Figure 11.

Second, we assume that global shocks do not directly influence US developments, by

[12]We use the quarterly average of the GFC factor updated to 2019Q1, as provided on `http://silviamirandaagrippino.com/code-data` [accessed 26 June 2023]. The GFC factor is contemporaneously included in all baseline equations, and all variables contribute contemporaneously to the development of the GFC factor in the additional equation. We set relatively wide priors on the additional contemporaneous parameters, Student t priors with mode zero and scale one.

Figure 11: Forecast error correlation decomposition between global and country-specific interest rates for alternative US specifications



NOTES: Figure shows the forecast error correlation decomposition between global and country-specific interest rates for the baseline model (bars) together with three alternative US specifications (lines).

setting US loadings to zero. This implies, conversely, that unexpected US developments are not contained in global shocks. By construction the contribution of global shocks to the correlation of global and US interest rates is extremely close to zero, as they affect the US economy only through their influence on foreign economic development. For the remaining countries foreign shocks gain importance while global shocks loose compared to the baseline model, see the dark dash-dotted lines in Figure 11, labeled *no global on US*. It shows that our global shocks are not only driven by US developments but are less important as the part of the US shocks correlated with other countries' shocks contributes strongly to the global shocks. Removing US loadings does not nullify the importance of global shocks, as the posterior distributions of other loadings are all positive, with posterior mass not pressed against zero.

Third, we allow for a dominant US central bank by including US interest rates directly in the policy rules of other countries, thus allowing for an additional direct spillover channel. The prior distribution of the new structural parameter $\psi^{c,r^{US}}$ follows a t-distribution with mean zero, scale 0.4 and 3 degrees of freedom. More than 80% of the posterior mass of this coefficient is negative for the EA, consistent with monetary policy anticipating a cooling of the economy after an interest rate increase in the US. Above 90% of the posterior is positive for CA, implying a central bank reaction more in-line with capital flow pressures. For AU, JP, KO, and UK, the coefficient is centered around zero. The contribution to the forecast error correlations of interest rates is nearly identical to the baseline, indicating robustness of our model in that regard, see dark dotted lines in Figure 11 labeled r^{US} *in MP*.

Overall, our tests indicate that US developments are important parts of identified global shocks due to its sheer size, but that the US is not "special" enough to be treated differently from the other countries in our sample.

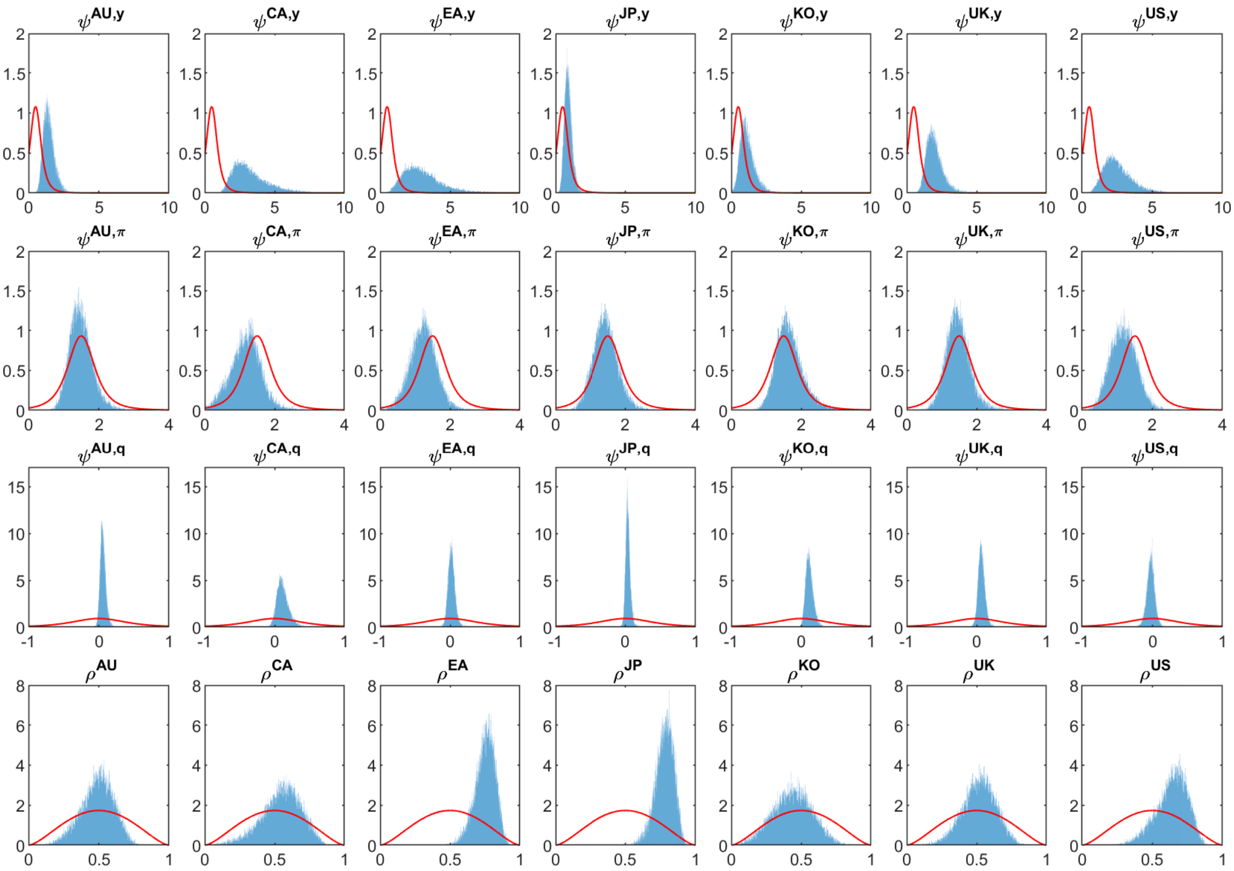
3.4 The monetary policy rule

We document above that only between 15% and 20% of the long-run correlation between country-specific and global interest rates are due to domestic, foreign and global monetary policy shocks. The remainder is therefore driven by the systematic response of monetary policy to other structural shocks.

Figure 12 shows the prior and posterior distributions of the structural contemporaneous coefficients of the monetary policy rules. The usual Taylor rule coefficients $(\psi^{c,y}, \psi^{c,\pi}, \rho^c)$ are all consistent with our priors. The comparably aggressive central bank reaction to output gaps is similar to findings of Baumeister and Hamilton (2018). The additional coefficient on exchange rate fluctuations $\psi^{c,q}$ is very precisely identified to be zero for the two largest economies (Euro Area and US). It is extremely small but positive for the remaining smaller economies. This indicates that the central banks in our sample largely conduct monetary policy according to a traditional mandate to reduce domestic inflation and output gaps. That is, interest rates movements are internationally correlated because business cycle fluctuations are. Thus, the main cause of international business cycle correlations – different types of demand shocks – are also the main cause of interest rate co-movement.

Relaxing the restrictions that monetary policy endogenously reacts only to contemporaneous changes in foreign interest rates, output and inflation does not significantly change

Figure 12: Posterior distributions of structural contemporaneous coefficients, monetary policy rule



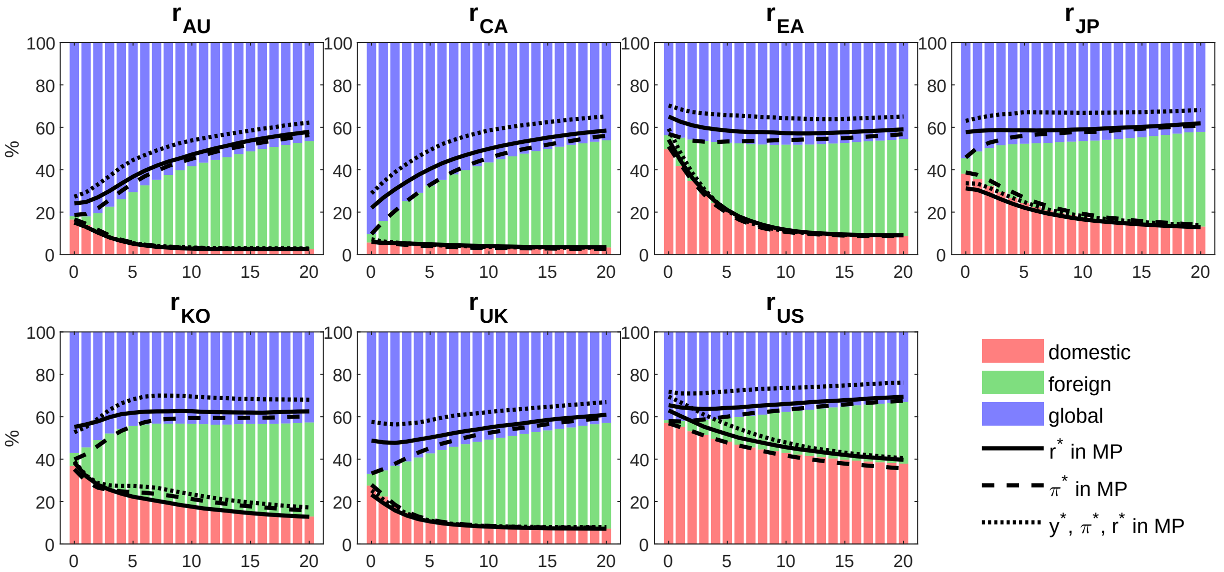
NOTES: The histograms show the posterior distribution of structural contemporaneous coefficients in the monetary policy equation together with the prior distribution (red line).

our results, see Figure 13. We extend the monetary policy rules to contain coefficients on foreign interest rates (labeled r^* *in MP*), foreign inflation (labeled π^* *in MP*) and foreign output gaps, inflation and interest rates (labeled y^*, π^*, r^* *in MP*).

The main difference to the baseline results is an increase in the importance of foreign shocks to around 20% at short horizons for UK and JP in the first (solid lines) and third (dotted lines) variant, at the expense of domestic shocks. This is because the coefficients on foreign interest rates in the British and Japanese monetary policy rule are both economically and statistically significantly positive. All other additional coefficients in the monetary policy rules are identified to be zero.

Our results are also robust to the explicit inclusion of a fifth variable capturing financial transmission channels. We discuss this in Appendix Section E.4.

Figure 13: Forecast error correlation decomposition between global and country-specific interest rates for alternative monetary policy rule specifications



NOTES: Figure shows the forecast error correlation decomposition between global and country-specific interest rates for the baseline model (bars) together with three alternative monetary policy rule specifications (lines).

3.5 Robustness checks

We check robustness to alternative specifications of aggregate supply and the exchange rate equations. To that end, we first add foreign output gaps in the supply equation, restricting the coefficient to be positive. Second, we adjust the exchange rate equation such that interest rate differential enter. Hence, we model directly the uncovered interest rates parity such that differences in interest rates between the countries should equalize relative changes in exchange rates, potentially strengthening the exchange rate channel for the transmission of monetary policy shocks. Results do not change considerably: the role of global shocks decreases marginally, when we include foreign output in the supply curve, and increases slightly under the assumption of UIP. Once we restrict our sample to the 21st century (following a recent argument of Engel and Wu, 2024), global shocks become much more important, decreasing the contribution of domestic shocks, see Appendix Figure E.4.

Moreover, our findings are not sensitive to the use of shadow rates in two robustness check that (a) use data only until 2007:Q3 only and (b) additionally exclude Japan, which hit the zero lower bound already in 1995Q3, see Appendix Figure E.3. We also test whether using log differences of real GDP instead of output gaps changes our main findings. It does so marginally, lowering the contribution of foreign shocks and increasing the importance of global shocks, see Appendix Figure E.2. Finally, we run a robustness check starting only in 1999, and find that global shocks become significantly more important, mostly at the expense of domestic shocks, see Appendix Figure E.3.

4 Conclusion

Central bank interest rates are internationally correlated. These correlations may stem from direct spillovers of monetary policy shocks from a dominant economy. Alternatively, monetary policy may respond systematically to common business cycle fluctuations, consistent with mandates focused on domestic inflation and output gaps.

To quantify the drivers of interest rate forecast error variance and cross-country correlations, we develop a flexible Bayesian structural panel VAR model. Our framework includes supply, demand, monetary policy, and exchange rate equations for seven advanced economies, with informative priors on contemporaneous coefficients. Country-specific policy rule parameters provide evidence on the systematic, rule-based conduct of monetary policy. We extend the model using a factor structure to distinguish between domestic, foreign, and global structural shocks. The relative importance of monetary policy shocks informs the question of whether direct spillovers are present.

Our results provide evidence that interest rates move together internationally because central banks systematically respond to correlated domestic fluctuations. These fluctuations are predominantly driven by domestic, foreign, and global demand shocks. Monetary policy does not respond directly to foreign monetary policy shocks. Hence, our findings imply that central banks in the economies of our sample act autonomously.

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Online Appendix to "What explains international interest rate co-movement?"

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July 24, 2025

A Data

Table A.1: Data

Variable	Description
	baseline VAR variables
y	output gap; source: Oxford Economics, datastream codes AUXOGAP.R, CNXOGAP.R, EKXOGAP.R, JPXOGAP.R, KOXOGAP.R, UKXOGAP.R, USXOGAP.R
π	year-on-year inflation rate calculated from consumer prices (all items); source: IMF-IFS and Eurostat, Datastream codes AUI64...F, CNI64...F, EMCONPRCF, JPI64...F, KOI64...F, UKI64...F, USI64...F
r	quarterly average of monthly policy rates and Krippner shadow rates; sources: https://www.ljkmfa.com/ and national central banks, Datastream codes (dates of shadow rate) AUPRATE., CNB14044, EMREPO.. (02/2009-12/2019), JPPRATE. (09/1995-12/2019), KOI60B.., UKPRATE. (12/2008-10/2017), FREFEDFD (11/2008-07/2016)
q	real effective exchange rate (narrow), defined as nominal effective exchange rate times the ratio of a weighted sum of foreign price indices relative to a domestic price index; source: BIS (we are using the inverse of the data), https://www.bis.org/statistics/eer.htm
	Variables for VAR extensions
sp	stock prices, yoy growth rates; source: Thomson Reuters, Datastream codes: AUSHRPRCF, CNSHRPRCF, EMSHRPRCF, JPSHRPRCF, KOSHRPRCF, UKSHRPRCF, USSHRPRCF
ts	term spread, long-term bond yields minus r; source (bond yields): Thomson Reuters, Datastream codes AUGBOND., CNGBOND., EMGBOND., JPGBOND., KOGBOND., UKGBOND., USGBOND.
$GDP(real)$	use to calculate *output growth* and *GDP weights*; source: OECD Quarterly National Accounts, Oxford Economics, Datastream codes: AUOEXP03D, CNOEXP03D, EKXGDSA.D, JPOEXP03D, KOOEXP03D, UKOEXP03D, USOEXP03D
	External instruments, used to cross-check with global shocks
BH2	Median of oil supply shocks identified from Baumeister and Hamilton (2019), updated to December 2022, https://sites.google.com/site/cjsbaumeister/research
Ksurp	monthly oil supply surprise shocks from Känzig (2021), updated to December 2022, https://github.com/dkaenzig/oilsupplynews
Knews	monthly oil supply news shocks from Känzig (2021), updated to December 2022, https://github.com/dkaenzig/oilsupplynews
JQ	Innovations to the financial conditions index of Jermann and Quadrini (2012) as used in Mumtaz et al. (2018)
EBP	Excess bond premium of Gilchrist and Zakrajšek (2012) as used in Mumtaz et al. (2018)
NEWS	Textual proxy for credit supply shocks, from Mumtaz et al. (2018)
SW	US Monetary policy shocks from Smets and Wouters (2007)
RR	US Romer-Romer monetary policy shocks, Romer and Romer (2004)
GKFF4	US High-frequency monetary policy shocks from Gertler and Karadi (2015)
EA BSUMP	EA High-frequency and narrative monetary policy shocks, Badinger and Schiman (2023)

B Theoretical model for identifying restrictions

Following Lubik and Schorfheide (2007), we consider a model with an international Phillips curve, IS curve, monetary policy rule, and (real) exchange rate equation, to which we add common shifts of country-specific supply and demand curves in the form of global supply and demand shocks. The set of equations for country $c \in \{1, \ldots, C\}$ is given by:

$$y_{ct} = \frac{\tilde{\tau}^c}{\kappa^c}\left[\mu^{c,s} + \pi_{ct} - \beta^c E[\pi_{ct+1}] + \frac{\alpha^c \beta^c}{1-\alpha^c} E[q_{ct+1}] - \frac{\alpha^c}{1-\alpha^c} q_{ct}\right] + \chi_c^s u_{gt}^s + u_{ct}^s \tag{B.1}$$

$$\begin{aligned} y_{ct} =& \mu^{c,d} + E[y_{ct+1}] - \tilde{\tau}^c \left(r_{ct} - E[\pi_{ct+1}]\right) + \alpha^c(2-\alpha^c)\frac{1-\tau^c}{\tau^c} E[y^*_{ct+1}] + \ldots \\ &+ \frac{\alpha^c \tilde{\tau}^c}{1-\alpha^c} E[q_{ct+1}] + \chi_c^d u_{gt}^d + u_{ct}^d \end{aligned} \tag{B.2}$$

$$r_{ct} = \rho^c r_{it-1} + (1-\rho^c)\left[\left(\psi^{c,1} + \psi^{c,3}\right)\pi_{ct} + \psi^{c,2} y_{ct} + \psi^{c,3}\left(q_{ct} - \pi^*_{ct}\right)\right] + \chi_c^m u_{gt}^m + u_{ct}^m \tag{B.3}$$

$$q_{ct} = \mu^{c,q} - \frac{1-\alpha^c}{\tilde{\tau}^c}\left(y^*_{ct} - y_{ct}\right) + u_{ct}^q \tag{B.4}$$

with

$$\tilde{\tau}^c = \tau^c + \alpha^c(2-\alpha^c)(1-\tau^c), \quad y^*_{ct} = \sum_{c^* \neq c} w_{cc^*} y_{c^*t}, \quad \pi^*_{ct} = \sum_{c^* \neq c} w_{cc^*} \pi_{c^*t},$$

Equation (B.1) expresses the open economy Phillips curve. Supply depends on inflation and changes in the exchange rate. The parameter α^c, $0 < \alpha^c < 1$, measures the import share. When $\alpha^c = 0$, the model reduces to a closed economy set-up. τ^c gives the intertemporal substitution elasticity, β^c the discount factor, and κ^c the slope coefficient of the Phillips curve.

Equation (B.2) models the open economy IS curve. Demand is expressed as a function of interest rates, inflation, foreign output, and changes in the exchange rate. We substitute expectations in equations (B.1) and (B.2) with simple autoregressive forecasts, following Baumeister and Hamilton (2018). We model the expected value of a variable z as $z_{t+1|t} = c^z + \phi^z z_{t|t}$. We set the autoregressive parameter equal for all variables, as $\phi^c = 0.75$.[1] The term c^z is absorbed in the constant terms. The parameter ζ^c weights expected output in the IS curve. The Phillips and IS curve can then be expressed as

$$y_{ct} = \frac{\tilde{\tau}^c}{\kappa^c}\left[\mu^{c,s} + (1-\beta^c\phi^c)\left[\pi_{ct} - \frac{\alpha^c}{1-\alpha^c} q_{ct}\right]\right] + u_{ct}^s \tag{B.5}$$

$$y_{ct} = \frac{1}{1-\zeta^c\phi^c}\left[\mu^{c,d} - \tilde{\tau}^c\left(r_{ct} - \phi^c\pi_{ct}\right) + \alpha^c(2-\alpha^c)\frac{1-\tau^c}{\tau^c}(\phi^c - 1) y^*_{ct} + \frac{\alpha^c\tilde{\tau}^c\phi^c}{1-\alpha^c} q_{ct}\right] + u_{ct}^d \tag{B.6}$$

[1] Note that Lubik and Schorfheide (2007) include in some equations expected changes in variables (as opposed to expected values of the variables). In such cases, we model the expected value of a change in variable z_{t+1}, denoted by Δz_{t+1}, as $E(\Delta z_{t+1}) = (0.75 - 1) z_t$.

The monetary policy authority sets interest rates according to the rule given in equation (B.3). The parameter $\psi^{c,1}$ captures the response of the monetary policy authority to changes in inflation, $\psi^{c,2}$ reflects the reaction to output, and $\psi^{c,3}$ to changes in the nominal exchange rate. ρ^c is a smoothing parameter, smoothing the implementation of monetary policy over time. Equation (B.4) relates changes in the exchange rate to differences in foreign and domestic output. We use that under PPP Lubik and Schorfheide (2007) express inflation as $\pi_{ct} = ex_{ct} + (1-\alpha^c)tot_{ct} + \pi^*_{ct}$ where ex_{ct} are changes in the nominal exchange rate and tot_{ct} changes in terms of trade. Thus, the real exchange rate relates to terms of trades as $tot_{ct} = -\frac{1}{1-\alpha^c} q_{ct}$.

Our coefficients in the panel SVAR model given in equations (s) to (er) are related to the structural parameters in the theoretical model in the following way:

$$
\begin{aligned}
\alpha^{c,\pi} &= \frac{\tilde{\tau}^c}{\kappa^c}(1-\beta^c\phi^c), \quad \alpha^{c,q} = -\frac{\tilde{\tau}^c}{\kappa^c}(1-\beta^c\phi^c)\frac{\alpha^c}{1-\alpha^c} \\
\beta^{c,r} &= -\frac{1}{1-\zeta^c\phi^c}\tilde{\tau}^c, \quad \beta^{c,\pi} = \frac{1}{1-\zeta^c\phi^c}\tilde{\tau}^c\phi^c \\
\beta^{c,y^*} &= \frac{1}{1-\zeta^c\phi^c}\alpha^c(2-\alpha^c)\frac{1-\tau^c}{\tau^c}(\phi^c-1), \quad \beta^{c,q} = \frac{1}{1-\zeta^c\phi^c}\frac{\alpha^c\tilde{\tau}^c\phi^c}{1-\alpha^c} \\
\psi^{c,\pi} &= \psi^{c,1} + \psi^{c,3}, \quad \psi^{c,y} = \psi^{c,2}, \quad \psi^{c,q} = \psi^{c,3} \\
\theta^{c,y} &= \frac{1-\alpha^c}{\tilde{\tau}^c}.
\end{aligned}
$$

C Econometric model

In this Appendix, we describe our econometric approach, and our algorithm (Metropolis-within-Gibbs) in more detail. First, we show how homogeneity restrictions simplify identification of the matrix of structural contemporaneous coefficients $\mathbf{A}$. Second, we explain the flexibility on the prior structure that we gain through hierarchical priors for structural variances $\mathbf{D}$ (Giannone et al., 2015). We can draw from the hierarchical priors using the same Metropolis-Hastings-Algorithm as the elements of $\mathbf{A}$, which simplifies our algorithm significantly. Third, we show how global shocks and their loadings can be included in the Gibbs sampler as in Korobilis (2022).

For ease of reading, we repeat the mathematical notation. As our baseline, we estimate a Bayesian structural panel VAR with global shocks using $J = 4$ variables from a panel of $C = 7$ countries. In matrix notation, our model is

$$
\begin{aligned}
\mathbf{A}\mathbf{y}_t &= \mathbf{B}\mathbf{x}_{t-1} + \chi\mathbf{u}_{gt} + \mathbf{u}_t \\
\mathbf{u}_t &\sim \mathcal{N}(\mathbf{0}, \mathbf{D}), \quad \mathbf{u}_{gt} \sim \mathcal{N}(\mathbf{0}, \mathbf{I}_G).
\end{aligned}
$$

The model has $n = CJ$ equations, jointly indexed by subscript $c \in \{1, \ldots, C\}$ and superscript $j \in \{s, d, m, er\}$. Variables are indexed by $i \in \{y, \pi, r, q\}$. The $[G \times 1]$ vector

$\mathbf{u}_{gt} = (u^s_{gt}, u^d_{gt}, u^m_{gt})$ contains global supply, demand and monetary policy shocks shocks. The loadings χ^j_c for structural equation j in country c are stacked in the $[n \times G]$ loading matrix χ:

$$\chi = \left[\begin{pmatrix} \text{diag}(\chi_1) \\ \mathbf{0}_{J-G\times G} \end{pmatrix}' \quad \begin{pmatrix} \text{diag}(\chi_2) \\ \mathbf{0}_{J-G\times G} \end{pmatrix}' \quad \cdots \quad \begin{pmatrix} \text{diag}(\chi_C) \\ \mathbf{0}_{J-G\times G} \end{pmatrix}' \right]'$$

with $\chi_c = (\chi^s_c, \chi^d_c, \chi^m_c)$.

Because the structural shocks of the model are assumed to be mutually independent (i.e., $\mathbf{D}$ is assumed to be diagonal), it is worthwhile to write down the individual structural equations. Let $\mathbf{a}^j_c$ and $\mathbf{b}^j_c$ be the $(1 \times n)$- and $(1 \times k)$- dimensional row vectors of structural coefficients (contemporaneous and lagged) in structural equation j of country c, and d^j_c the variance of the corresponding structural domestic shocks. Differentiating between equations with and without global shocks, the structural equation j in country c is

$$\mathbf{a}^j_c \mathbf{y}_t = \begin{cases} \mathbf{b}^j_c \mathbf{x}_{t-1} + \chi^j_c u^j_{gt} + u^j_{ct}, & u^j_{ct} \sim \mathcal{N}(0, d^j_c), \quad u^j_{gt} \sim \mathcal{N}(0,1) \quad \text{, if } j \in \{s, d, m\} \\ \mathbf{b}^j_c \mathbf{x}_{t-1} + u^j_{ct}, & u^j_{ct} \sim \mathcal{N}(0, d^j_c) \quad \text{, if } j \in \{er\} \end{cases}.$$

C.1 Identification of the global shocks

Conditional on $\mathbf{A}$, we assume that the structural shocks have a "combined" representation $\breve{u}^j_{ct} = \chi^j_c u^j_{gt} + u^j_{ct}$. To identify the latent global shocks, u^j_{gt}, and their loadings, χ^j_c, we assume the following. First, we let u^j_{gt} follow a standard normal distribution and restrict the correlation across global shocks to be zero (as done in static factor models). Second, following Korobilis (2022), we impose sign restrictions on loadings $\chi^j_c > 0$, which are used to identify economically interpretable global shocks, and make the structural factor model more explicit.

In the reduced form the combined shocks $\breve{u}^j_{ct}$ feature a full covariance matrix,

$$\mathbf{A}^{-1}\chi\chi'(\mathbf{A}^{-1})' + \mathbf{A}^{-1}\mathbf{D}(\mathbf{A}^{-1})'.$$

In order to identify global shocks in the reduced form, one would usually need sign restrictions on all reduced-form loadings $\mathbf{A}^{-1}\chi$. That requires up to $n \times G$ restrictions instead of only n restrictions on the structural model.

C.2 Homogeneity restrictions on structural contemporaneous parameters

The total number of structural contemporaneous coefficients $\mathbf{A}$ increases quadratically with the number of countries in the panel. In order to deal with the resulting curse of dimensionality in the identification of $\mathbf{A}$, we assume that the foreign coefficients in the vector $\mathbf{a}_c^j$ are identical up to a scaling constant $w_{cc'}$, which is the average trade share of foreign country c' in the total trade of country c. This homogeneity restriction means that we only have to identify posterior distributions for the J domestic and J foreign coefficients of structural equation j in country c. These coefficients can be summarized in the $[2J \times 1]$ vector $\tilde{\mathbf{a}}_c^j$.

As an example, let us look at the structural contemporaneous coefficients in the first four equations, which correspond to Australia. Combining the four row vectors into one 4×8-matrix $\tilde{\mathbf{A}}_{AU}$, our baseline model identifies

$$\tilde{\mathbf{A}}_{AU} = \begin{pmatrix} 1 & -\alpha^{AU,\pi} & 0 & \alpha^{AU,q} & 0 & 0 & 0 & 0 \\ 1 & -\beta^{AU,\pi} & -\beta^{AU,r} & -\beta^{AU,q} & -\beta^{AU,y^*} & 0 & 0 & 0 \\ -(1-\rho^{AU})\psi^{AU,y} & -(1-\rho^{AU})\psi^{AU,\pi} & 1 & -(1-\rho^{AU})\psi^{AU,q} & 0 & (1-\rho^{AU})\psi^{AU,q} & 0 & 0 \\ -\theta^{AU,y} & 0 & 0 & 1 & \theta^{AU,y} & 0 & 0 & 0 \end{pmatrix}.$$

Returning to individual equations, the full coefficient vector $\mathbf{a}_c^j$, and thereby $\mathbf{A}$, can be derived from the restricted vector $\tilde{\mathbf{a}}_c^j$ as

$$\mathbf{a}_c^j = \tilde{\mathbf{a}}_c^j \mathbf{R}_c. \tag{C.1}$$

The $2J \times n$ restriction matrix $\mathbf{R}_c$ applies the appropriate scaling and allocates coefficients to the right position in $\mathbf{a}_c^j$. It is defined as

$$\underset{(2J\times n)}{\mathbf{R}_c} = \begin{pmatrix} \mathbf{0}_{J\times J} & \cdots & \mathbf{0}_{J\times J} & \mathbf{I}_J & \mathbf{0}_{J\times J} & \cdots & \mathbf{0}_{J\times J} \\ w_{c,1}\mathbf{I}_J & \cdots & w_{c,c-1}\mathbf{I}_J & \mathbf{0}_{J\times J} & w_{c,c+1}\mathbf{I}_J & \cdots & w_{c,C}\mathbf{I}_J \end{pmatrix}.$$

C.3 Prior distributions on lag coefficients and shock variances

Let $\iota_{r,c}$ be a $[k \times 1]$-vector that is one for the first lag of interest rates from country c, and zero otherwise. Let $\mathbf{v}_1 = (1/1^2, 1/2^2, \ldots, 1/p^2)'$ be a $(p \times 1)$ vector for lag scaling, and $\mathbf{v}_2(\mathbf{s}) = (1/s_{AU}^y, 1/s_{AU}^\pi, 1/s_{AU}^r, 1/s_{AU}^q, \ldots, 1/s_{US}^q)'$ a $(CJ \times 1)$ vector of inverse variances given by the $(CJ \times 1)$ vector $\mathbf{s}$. Let $\hat{\rho}_{cc'}^{ii'}$ be the correlation of residuals $\hat{\mathbf{e}}_{ct}^i$ and $\hat{\mathbf{e}}_{c't}^{i'}$ from fourth order univariate regressions of variable i (i') in country c (c'). Let finally $\mathbf{S}(\mathbf{s})$ be a prior variance-covariance matrix with entries $\mathbf{S}_{cc'}^{ii'}(\mathbf{s}) = \sqrt{s_c^i s_{c'}^{i'}}\hat{\rho}_{cc'}^{ii'}$.

Conditional on $\mathbf{A}$, we set the following prior distributions:

$$p\left(d_c^j|\mathbf{A}, \mathbf{s}\right) = \gamma\left((d_c^j)^{-1}; \kappa, \tau_c^j(\mathbf{A}, \mathbf{s})\right) \tag{C.2}$$

$$p\left(\mathbf{b}_c^j|\mathbf{A}, \lambda_0, \mathbf{s}, d_c^j\right) = \phi\left(\mathbf{b}; \mathbf{m}_c^j(\mathbf{A}), d_c^j\mathbf{M}_c^j(\lambda_0, \mathbf{s})\right) \tag{C.3}$$

$$p(s_c^i) = \gamma((s_c^i)^{-1}; \kappa_s, \tau_{s_c^i}) \tag{C.4}$$

$$p(\lambda_0) = \gamma(\lambda_0; \kappa_{\lambda_0}, \tau_{\lambda_0}) \tag{C.5}$$

with

$$\tau_c^j(\mathbf{A},\mathbf{s})) = \kappa \mathbf{a}_c^j \mathbf{S}(\mathbf{s})\, \mathbf{a}_c^{j\,'}$$

$$\mathbf{m}_c^j(\mathbf{A},\mathbf{s},\lambda_0) = \begin{cases} \mathbf{M}_c^j(\lambda_0,\mathbf{s}) \left[\left(\mathrm{diag}\left(\mathbf{v}_3(\lambda_0,\mathbf{s})\right)\right)^{-1} \eta\, \mathbf{a}_c^{j\,'} + \left(\frac{\rho_c}{V_\rho}\iota_{r,c}\right)' \right] & j = m \\ \eta\, \mathbf{a}_c^{j\,'} & \text{otherwise} \end{cases}$$

$$\mathbf{M}_c^j(\lambda_0,\mathbf{s}) = \begin{cases} \left(\left(\mathrm{diag}\left(\mathbf{v}_3(\lambda_0,\mathbf{s})\right)\right)^{-1} + \mathrm{diag}\left(\frac{1}{V_\rho}\iota_{r,c}\right) \right)^{-1} & \text{if } j = m \\ \mathrm{diag}\left(\mathbf{v}_3(\lambda_0,\mathbf{s})\right) & \text{otherwise} \end{cases}$$

$$\mathbf{v}_3(\lambda_0,\mathbf{s}) = \lambda_0^2 \begin{pmatrix} \mathbf{v}_1 \otimes \left(\mathrm{diag}(\omega \otimes \mathbf{1}_{(J\times 1)})\mathbf{v}_2(\mathbf{s})\right) \\ 100^2 \times \mathbf{1}_{(2\times 1)} \end{pmatrix}$$

As in Baumeister and Hamilton (2018), the prior mean of structural lag coefficients distribution combines (a) the prior belief that the data follow an AR(1) process with AR-coefficient $\phi = 0.75$, and (b) that the central bank engages in interest-rate smoothing, as described by the structural coefficient ρ^c. We give this prior a variance of $V_\rho = 0.1$.

We increase the tightness of priors on structural lag coefficients considerably in order to deal with the curse-of-dimensionality. We do this by multiplying the prior variance of every coefficient related to a variable from country c by ω_c^2, where ω_c is the average share of GDP of country c in our sample.

We choose $\kappa_s = 0.1$. The scale $\tau_{s_c^i}$ is set to 0.05 for output gaps, inflation and interest rates, and to 2 for real effective exchange rate growth. This ensures that $\mathbf{E}(1/s_c^i) = \frac{\tau_{s_c^i}}{\kappa_{s_c^i}}$ is roughly similar to the variance of residuals $\hat{\mathbf{e}}_{ct}^i$. For λ_0, we use a gamma-distribution with mode 0.2 and standard deviation 0.4 (i.e., $\kappa_{\lambda_0} = 1.64, \tau_{\lambda_0} = 0.3125$) (Sims and Zha, 1998; Giannone et al., 2015).

C.4 Posterior distributions and algorithm

C.4.1 Posterior distributions

Out posterior sample is an extension of the Metropolis-within-Gibbs of Baumeister and Hamilton (2015) that accounts for hyperparameters for the Minnesota prior, as well as global shocks and their loadings. Two things are worthwhile to note before we develop the full sampler. First, we can sample from the known conditional posterior distributions of global shocks and loadings in the Gibbs sampler as in Korobilis (2022). Second, the sampling of hyperparameters in Giannone et al. (2015) requires a Metropolis step that uses the same likelihood kernel as the sampling of structural contemporaneous coefficients $\mathbf{A}$. That is, we can sample $(\mathbf{A}, \lambda_0, \mathbf{s})$ together in the same step, as proven in subsection C.4.2.

Let the $(T \times n)$-dimensional matrix $\mathbf{Y} = (\mathbf{y}_1, \ldots, \mathbf{y}_T)'$ and $(T \times k)$-dimensional matrix $\mathbf{X} = (\mathbf{x}_0, \ldots, \mathbf{x}_{T-1})'$ collect all observations. For country c and structural equation j, we construct extended data $\tilde{\mathbf{Y}}_c^j$ and $\tilde{\mathbf{X}}_c^j$ by applying two data modifications. First, we condition on $\mathbf{A}$ and hyperparameters $\lambda_0, \mathbf{s}$, and append the corresponding normal prior $\mathcal{N}(\mathbf{m}_c^j, d_c^j \mathbf{M}_c^j)$ as dummy observations (for ease of readability, we drop the dependence of $\mathbf{m}_c^j$ and $\mathbf{M}_c^j$ on $(\mathbf{A}, \lambda_0, \mathbf{s})$). Let $\mathbf{P}_c^j$ be the Cholesky factor of $(\mathbf{M}_c^j)^{-1}$, i.e. $(\mathbf{M}_c^j)^{-1} = \mathbf{P}_c^j \mathbf{P}_c^{j'}$. Second, we shift global shocks to the left-hand side of the equation such that – conditional on all other parameters of the model – structural lags and variances follow normal-inverse gamma distributions as in Baumeister and Hamilton (2018). The properly augmented data for country c and structural equation j are defined as:

$$\underset{(T+k+1)\times 1}{\tilde{\mathbf{Y}}_c^j} = \begin{pmatrix} \left(\mathbf{a}_c^j \mathbf{Y}' - \chi_c^j \mathbf{U}_{gT}^{j'}\right)' \\ \mathbf{m}_c^j \mathbf{P}_c^{j'} \end{pmatrix}, \qquad \underset{(T+k+1)\times(k+1)}{\tilde{\mathbf{X}}_c^j} = \begin{pmatrix} \mathbf{X} \\ \mathbf{P}_c^{j'} \end{pmatrix}$$

Conditional on global shocks $\mathbf{U}_{gT}$, these augmented data can be used to derive the posterior distributions of $\mathbf{A}, \tilde{\mathbf{B}}, \chi, \mathbf{D}$ just as in Baumeister and Hamilton (2015) and Giannone et al. (2015), see also the following subsection C.4.2:

$$\begin{aligned} p\left(\mathbf{A}, \lambda_0, \mathbf{s} | \mathbf{Y}_T, \chi, \mathbf{U}_{gT}\right) \propto & p(\mathbf{A}) p(\lambda_0) p(\mathbf{s}) [\det(\mathbf{A}\hat{\Omega}\mathbf{A'})]^{T/2} \\ & \prod_{c=1}^{C} \prod_{j=1}^{J} \frac{\left|\mathbf{M}_c^{j^*}\right|^{1/2}}{\left|\mathbf{M}_c^j\right|^{1/2}} \frac{(\tau_c^j)^{\kappa_c^j}}{\left(2\tau_c^{j^*}/T\right)^{\kappa_c^{j^*}}} \frac{\Gamma\left((\kappa_c^j)^*\right)}{\Gamma\left(\kappa_c^j\right)} \end{aligned} \tag{C.6}$$

$$p\left(\mathbf{D} | \mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{Y}_T, \chi, \mathbf{U}_{gT}\right) = \prod_{c=1}^{C} \prod_{j=1}^{J} \gamma\left(\left(d_c^j\right)^{-1}; \kappa_c^{j^*}, \tau_c^{j^*}\right) \tag{C.7}$$

$$p\left(\mathbf{B} | \mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{D}, \mathbf{Y}_T, \chi, \mathbf{U}_{gT}\right) = \prod_{c=1}^{C} \prod_{j=1}^{J} \phi\left(\mathbf{b}_c^j; \mathbf{m}_c^{j^*}, \mathbf{M}_c^{j^*}\right) \tag{C.8}$$

with

$$\begin{aligned} \kappa_c^{j^*} &= \kappa_c^j + T/2 \\ \tau_c^{j^*} &= \tau_c^j + \zeta_c^{j^*}/2 \\ \zeta_c^{j^*} &= \tilde{\mathbf{Y}}_c^{j'} \tilde{\mathbf{Y}}_c^j - \tilde{\mathbf{Y}}_c^{j'} \tilde{\mathbf{X}}_c^j \left(\tilde{\mathbf{X}}_c^{j'} \tilde{\mathbf{X}}_c^j\right)^{-1} \tilde{\mathbf{X}}_c^{j'} \tilde{\mathbf{Y}}_c^j \\ \mathbf{m}_c^{j^*} &= \left[\left(\tilde{\mathbf{X}}_c^{j'} \tilde{\mathbf{X}}_c^j\right)^{-1} \tilde{\mathbf{X}}_c^{j'} \tilde{\mathbf{Y}}_c^j\right]' \\ \mathbf{M}_c^{j^*} &= \left(\tilde{\mathbf{X}}_c^{j'} \tilde{\mathbf{X}}_c^j\right)^{-1}. \end{aligned}$$

Using Korobilis (2022), we can derive the conditional posterior distributions of global shocks and their loadings. This is particularly easy in our baseline case, where each structural equation is affected by at most one global shock.[2]

[2] A possible extension where multiple shocks load onto one equation, such as in a robustness check where global supply and demand shocks load onto exchange rates of commodity-exporting countries, are also easy to derive.

The conditional posterior distributions of loading χ_c^j depends on the correlation between global shocks $\mathbf{u}_{gt}^j$ and combined structural shocks $\check{\mathbf{u}}_{ct}^j = \left[\mathbf{a}_c^j \mathbf{Y}_T' - \mathbf{b}_c^j \mathbf{X}_T'\right]'$:

$$p(\chi_c^j|\mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{B}, \mathbf{D}, \mathbf{Y}_T, \mathbf{U}_{gT}) = \mathcal{TN}_{\chi_c^j>0}\left(V_{\chi,c}^{j^*}\left(\mathbf{u}_{gt}^j\right)'\check{\mathbf{u}}_{ct}^j, d_c^j V_{\chi,c}^{j^*}\right) \tag{C.9}$$
$$V_{\chi,c}^{j^*} = \left(\left(\mathbf{u}_{gt}^j\right)'\mathbf{u}_{gt}^j + V_\chi^{-1}\right)^{-1}$$

With $\check{\mathbf{U}}_T$ the $(T \times n)$ matrix of correlated structural shocks from all equations, the posterior distribution of global shocks $\mathbf{U}_{gT}$ is given by:

$$p(U_{gT}|\mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{B}, \mathbf{D}, \mathbf{Y}_T, \chi) = \mathcal{N}\left(\left(\mathbf{M}_g^* \chi' \mathbf{D}^{-1}\check{\mathbf{U}}_T'\right)', \mathbf{M}_g^*\right) \tag{C.10}$$
$$\mathbf{M}_g^* = \left(I_G + \chi'\mathbf{D}^{-1}\chi\right)^{-1}$$

C.4.2 Derivation of equation (C.6)

Considering that the prior distributions $p(\mathbf{A}, \lambda_0, \mathbf{s})$ are independent, and that the prior distributions of $p(\mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{B}, \mathbf{D})$ do not depend on global shocks, we have after rearranging:

$$\begin{aligned}
&p(\mathbf{Y}_T, \mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{B}, \mathbf{D}|\chi, \mathbf{U}_{gT}) \\
&\quad= p(\mathbf{A}, \lambda_0, \mathbf{s})p(\mathbf{D}|\mathbf{A}, \lambda_0, \mathbf{s})p(\mathbf{B}|\mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{D})p(\mathbf{Y}_T|\mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{B}, \mathbf{D}, \chi, \mathbf{U}_{gT}) \\
&\quad= p(\mathbf{A})p(\lambda_0)p(\mathbf{s})(2\pi)^{-Tn/2}|\det(\mathbf{A})|^T \\
&\quad\quad\times \prod_{c=1}^{C}\prod_{j=1}^{J}\left\{(d_c^j)^{-(\kappa_c^j-1)}\frac{(\tau_c^j)^{\kappa_c^j}}{\Gamma(\kappa_c^j)}\frac{\Gamma(\kappa_c^{j^*})}{(\tau_c^{j^*})^{\kappa_c^{j^*}}}\frac{(\tau_c^{j^*})^{\kappa_c^{j^*}}}{\Gamma(\kappa_c^{j^*})}(d_c^j)^{-T/2}\exp\left[-\frac{\tau_c^{j^*}}{d_c^j}\right]\right. \\
&\quad\quad\times \left.\frac{|\mathbf{M}_c^{j^*}|^{1/2}}{|\mathbf{M}_c^j|^{1/2}}\frac{1}{(2\pi)^{k/2}|d_c^j\mathbf{M}_c^{j^*}|^{1/2}}\exp\left[-\frac{\left(\mathbf{b}_c^j - \mathbf{m}_c^{j^*}\right)'\left(\mathbf{M}_c^{j^*}\right)^{-1}\left(\mathbf{b}_c^j - \mathbf{m}_c^{j^*}\right)}{2d_c^j}\right]\right\} \\
&\quad= p(\mathbf{A})p(\lambda_0)p(\mathbf{s})(2\pi)^{-Tn/2}|\det(\mathbf{A})|^T \\
&\quad\quad\times \prod_{c=1}^{C}\prod_{j=1}^{J}\left\{\frac{|\mathbf{M}_c^{j^*}|^{1/2}}{|\mathbf{M}_c^j|^{1/2}}\frac{(\tau_c^j)^{\kappa_c^j}}{\Gamma(\kappa_c^j)}\frac{\Gamma(\kappa_c^{j^*})}{(\tau_c^{j^*})^{\kappa_c^{j^*}}}\right\}\gamma\left(\left(d_c^j\right)^{-1}; \kappa_c^{j^*}, \tau_c^{j^*}\right)\phi\left(\mathbf{b}_c^j; \mathbf{m}_c^{j^*}, \mathbf{M}_c^{j^*}\right)
\end{aligned} \tag{C.11}$$

Equation (C.11) means that, conditional on global shocks and their loading, the posterior distribution of $(\mathbf{A}, \lambda_0, \mathbf{s})$ is proportional to

$$p(\mathbf{A}, \lambda_0, \mathbf{s}|\mathbf{Y}_T, \chi, \mathbf{U}_{gT}) \propto p(\mathbf{A})p(\lambda_0)p(\mathbf{s})(2\pi)^{-Tn/2}|\det(\mathbf{A})|^T\prod_{c=1}^{C}\prod_{j=1}^{J}\left\{\frac{|\mathbf{M}_c^{j^*}|^{1/2}}{|\mathbf{M}_c^j|^{1/2}}\frac{(\tau_c^j)^{\kappa_c^j}}{\Gamma(\kappa_c^j)}\frac{\Gamma(\kappa_c^{j^*})}{(\tau_c^{j^*})^{\kappa_c^{j^*}}}\right\}.$$

Equation (C.6) can be obtained after removing some constants and multiplying with $|\hat{\Omega}_T|^{T/2}$, where $\hat{\Omega}_T$ is the variance-covariance matrix of reduced-form errors, which does not depend on any of the model parameters:

$$p(\mathbf{A}, \lambda_0, \mathbf{s}|\mathbf{Y}_T, \chi, \mathbf{U}_{gT}) \propto p(\mathbf{A})p(\lambda_0)p(\mathbf{s})|\det(\mathbf{A}\hat{\Omega}_T\mathbf{A})|^T \prod_{c=1}^{C}\prod_{j=1}^{J}\left\{\frac{|\mathbf{M}_c^{j*}|^{1/2}}{|\mathbf{M}_c^{j}|^{1/2}}\frac{(\tau_c^j)^{\kappa_c^j}}{\Gamma(\kappa_c^j)}\frac{\Gamma(\kappa_c^{j*})}{(2\tau_c^{j*}/T)^{\kappa_c^{j*}}}\right\}.$$

C.4.3 Drawing from the posterior distribution

We use a Metropolis-within-Gibbs to generate draws from the posterior distributions in the following sequence:

1. Draw $(\mathbf{A}, \lambda_0, \mathbf{s})$ from $p(\mathbf{A}, \lambda_0, \mathbf{s}|\mathbf{Y}_T, \chi, \mathbf{U}_{gT})$, equation (C.6), in a Metropolis-Hastings step evaluated for $C+1$ parameter blocks, namely C random blocks for coefficients in $\mathbf{A}$, followed by a single block for hyperparameters $(\lambda_0, \mathbf{s})$.

2. For every country c and structural equation j, ...

   (a) ... draw $(d_c^j)^{-1}$ from $p\left(\mathbf{D}|\mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{Y}_T, \chi, \mathbf{U}_{gT}\right)$, equation (C.7).

   (b) ... draw $\mathbf{b}_c^j$ jointly from $p\left(\mathbf{B}|\mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{D}, \mathbf{Y}_T, \chi, \mathbf{U}_{gT}\right)$, equation (C.8).

   (c) ... calculate combined structural shocks $\check{\mathbf{u}}_t = \mathbf{A}\mathbf{y}_t - \mathbf{B}\mathbf{x}_{t-1}$.

3. Draw χ from $p(\chi|\mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{D}, \mathbf{B}, \mathbf{Y}_T, \mathbf{U}_{gT})$, equation (C.9).

4. Draw $\mathbf{U}_{gT}$ from $p(\mathbf{U}_{gt}|\mathbf{A}, \lambda_0, \mathbf{s}, \mathbf{D}, \mathbf{B}, \mathbf{Y}_T, \chi)$, equations (C.10).

Following Baumeister and Hamilton (2015), we calculate the mode of the posterior likelihood $p(\mathbf{A}, \lambda_0, \mathbf{s}|\mathbf{Y}_T, \mathbf{0}_{n\times G}, \mathbf{U}_{gT})$, albeit at loadings of zero. We take the parameters $(\mathbf{A}, \lambda_0, \mathbf{s})$ at the mode as initial draws, and their Hessian as the most promising search direction for the Metropolis-Hastings steps.

Because of global shocks, we want to fine-tune this initial draw further. Thus, we run a pre-sampling such that we obtain twice as many draws as the number of free parameters of $\mathbf{A}$ after 400'000 burn-in and thinning of 500 draws (total chain length of baseline model: 484'000 draws). Taking the median retained draw as starting point, and the variance-covariance of the retained draws as search direction, the main algorithm keeps a total of 20'000 retained draws after a burn-in of 200,000 draws and thinning of 50 draws (total chain length: 1'200'000). For the prior distribution and robustness checks, we reduce the number of retained draws to 5'000. This "thins" out posterior distributions of structural coefficients, but does not materially affect the quantile-based results for forecast-error-variance decompositions and impulse-response-functions. We adapt the step-size during the burn-in phase to achieve an acceptance probability of 30% for the Metropolis-Hastings step.

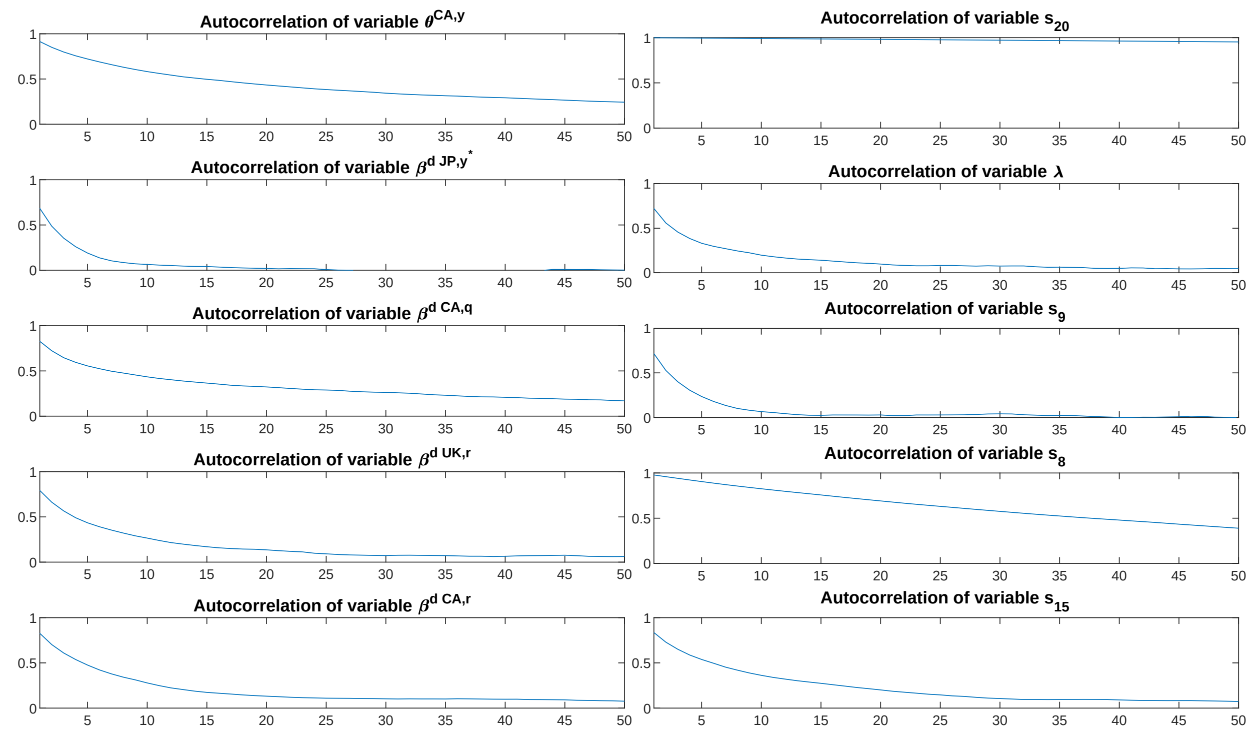
C.5 Convergence statistics

Figure C.1 and C.2 plot the autocorrelation across draws (after burn-in) and all draws for the four chains exemplary for the coefficients which have the weakest convergence statistics according to Geweke (1992) test for equality in means. These coefficients have by far the lowest p-values, mostly between 0.01 and 0.05. Across the 106 coefficients drawn in the Metropolis-Hastings step, only 10 have a p-value below 0.05, which is only marginally larger than the share one would expect from multiple testing.

We see that the autocorrelation and trace plots for structural contemporaneous coefficients $\mathbf{A}$ is well behaved. Some of the prior variances $\mathbf{s}$, however, are be correlated across retained draws. We set the same hyperparameters for countries in all prior distributions. However, since half of the parameters ($\alpha^{sJP,q}, \beta^{dJP,\pi}, \mathbf{s}_{13}, \mathbf{s}_{14}, \mathbf{s}_{16}$) are related to Japanese variables, this might indicate that setting country-specific hyperparameters in the prior distributions of $\mathbf{A}$ and $\mathbf{s}$ could improve convergence.

Figure C.1: Autocorrelations of draws

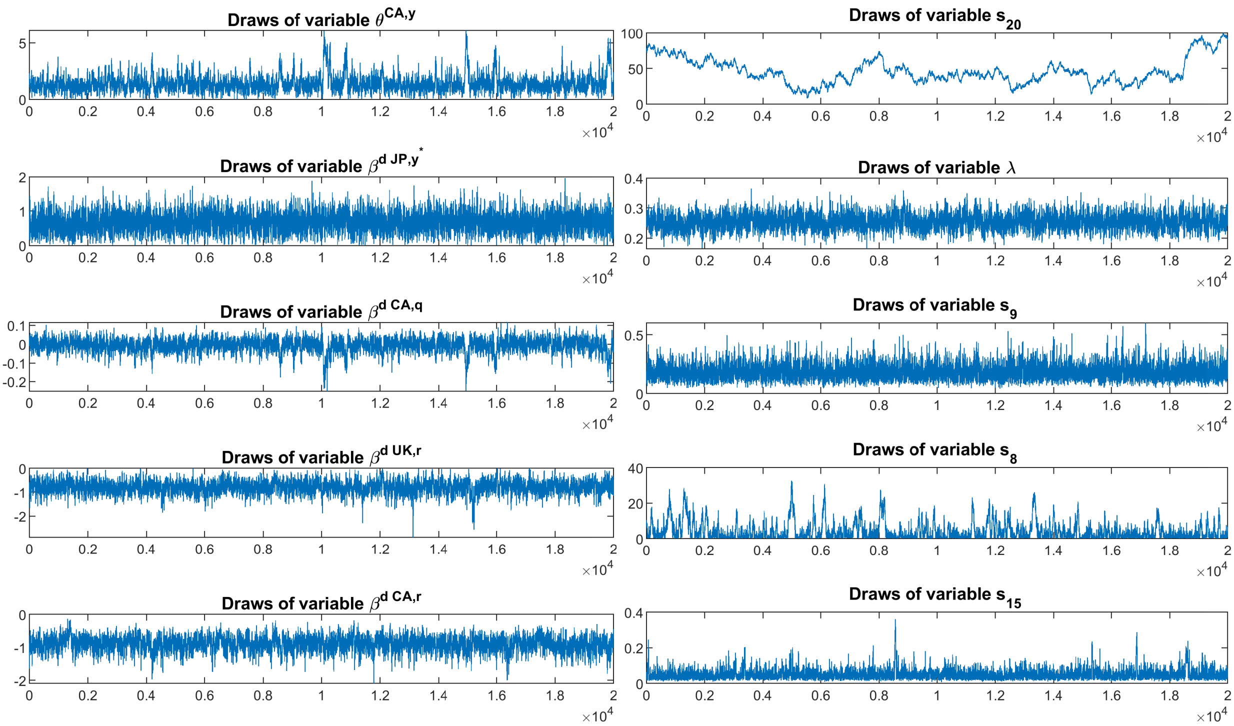
(a) A **(b) $\mathbf{s}, \lambda$**



NOTES: The plots show the autocorrelation across draws (after burn-in) of the structural parameters in $\mathbf{A}$ (left subplot) and hyperparameters ($\mathbf{s}, \lambda$) (right subplot) with the weakest convergence statistics.

Figure C.2: Trace plot of draws

(a) A **(b)** $\mathbf{s}, \lambda$



NOTES: Trace plots of the structural parameters in $\mathbf{A}$ (left subplot) and hyperparameters $(\mathbf{s}, \lambda)$ (right subplot) with the weakest convergence statistics.

D Further results

D.1 Forecast error variance and correlation decompositions in the baseline model

The following figures show the forecast error variance decomposition of country-specific variables (in subplots) to domestic (dom) and foreign (for) supply, demand, monetary policy and exchange rate shocks and global (glob) supply, demand and monetary policy shocks over 20 quarters.

Figure D.1: Forecast error variance decomposition of output gaps

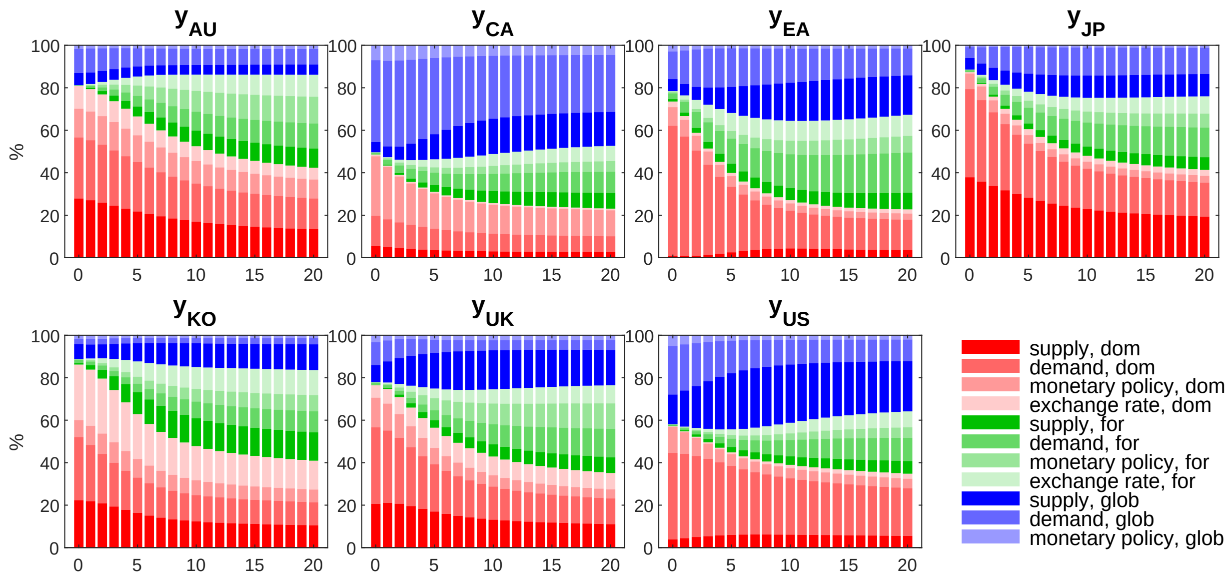


Figure D.2: Forecast error variance decomposition of inflation

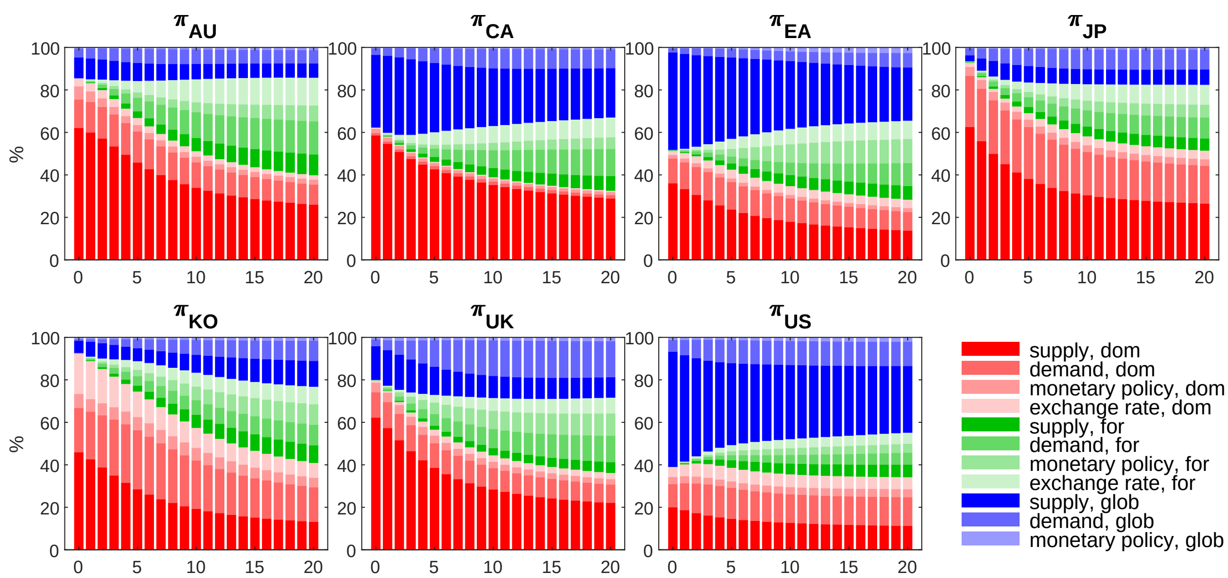


Figure D.3: Forecast error variance decomposition of exchange rate growth

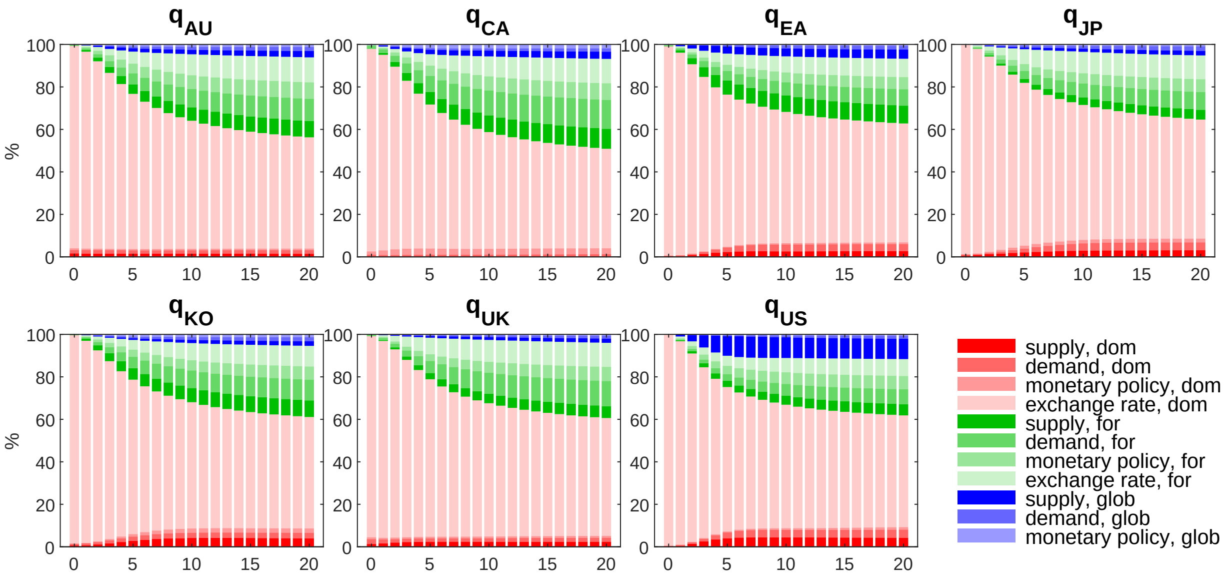


Figure D.4: Forecast error correlation decomposition between US and country-specific interest rates

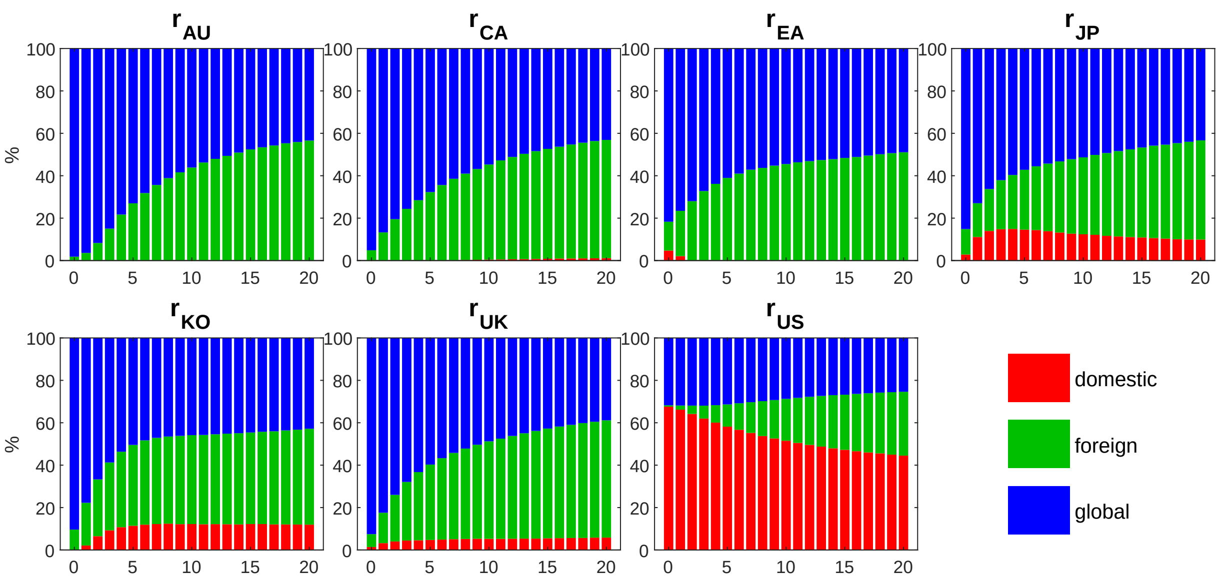


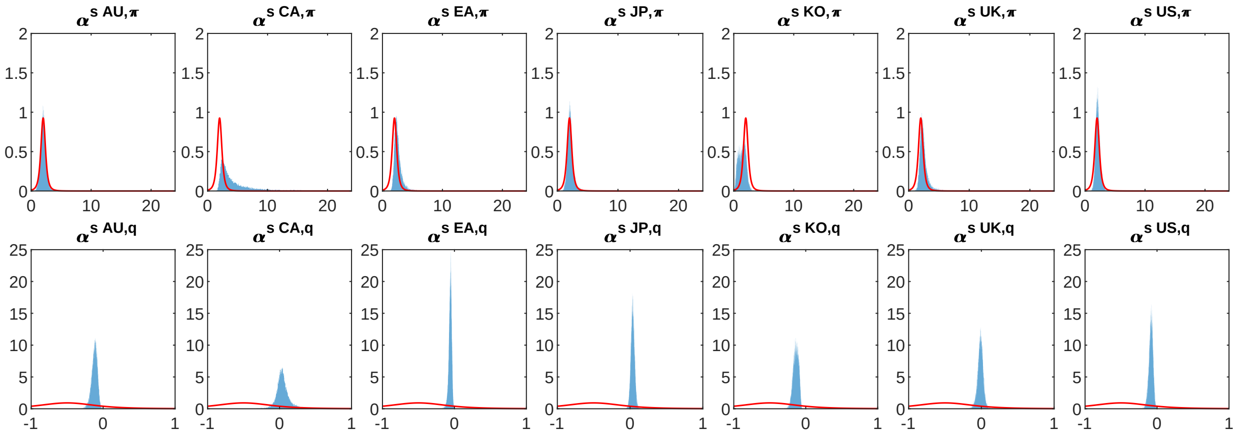
Table D.1: Decomposition of correlations of interest rates, $H = 20$

	AU	CA	EA	JP	KO	UK	US
				domestic			
group agg	-0.02	0.01	-0.01	0.10	0.12	**0.06**	**0.44**
	[-0.29,0.10]	[-0.03,0.09]	[-0.30,0.19]	[-0.03,0.32]	[-0.09,0.39]	[0.00,0.19]	[0.28,0.61]
supply, dom	0.00	0.00	-0.02	0.01	0.02	0.02	**0.02**
	[-0.03,0.03]	[-0.00,0.01]	[-0.16,0.04]	[-0.03,0.11]	[-0.03,0.12]	[-0.00,0.10]	[0.00,0.08]
demand, dom	-0.01	0.01	-0.01	0.05	0.08	0.02	**0.29**
	[-0.16,0.07]	[-0.01,0.06]	[-0.12,0.09]	[-0.01,0.19]	[-0.05,0.27]	[-0.01,0.11]	[0.10,0.46]
mon. pol., dom	-0.01	-0.00	0.01	0.01	0.00	0.00	**0.12**
	[-0.12,0.04]	[-0.03,0.04]	[-0.08,0.14]	[-0.05,0.16]	[-0.04,0.06]	[-0.01,0.04]	[0.02,0.32]
er, dom	0.00	0.00	0.01	0.01	0.01	0.01	**0.02**
	[-0.02,0.03]	[-0.00,0.04]	[-0.11,0.13]	[-0.04,0.10]	[-0.04,0.10]	[-0.01,0.07]	[0.00,0.11]
				foreign			
group agg	**0.57**	**0.56**	**0.51**	**0.47**	0.45	**0.55**	**0.30**
	[0.05,0.85]	[0.28,0.79]	[0.15,0.88]	[0.14,0.75]	[-0.06,0.82]	[0.33,0.77]	[0.15,0.50]
supply, for	0.11	**0.09**	0.09	0.07	0.06	**0.08**	**0.08**
	[-0.03,0.32]	[0.01,0.21]	[-0.02,0.24]	[-0.02,0.21]	[-0.14,0.24]	[0.01,0.20]	[0.02,0.17]
demand, for	**0.29**	**0.33**	**0.31**	**0.27**	**0.28**	**0.32**	**0.11**
	[0.00,0.63]	[0.12,0.54]	[0.08,0.65]	[0.04,0.51]	[0.04,0.62]	[0.11,0.51]	[0.03,0.22]
mon. pol., for	0.12	0.08	0.07	0.09	0.06	**0.10**	**0.05**
	[-0.08,0.36]	[-0.01,0.26]	[-0.10,0.31]	[-0.06,0.31]	[-0.21,0.27]	[0.01,0.27]	[0.01,0.13]
er, for	0.05	0.06	0.04	0.04	0.06	0.06	**0.06**
	[-0.32,0.28]	[-0.03,0.21]	[-0.23,0.21]	[-0.16,0.23]	[-0.17,0.30]	[-0.01,0.22]	[0.02,0.17]
				global			
group agg	**0.45**	**0.43**	**0.49**	**0.43**	**0.43**	**0.39**	**0.25**
	[0.20,1.12]	[0.20,0.69]	[0.15,1.00]	[0.15,0.80]	[0.12,0.88]	[0.16,0.61]	[0.10,0.45]
supply, glob	0.04	0.03	0.11	0.04	0.05	0.08	**0.04**
	[-0.11,0.22]	[-0.01,0.18]	[-0.16,0.36]	[-0.04,0.21]	[-0.07,0.28]	[-0.02,0.24]	[0.00,0.17]
demand, glob	**0.34**	**0.35**	**0.32**	**0.33**	**0.29**	**0.26**	**0.16**
	[0.05,0.99]	[0.08,0.60]	[0.03,0.77]	[0.04,0.64]	[0.02,0.68]	[0.04,0.47]	[0.02,0.35]
mon. pol., glob	0.08	**0.05**	0.07	0.07	0.08	**0.06**	**0.04**
	[-0.01,0.38]	[0.00,0.23]	[-0.00,0.38]	[-0.00,0.30]	[-0.01,0.37]	[0.00,0.24]	[0.00,0.20]

Note: Contributions relate to correlation between the interest rate of country c and the global interest rate at horizon $H = 20$. We first report aggregates for domestic, foreign and global shocks before breaking them down into different types of shocks. We report the mode of the decomposition (scaled such that they sum up to 1) together with 95% credibility sets. Correlations where the 95% credibility sets do not contain zero are marked in bold.

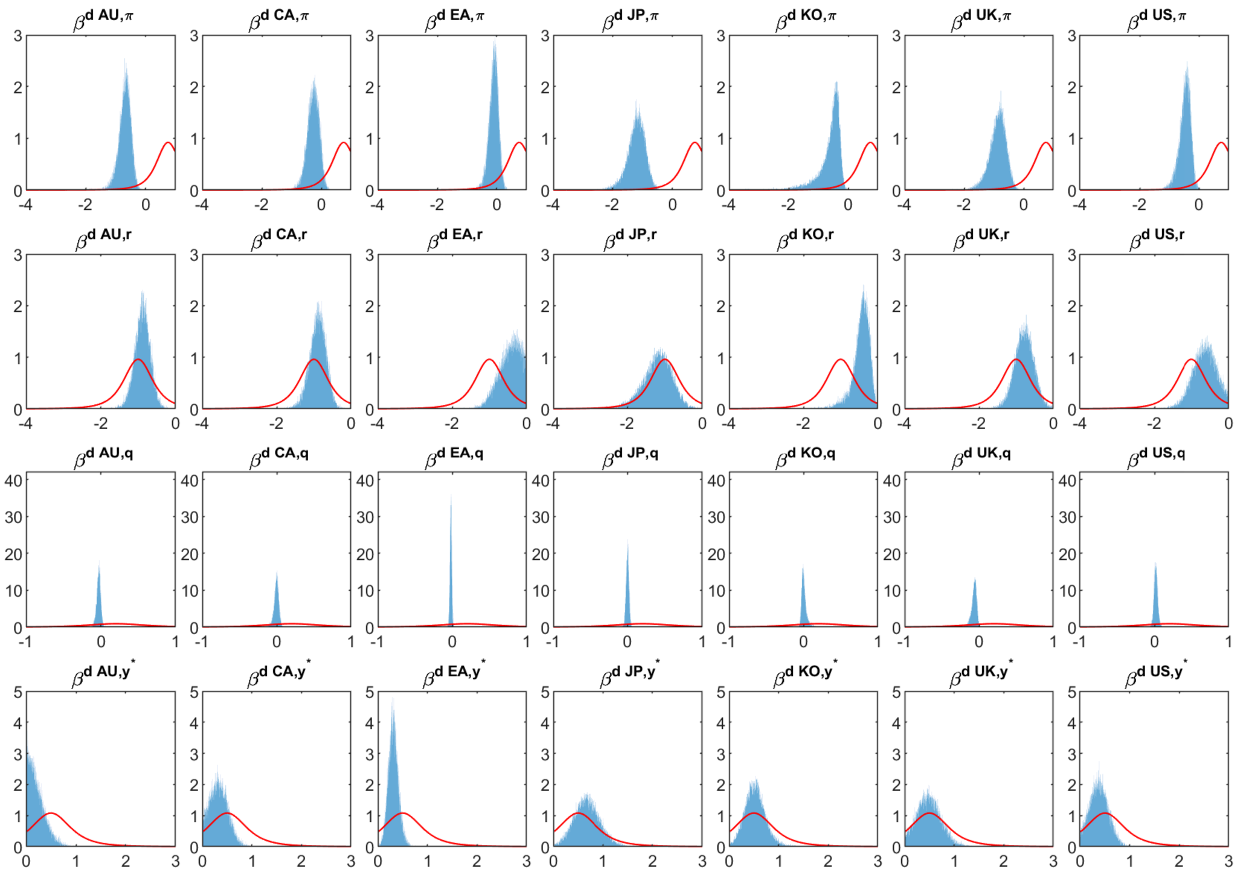
D.2 Structural contemporaneous coefficients

Figure D.5: Posterior distributions of structural contemporaneous coefficients in the supply equation



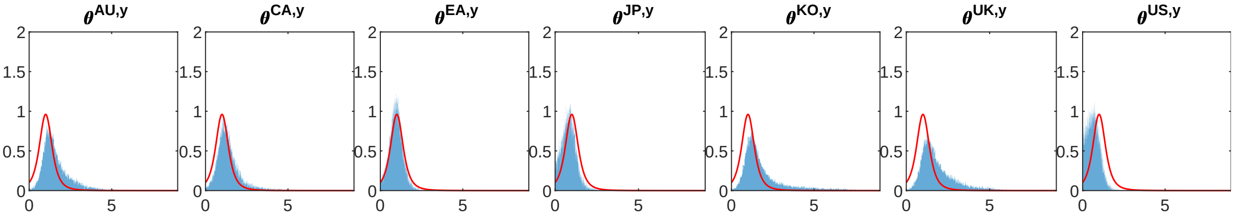
NOTES: The histograms show the posterior distribution of structural contemporaneous coefficients in the supply equation together with the prior distribution (red line).

Figure D.6: Posterior distributions of structural contemporaneous coefficients in the demand equation



NOTES: The histograms show the posterior distribution of structural contemporaneous coefficients in the demand equation together with the prior distribution (red line).

Figure D.7: Posterior distributions of structural contemporaneous coefficients in the exchange rate equation



NOTES: The histograms show the posterior distribution of structural contemporaneous coefficients in the exchange rate equation together with the prior distribution (red line).

Table D.2: Prior and posterior mean and credibility set of structural contemporaneous coefficients

	prior		posterior			prior		posterior	
	mean	95% cs	mean	95% cs		mean	95% cs	mean	95% cs
$\alpha^{sAU,\pi}$	2.02	[0.85,3.29]	2.04	[1.16,3.14]	$\alpha^{sCA,\pi}$	2.02	[0.85,3.29]	5.77	[1.95,25.31]
$\alpha^{sAU,q}$	-0.50	[-1.79,0.79]	-0.12	[-0.22,-0.05]	$\alpha^{sCA,q}$	-0.50	[-1.79,0.79]	0.05	[-0.13,0.33]
$\beta^{dAU,\pi}$	0.75	[-0.53,2.03]	-0.70	[-1.16,-0.35]	$\beta^{dCA,\pi}$	0.75	[-0.53,2.03]	-0.27	[-0.71,0.08]
$\beta^{dAU,r}$	-1.07	[-2.31,-0.17]	-0.89	[-1.32,-0.49]	$\beta^{dCA,r}$	-1.07	[-2.31,-0.17]	-0.90	[-1.38,-0.49]
$\beta^{dAU,q}$	0.20	[-1.09,1.49]	-0.03	[-0.09,0.02]	$\beta^{dCA,q}$	0.20	[-1.09,1.49]	-0.01	[-0.09,0.05]
β^{dAU,y^*}	0.67	[0.05,1.99]	0.22	[0.01,0.62]	β^{dCA,y^*}	0.67	[0.05,1.99]	0.33	[0.03,0.71]
$\psi^{AU,y}$	0.67	[0.05,1.99]	1.46	[0.81,2.42]	$\psi^{CA,y}$	0.67	[0.05,1.99]	3.43	[1.53,7.62]
$\psi^{AU,\pi}$	1.54	[0.47,2.79]	1.49	[0.93,2.24]	$\psi^{CA,\pi}$	1.54	[0.47,2.79]	1.18	[0.31,2.05]
$\psi^{AU,q}$	-0.00	[-1.29,1.29]	0.06	[-0.01,0.15]	$\psi^{CA,q}$	-0.00	[-1.29,1.29]	0.12	[-0.03,0.34]
$\theta^{AU,y}$	1.07	[0.19,2.33]	1.64	[0.51,3.97]	$\theta^{CA,y}$	1.07	[0.19,2.33]	1.41	[0.35,3.48]
ρ^{AU}	0.50	[0.13,0.87]	0.49	[0.24,0.69]	ρ^{CA}	0.50	[0.13,0.87]	0.55	[0.23,0.79]
$\alpha^{sEA,\pi}$	2.02	[0.85,3.29]	2.60	[1.73,4.47]	$\alpha^{sJP,\pi}$	2.02	[0.85,3.29]	2.17	[1.33,3.38]
$\alpha^{sEA,q}$	-0.50	[-1.79,0.79]	-0.06	[-0.11,-0.02]	$\alpha^{sJP,q}$	-0.50	[-1.79,0.79]	0.04	[-0.01,0.09]
$\beta^{dEA,\pi}$	0.75	[-0.53,2.03]	-0.11	[-0.46,0.17]	$\beta^{dJP,\pi}$	0.75	[-0.53,2.03]	-1.21	[-1.92,-0.69]
$\beta^{dEA,r}$	-1.07	[-2.31,-0.17]	-0.46	[-1.09,-0.03]	$\beta^{dJP,r}$	-1.07	[-2.31,-0.17]	-1.21	[-2.13,-0.51]
$\beta^{dEA,q}$	0.20	[-1.09,1.49]	-0.02	[-0.04,0.01]	$\beta^{dJP,q}$	0.20	[-1.09,1.49]	-0.00	[-0.04,0.04]
β^{dEA,y^*}	0.67	[0.05,1.99]	0.31	[0.12,0.53]	β^{dJP,y^*}	0.67	[0.05,1.99]	0.68	[0.18,1.23]
$\psi^{EA,y}$	0.67	[0.05,1.99]	3.10	[1.02,7.06]	$\psi^{JP,y}$	0.67	[0.05,1.99]	0.91	[0.44,1.57]
$\psi^{EA,\pi}$	1.54	[0.47,2.79]	1.23	[0.48,2.00]	$\psi^{JP,\pi}$	1.54	[0.47,2.79]	1.43	[0.77,2.20]
$\psi^{EA,q}$	-0.00	[-1.29,1.29]	0.02	[-0.07,0.14]	$\psi^{JP,q}$	-0.00	[-1.29,1.29]	0.04	[-0.02,0.11]
$\theta^{EA,y}$	1.07	[0.19,2.33]	0.95	[0.17,1.82]	$\theta^{JP,y}$	1.07	[0.19,2.33]	0.82	[0.08,1.75]
ρ^{EA}	0.50	[0.13,0.87]	0.75	[0.59,0.87]	ρ^{JP}	0.50	[0.13,0.87]	0.78	[0.63,0.89]
$\alpha^{sKO,\pi}$	2.02	[0.85,3.29]	1.41	[0.36,2.55]	$\alpha^{sUK,\pi}$	2.02	[0.85,3.29]	2.50	[1.58,4.60]
$\alpha^{sKO,q}$	-0.50	[-1.79,0.79]	-0.14	[-0.22,-0.07]	$\alpha^{sUK,q}$	-0.50	[-1.79,0.79]	-0.01	[-0.10,0.06]
$\beta^{dKO,\pi}$	0.75	[-0.53,2.03]	-0.64	[-1.84,-0.20]	$\beta^{dUK,\pi}$	0.75	[-0.53,2.03]	-0.88	[-1.56,-0.39]
$\beta^{dKO,r}$	-1.07	[-2.31,-0.17]	-0.45	[-1.09,-0.12]	$\beta^{dUK,r}$	-1.07	[-2.31,-0.17]	-0.79	[-1.33,-0.28]
$\beta^{dKO,q}$	0.20	[-1.09,1.49]	0.01	[-0.04,0.12]	$\beta^{dUK,q}$	0.20	[-1.09,1.49]	-0.07	[-0.16,-0.01]
β^{dKO,y^*}	0.67	[0.05,1.99]	0.53	[0.15,1.02]	β^{dUK,y^*}	0.67	[0.05,1.99]	0.50	[0.06,1.01]
$\psi^{KO,y}$	0.67	[0.05,1.99]	1.16	[0.35,2.47]	$\psi^{UK,y}$	0.67	[0.05,1.99]	2.01	[1.08,3.53]
$\psi^{KO,\pi}$	1.54	[0.47,2.79]	1.67	[1.01,2.60]	$\psi^{UK,\pi}$	1.54	[0.47,2.79]	1.47	[0.83,2.25]

Table D.2: Prior and posterior mean and credibility set of structural contemporaneous coefficients

	prior		posterior			prior		posterior	
	mean	95% cs	mean	95% cs		mean	95% cs	mean	95% cs
$\psi^{KO,q}$	-0.00	[-1.29,1.29]	0.12	[0.03,0.25]	$\psi^{UK,q}$	-0.00	[-1.29,1.29]	0.07	[-0.02,0.18]
$\theta^{KO,y}$	1.07	[0.19,2.33]	1.94	[0.51,6.34]	$\theta^{UK,y}$	1.07	[0.19,2.33]	1.87	[0.56,4.79]
ρ^{KO}	0.50	[0.13,0.87]	0.44	[0.16,0.71]	ρ^{UK}	0.50	[0.13,0.87]	0.52	[0.27,0.74]
$\alpha^{sUS,\pi}$	2.02	[0.85,3.29]	2.13	[1.46,3.02]					
$\alpha^{sUS,q}$	-0.50	[-1.79,0.79]	-0.08	[-0.15,-0.03]					
$\beta^{dUS,\pi}$	0.75	[-0.53,2.03]	-0.47	[-0.94,-0.15]					
$\beta^{dUS,r}$	-1.07	[-2.31,-0.17]	-0.65	[-1.33,-0.08]					
$\beta^{dUS,q}$	0.20	[-1.09,1.49]	0.02	[-0.03,0.07]					
β^{dUS,y^*}	0.67	[0.05,1.99]	0.39	[0.05,0.77]					
$\psi^{US,y}$	0.67	[0.05,1.99]	2.57	[0.99,5.25]					
$\psi^{US,\pi}$	1.54	[0.47,2.79]	1.18	[0.50,1.94]					
$\psi^{US,q}$	-0.00	[-1.29,1.29]	-0.02	[-0.14,0.09]					
$\theta^{US,y}$	1.07	[0.19,2.33]	0.65	[0.04,1.45]					
ρ^{US}	0.50	[0.13,0.87]	0.64	[0.38,0.82]					

D.3 Impulse-response functions in the baseline model

The solid lines in all the following figures show median impulse responses of country-specific variables (in columns, variable name in title) to country-specific demand, supply, monetary policy and exchange rate shocks (in rows) over 20 quarters. The shaded areas (dotted lines) show the 68% (95%) posterior credibility sets. Red dashed lines show median prior impulse response functions. The shocks have size of one unit. Figures D.8 to D.11 show the impulse responses to domestic shocks; Figures D.12 to D.14 those to global shocks; Figures D.15 to D.17 foreign responses to Euro Area and US demand and MP shocks; Figures D.18 and D.19 the foreign interest rate responses to demand and MP shocks from other countries.

Figure D.8: Impulse responses of output gaps to domestic shocks

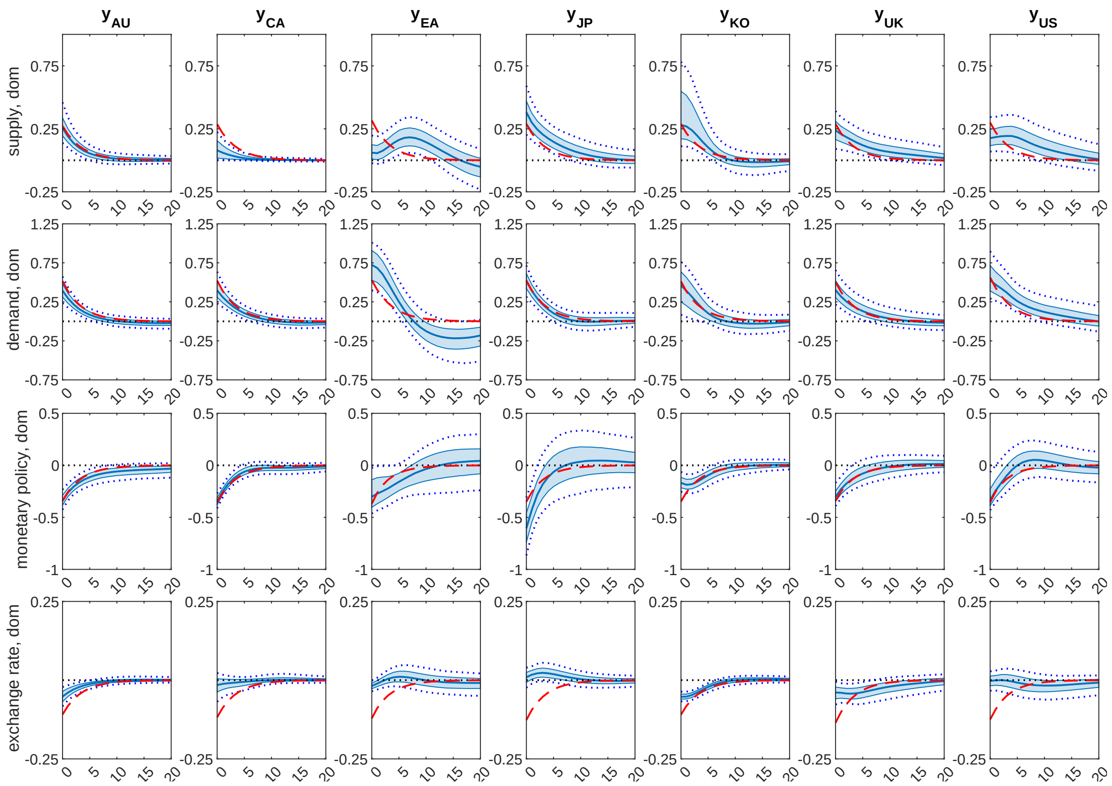


Figure D.9: Impulse responses of inflation to domestic shocks

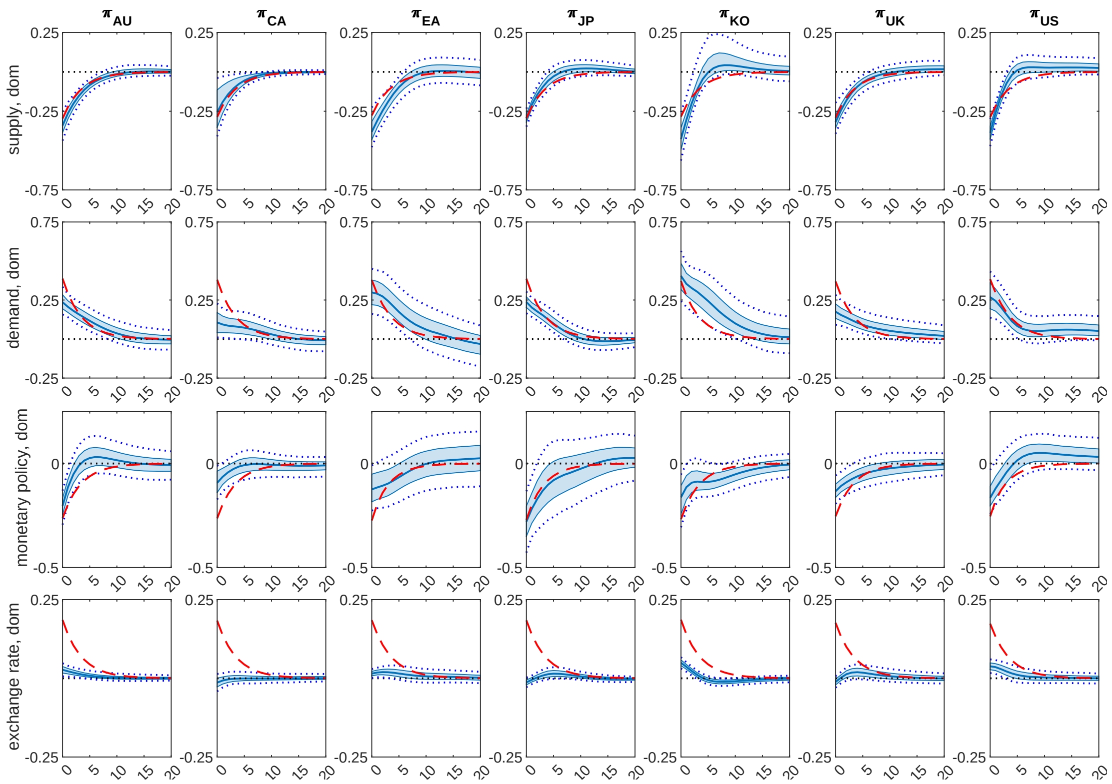


Figure D.10: Impulse responses of interest rates to domestic shocks

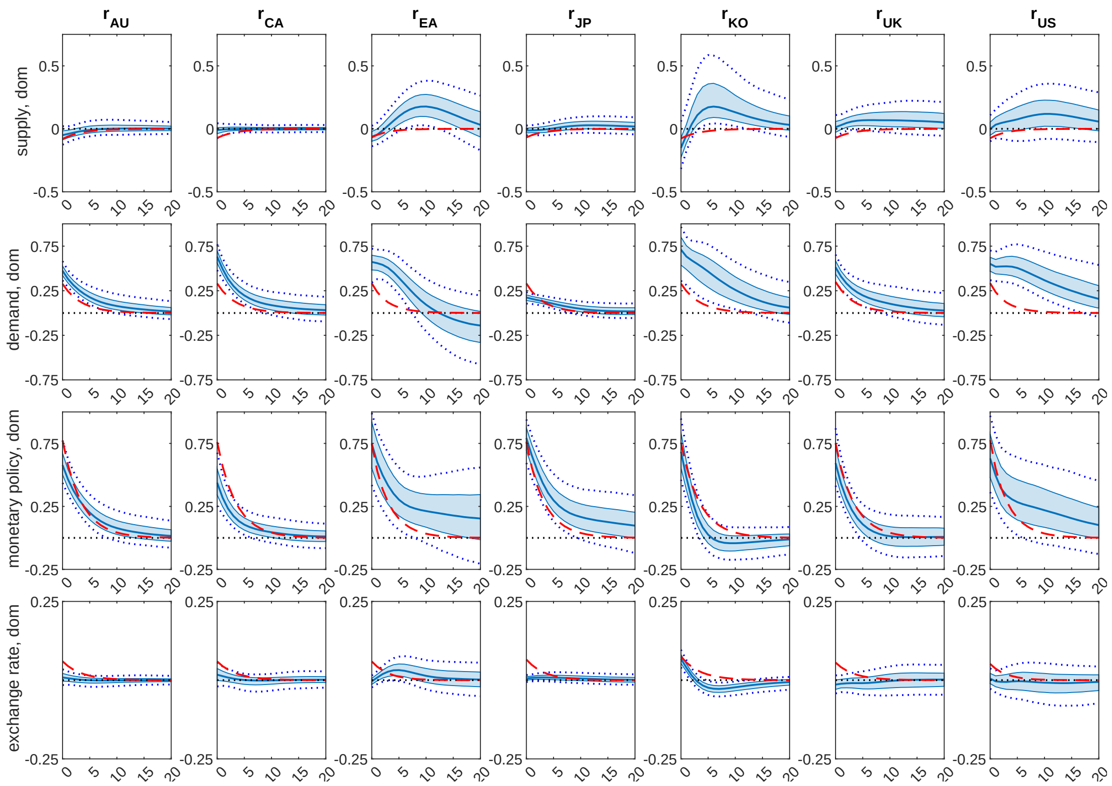


Figure D.11: Impulse responses of exchange rate growth to domestic shocks

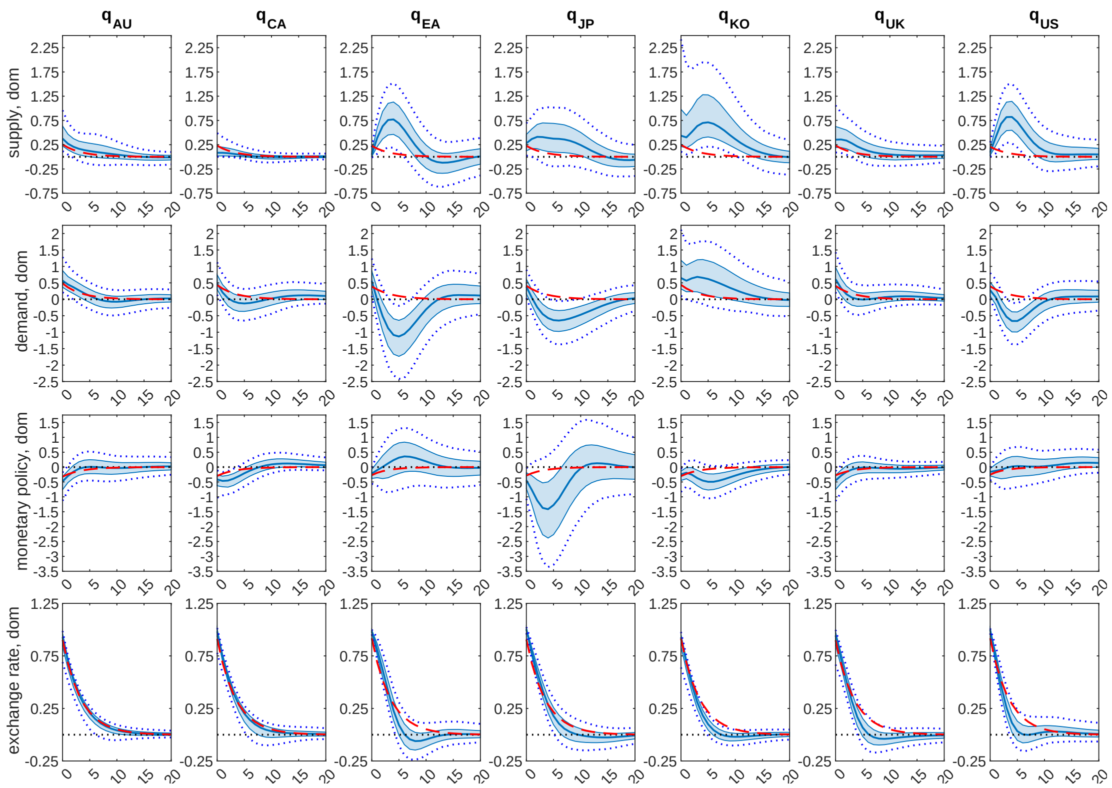


Figure D.12: Impulse responses of output gaps to global shocks

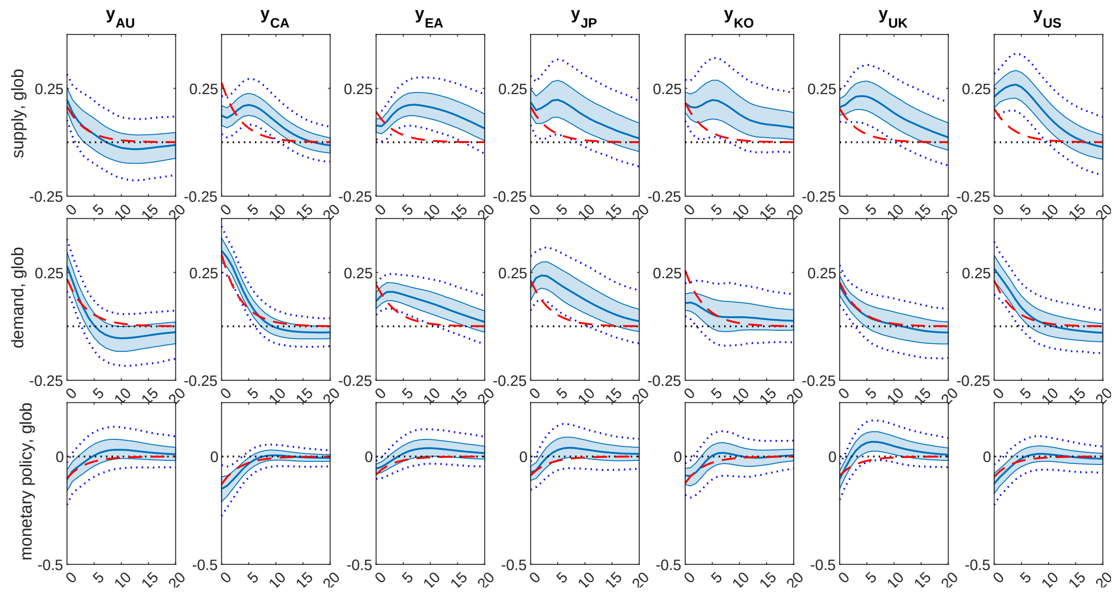


Figure D.13: Impulse responses of inflation to global shocks

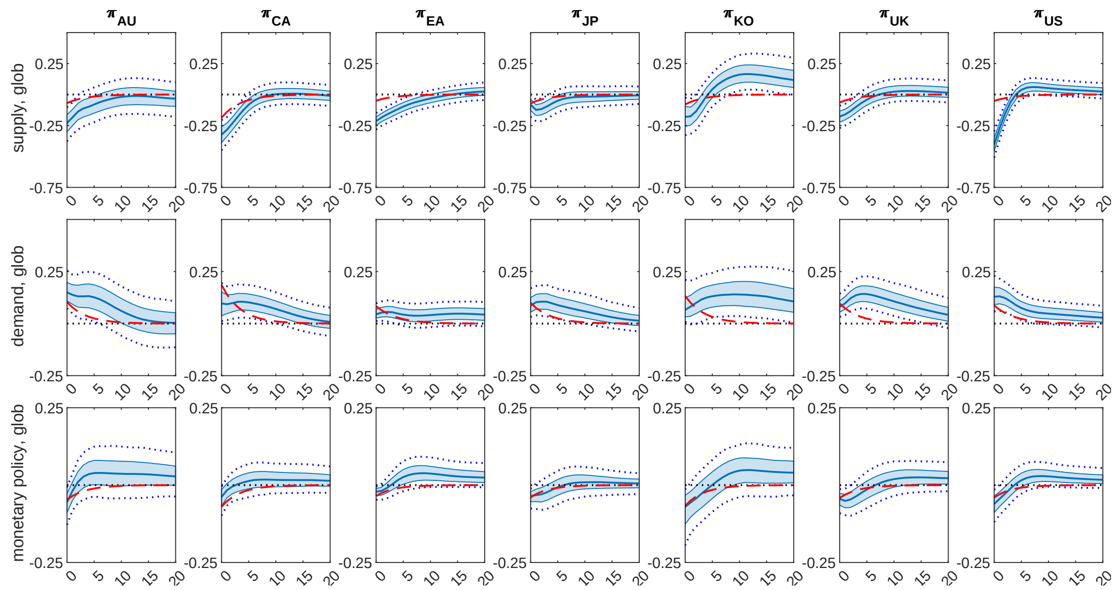


Figure D.14: Impulse responses of exchange rate growth to global shocks

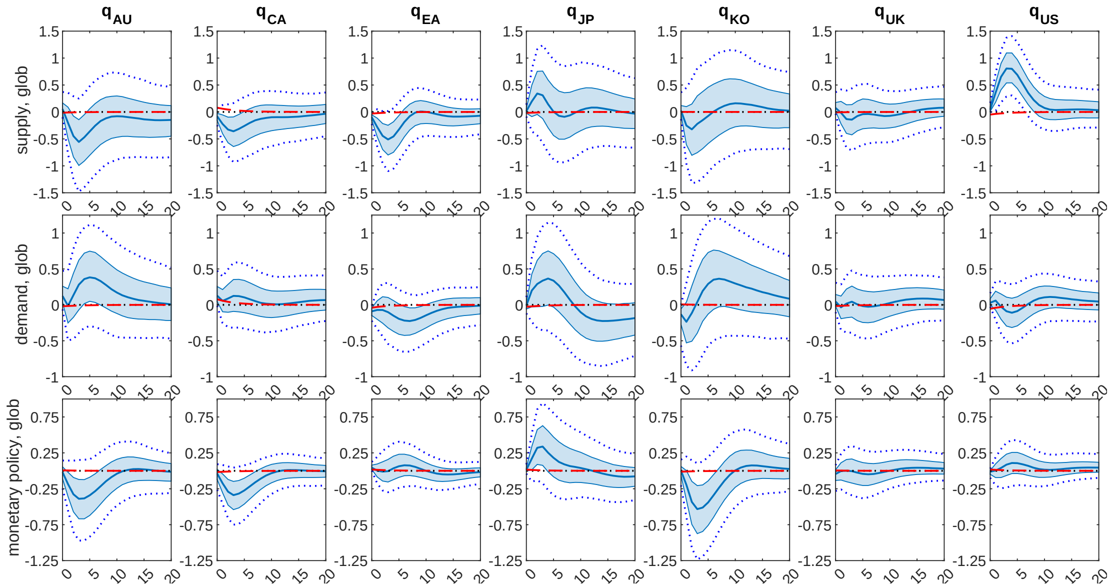


Figure D.15: Impulse responses of output gaps to EA and US demand and MP shocks

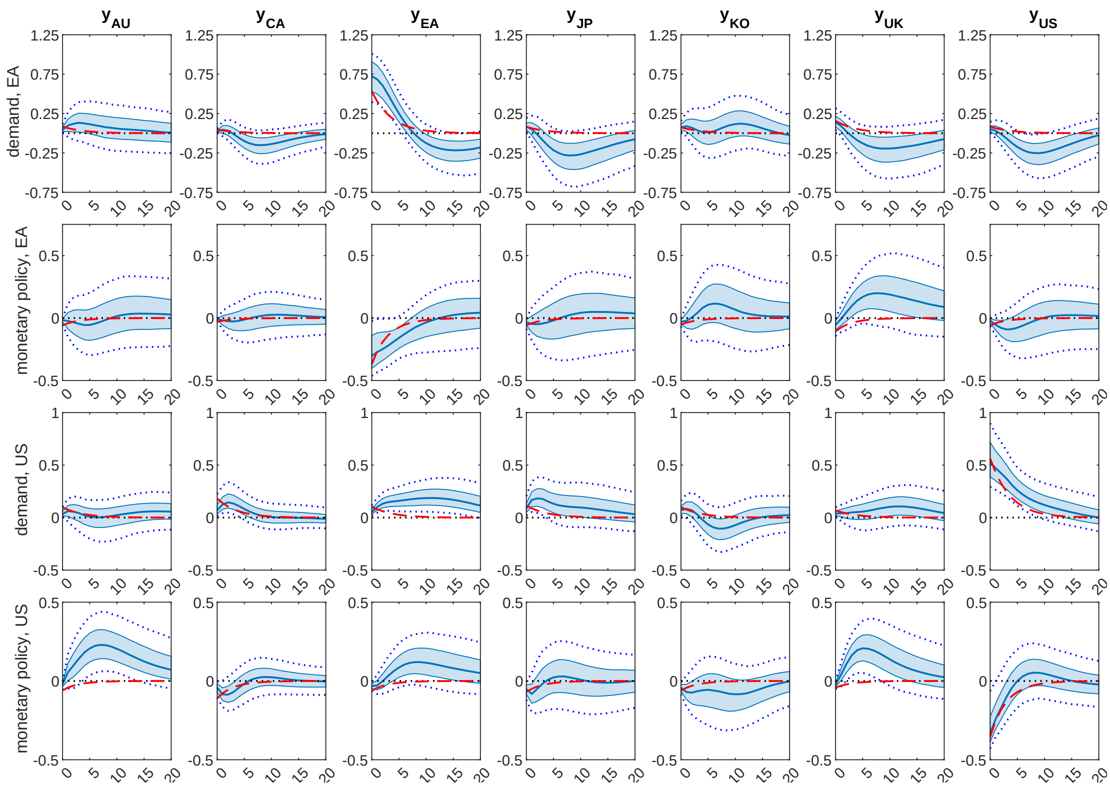


Figure D.16: Impulse responses of inflation to EA and US demand and MP shocks

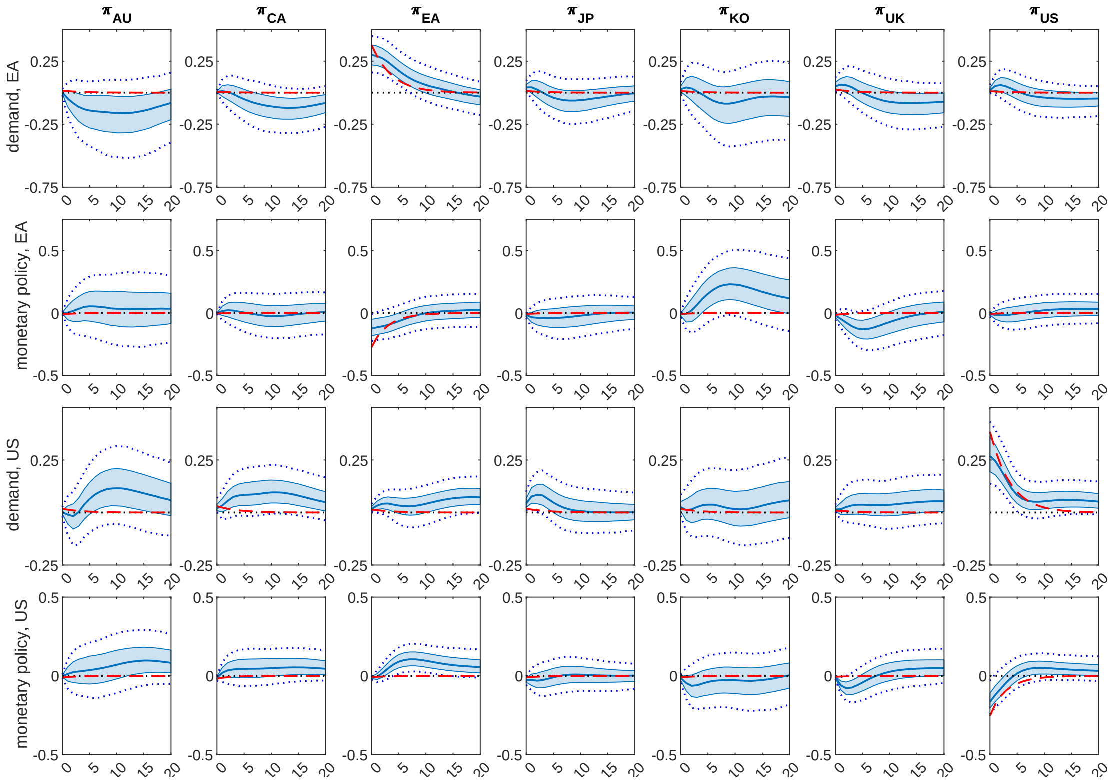


Figure D.17: Impulse responses of exchange rate growth to EA and US demand and MP shocks

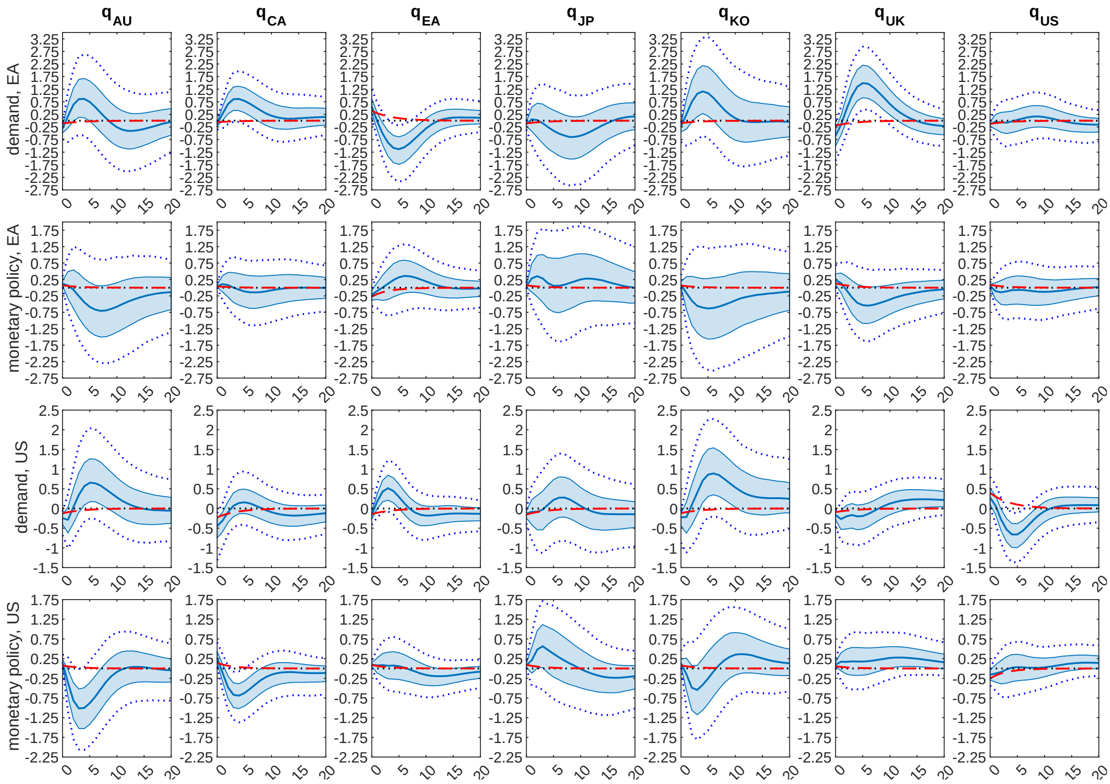


Figure D.18: Impulse responses of interest rates to demand shocks (AU, CA, JP, KO, UK)

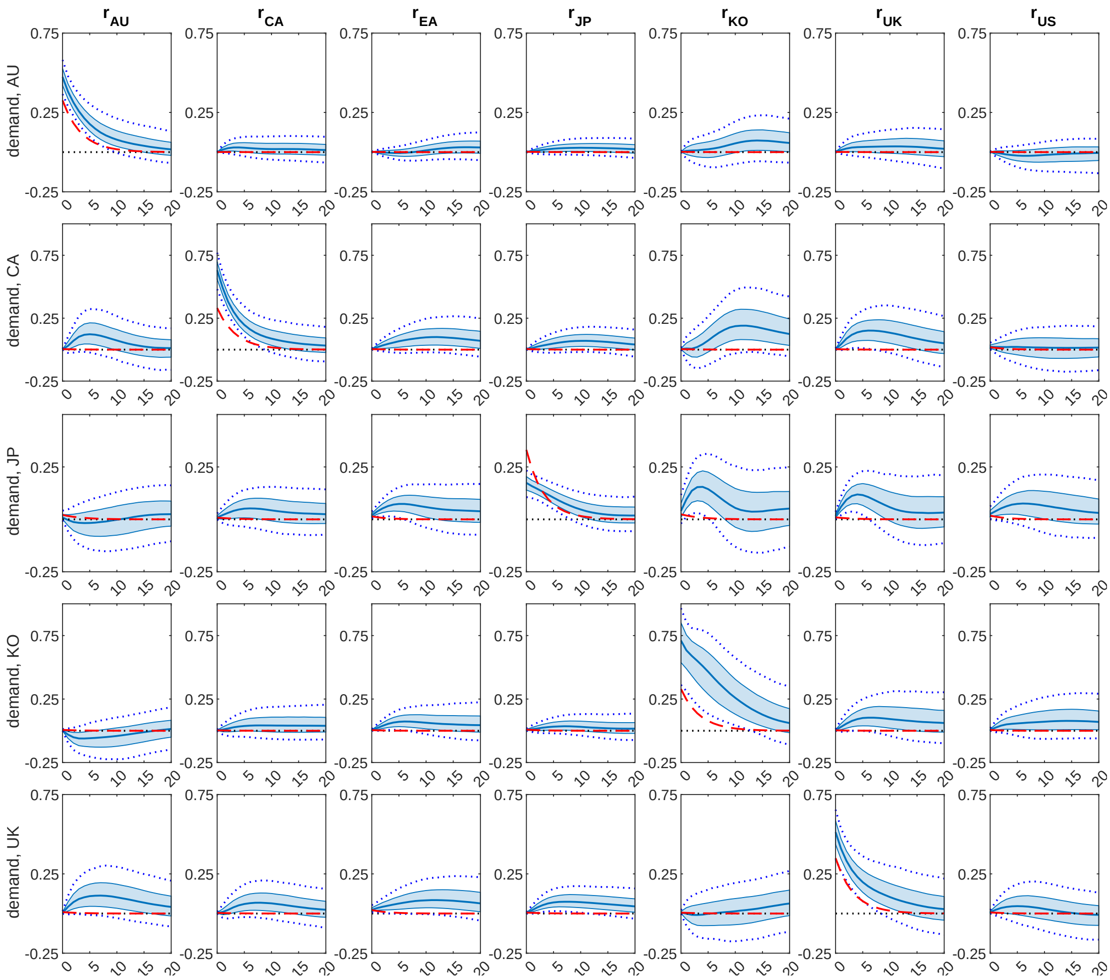


Figure D.19: Impulse responses of interest rates to MP shocks (AU, CA, JP, KO, UK)

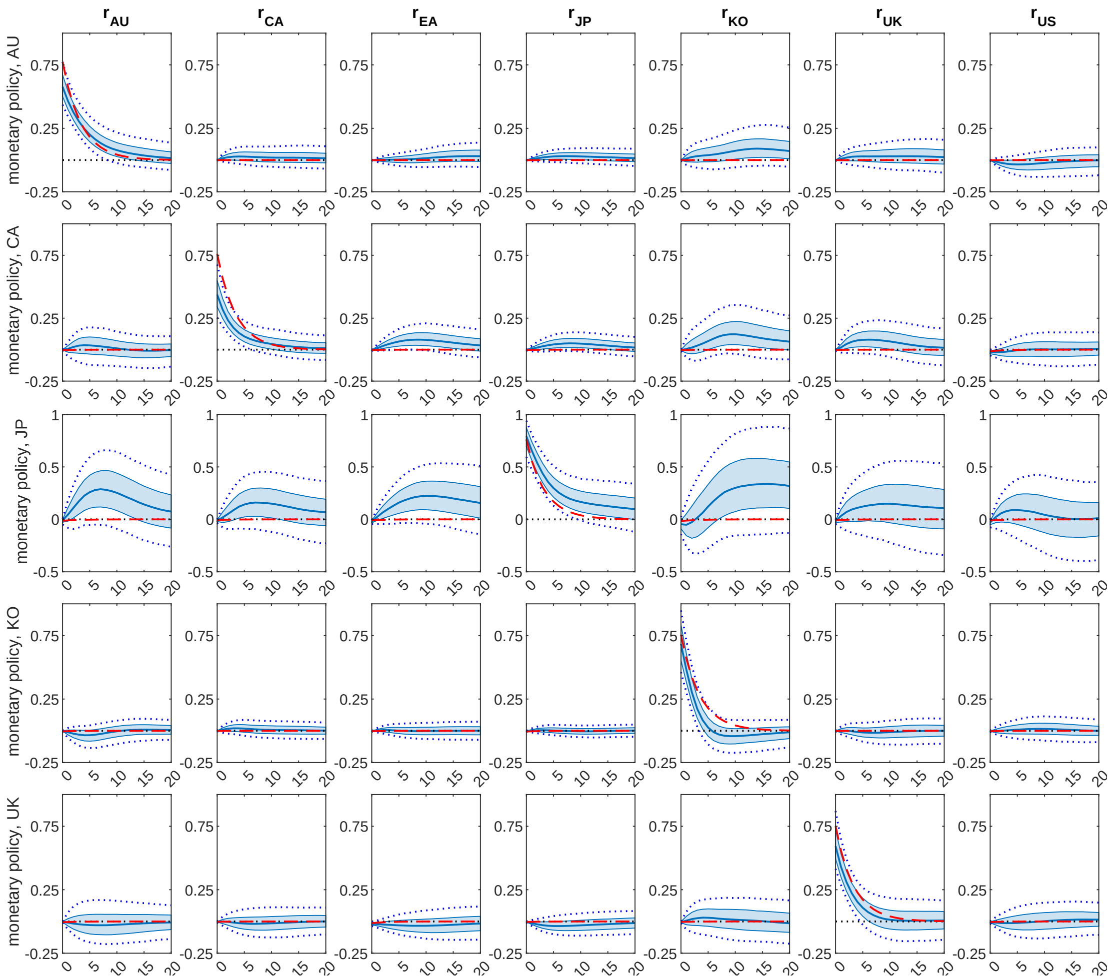
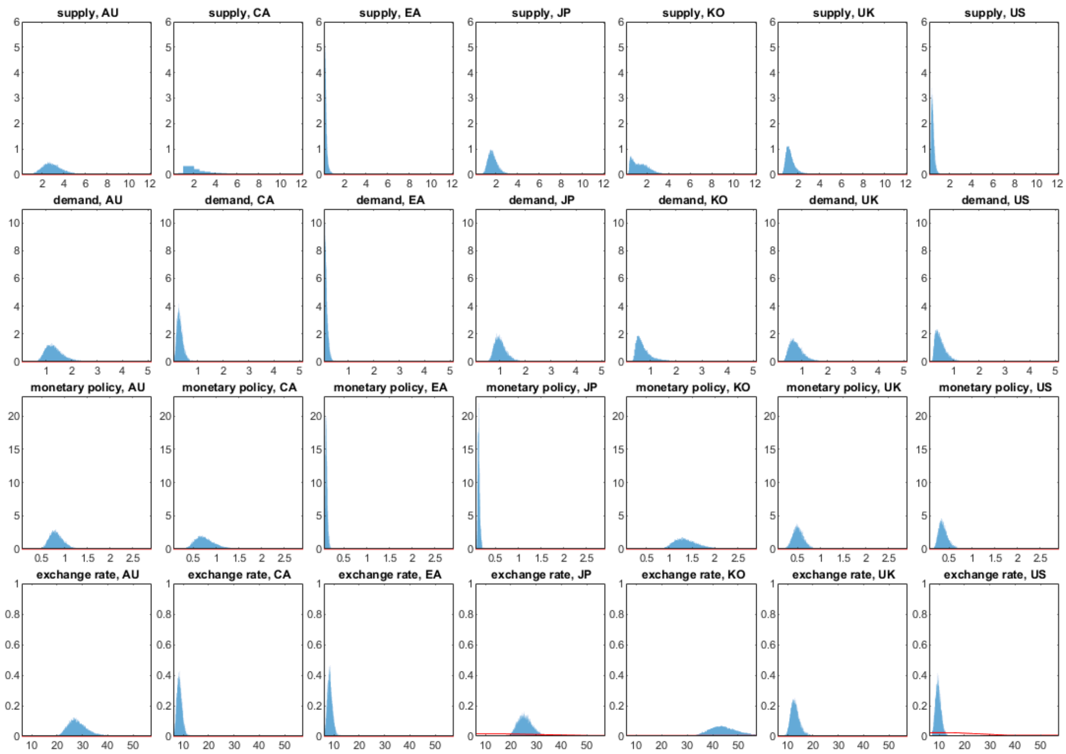


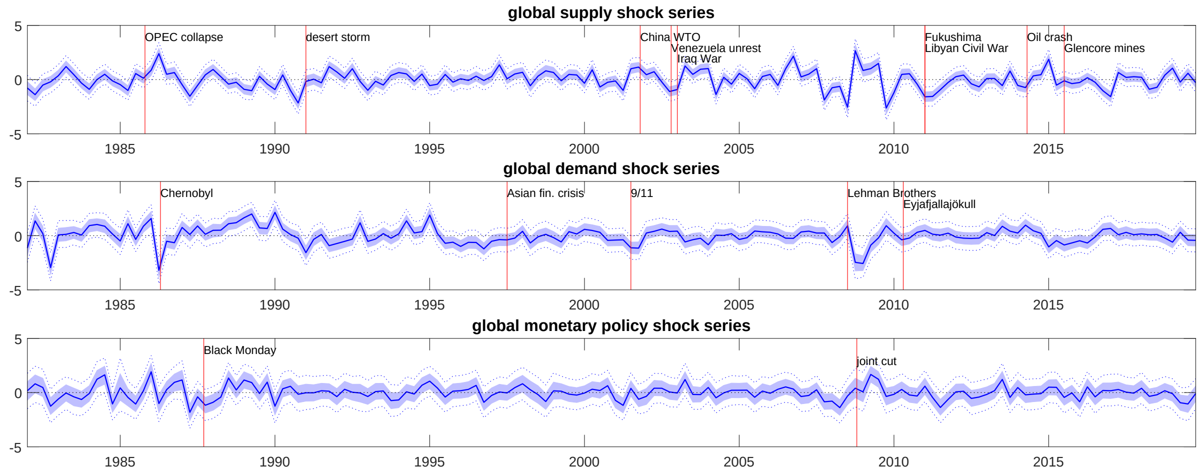
Figure D.20: Variances of structural shocks



NOTES: The histograms show the posterior distribution of variances of structural shocks, together with the uninformative prior distribution (red line). Global shocks have unit variance.

D.4 Global shocks

Figure D.21: Estimated time series of global shocks



NOTES: The solid lines in the figure show median global shock series. The shaded areas (dotted lines) show the respective 68% (95%) posterior credibility sets. Oil supply events are drawn from Antolín-Díaz and Rubio-Ramírez (2018); Känzig (2021) and reference therein.

E Further robustness analysis

This section shows results for several robustness checks changing the model set-up and adjusting the structural contemporaneous relations. We show the forecast error correlation decomposition between global and country-specific interest rates for the alternative specifications in comparison to our main results (baseline). Further results are available on request.

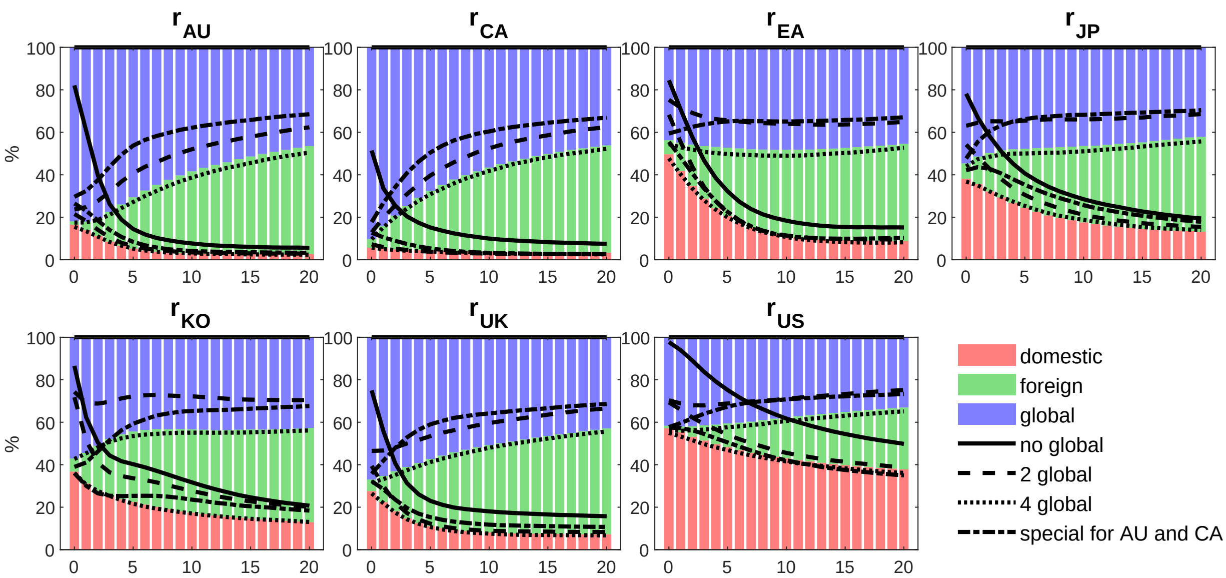
E.1 Alternative global shock specifications

Four specifications change the set-up of the global shocks, shown in Figure E.1: a model without global shocks, a model with two global shocks (demand and supply shocks), a model with four global shocks (additionally a global exchange rate shock), and a model where for Australia and Canada global supply and demand shocks enter the exchange rate equation.

E.2 Alternative model set-ups and samples

Four specifications change the model-set up, shown in Figure E.2: a model with a deterministic trend, a model with 8 lags, a model with no size-dependent shrinkage (no GDP weight included in the Minnesota prior), and a model with log differences of real GDP instead of

Figure E.1: Forecast error correlation decomposition between global and country-specific interest rates for alternative global shock specifications



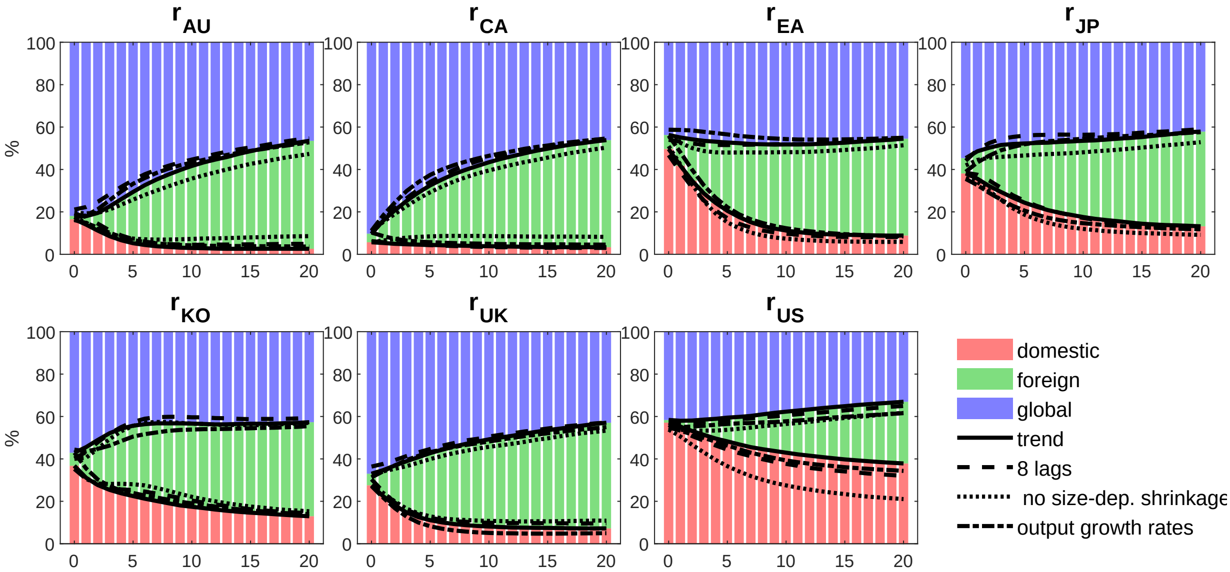
NOTES: Figure shows the forecast error correlation decomposition between global and country-specific interest rates for the baseline model (bars) together with a model without global shocks (solid lines), a model with two global shocks (dashed lines), a model with four global shocks (dotted lines), and a model where for Australia and Canada global supply and demand shocks enter the exchange rate equation (dash-dotted lines).

output gap. Three additional specifications use shorter samples, shown in Figure E.3: a model with data starting in 1999:Q1, a model with data ending in 2007:Q2, and a model with data ending in 2007:Q2 that additionally excludes Japan.

E.3 Alternative AS and ER equation specifications

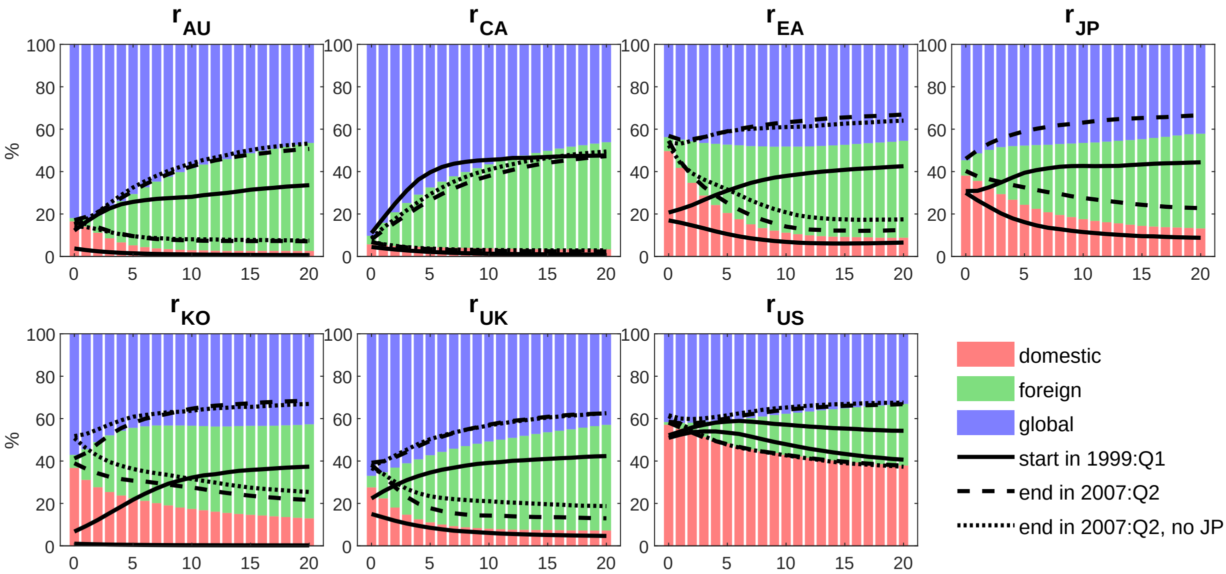
Four specifications adjust the structural contemporaneous relations of supply and exchange rate equations in the baseline model, shown in Figure E.4: The first adds foreign output gaps to the supply equation, restricting the coefficient to be positive (labeled "y^* in AS"). The second adjusts the exchange rate equation such that interest rate differentials enter (labeled "UIP"). Our third specification ("UIP, start in 1999Q1") also adds interest rate differentials, but restricts the sample to 1999:Q1 to 2019:Q4. Our fourth specification uses priors with a variance of 1 (instead of 0.4) for the coefficients $\alpha^{c,\pi}, \theta^{c,y}$ (labeled "wide prior").

Figure E.2: Forecast error correlation decomposition between global and country-specific interest rates for alternative model set-ups



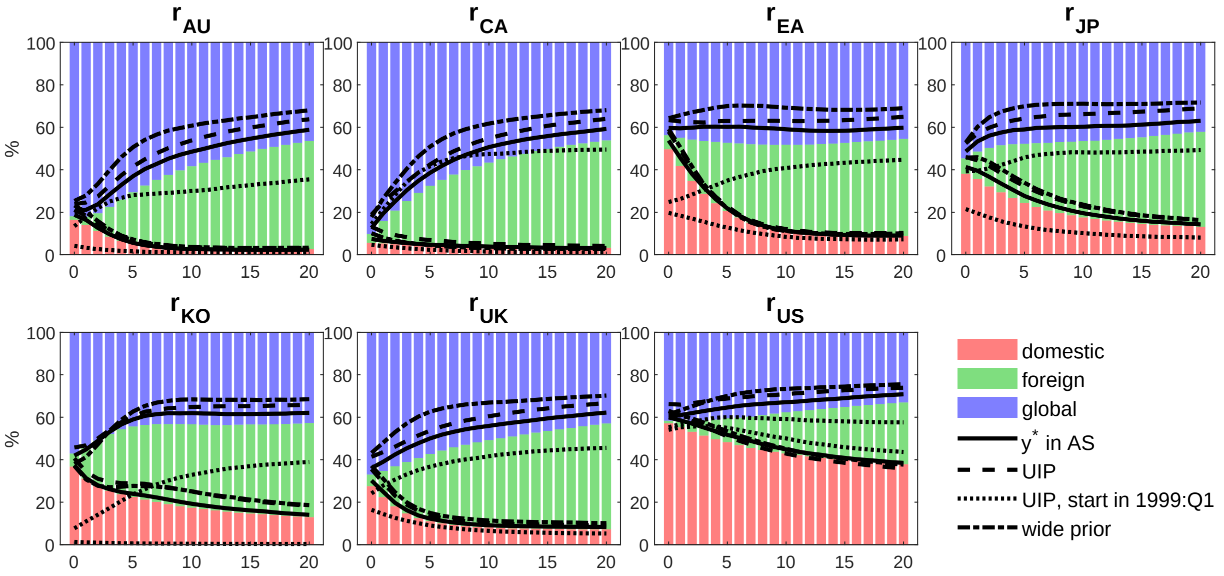
NOTES: Figure shows the forecast error correlation decomposition between global and country-specific interest rates for the baseline model (bars) together with a model with a deterministic trend (solid lines), with 8 lags (dashed lines), no size-dependent shrinkage (dotted lines), and with output growth rates (dash-dotted lines).

Figure E.3: Forecast error correlation decomposition between global and country-specific interest rates with alternative samples



NOTES: Figure shows the forecast error correlation decomposition between global and country-specific interest rates for the baseline model (bars) together with a model with data starting in 1999:Q1 (solid lines), with data ending in 2007:Q2 (dashed lines), and additionally excluding Japan (dotted lines).

Figure E.4: Forecast error correlation decomposition between global and country-specific interest rates for alternative AS and ER equation specifications



NOTES: Figure shows the forecast error correlation decomposition between global and country-specific interest rates for the baseline model (bars) together with a model including foreign output gaps in the supply equation (solid lines), a model with an interest rate differential in the exchange rate equation (dashed lines), the same model starting in 1999:Q1 (dotted lines), and a model with wide priors (dash-dotted lines).

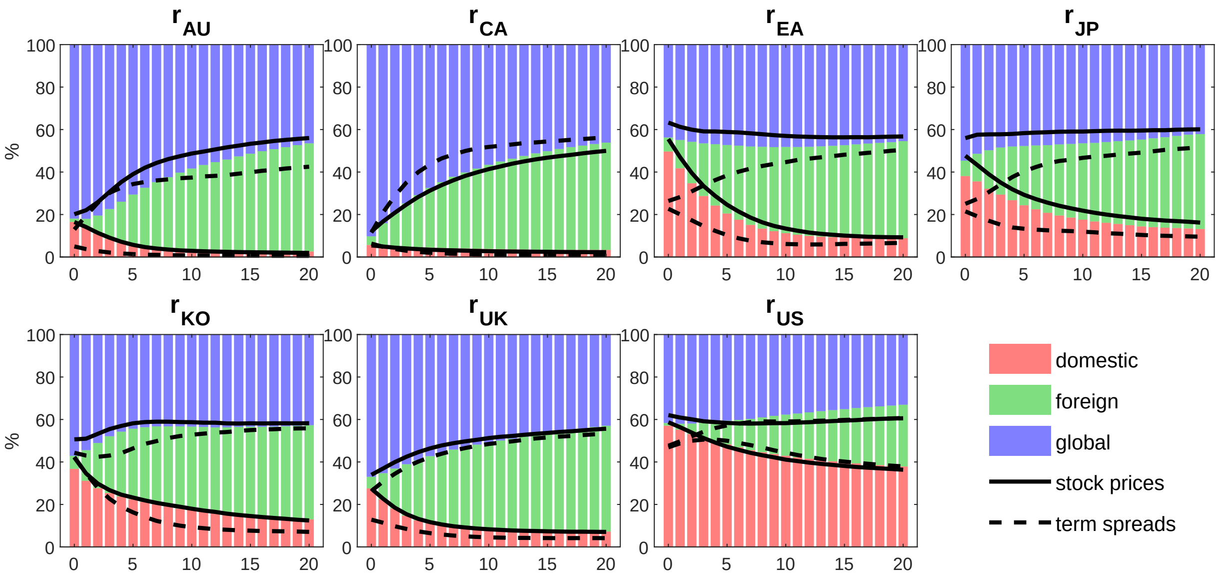
E.4 Accounting for financial transmission channels

Monetary policy shocks might be transmitted internationally via financial channels, such as the wealth and credit channel (for a description of different financial channels see, for example, Bauer and Neely, 2014; Neely, 2015; Fratzscher et al., 2018). The central banks' actions can alter asset prices by stimulating or dampening the demand for assets. Changes in asset prices impact the wealth of households and companies leading to adjusted spending (wealth channel). The monetary authority actions can affect the availability of credit in the market which in turn alters spending and investments (credit channel).

We augment our model by growth in stock prices (wealth channel) or term spreads (credit channel), measured as 5-year yields minus shadow interest rates. We include the additional channel variable (domestic and foreign) contemporaneously in all baseline equations and allow all variables to contribute contemporaneously to the development of the channel variable, with the exception of foreign exchange rates and foreign channel variables. We set Student t priors with mode zero and scale one on the additional contemporaneous parameters. We add prior beliefs that contractionary monetary policy increases domestic interest rates, lowers domestic output gaps and inflation, and decreases competitiveness, by setting asymmetric t distributions with $\mu = 0$, $\sigma = 1$, $v = 3$, and $\lambda = 20$ on the impact effects of monetary policy shocks. We restrict the impact response of the stock prices or term spreads to a domestic monetary policy shock to be negative.

Adding stock prices does not alter the main results substantially, while including term spreads strengthens the importance of global shocks for all countries except Canada, see Figure E.5.

Figure E.5: Forecast error correlation decomposition between global and country-specific interest rates with additional variables

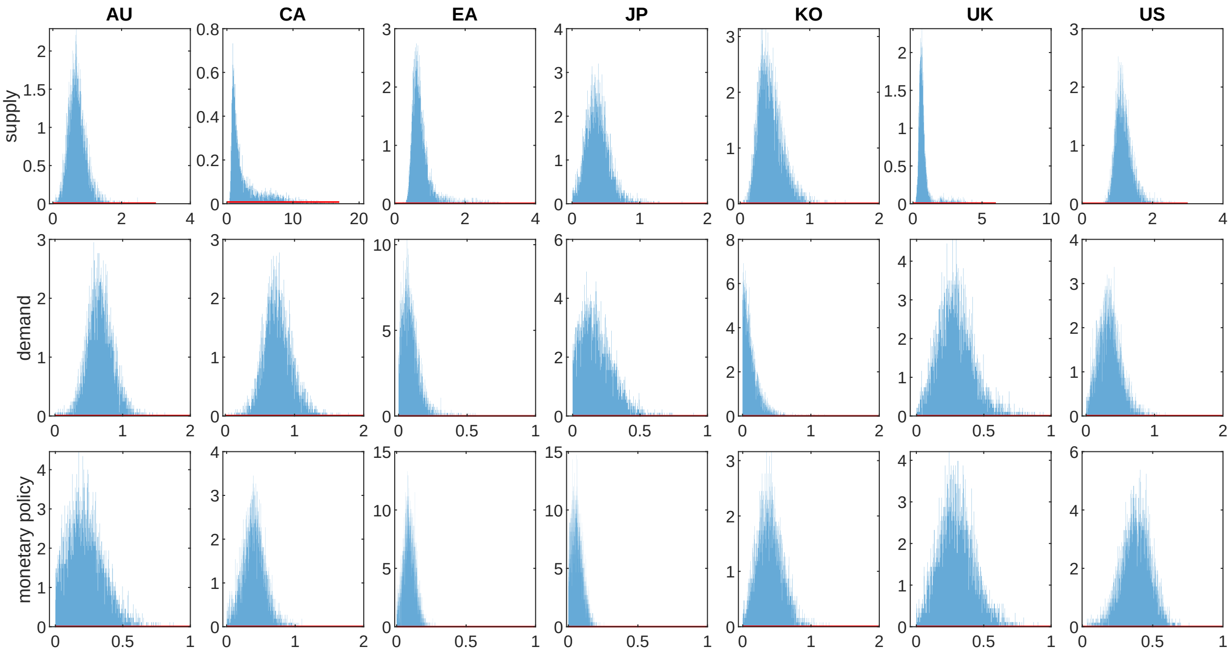


NOTES: Figure shows the forecast error correlation decomposition between global and country-specific interest rates for the baseline model (bars) together with a model including stock prices (solid lines) and a model including term spreads (dotted lines).

E.5 Including the global financial cycle

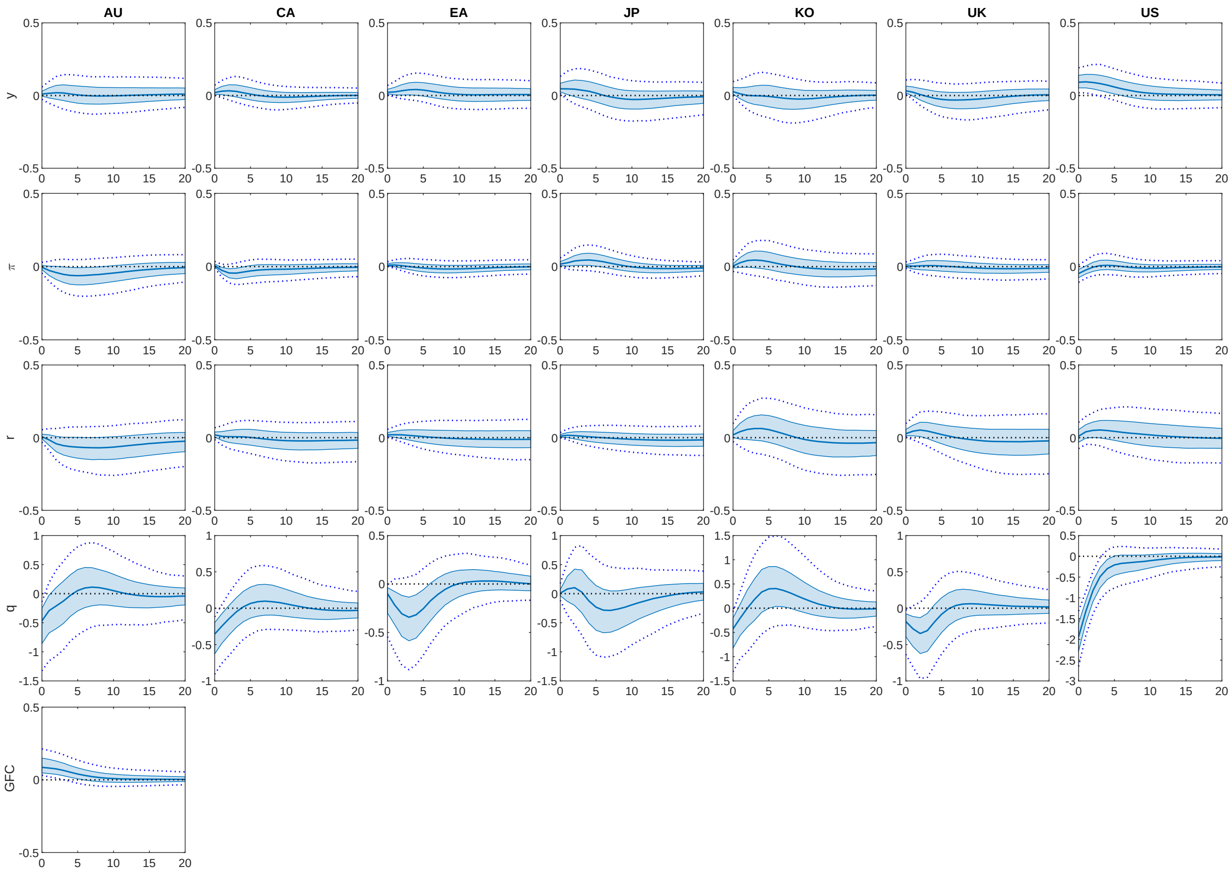
For the specification including the global financial cycle (GFC), we provide additional evidence on the loadings of global shocks in Figure E.6 and on impulse responses to the residual shock to the GFC.

Figure E.6: Posterior distributions of loadings, model including GFC



NOTES: The histograms show the posterior distribution of global shock loadings from the model including GFC, together with the prior distribution (red line).

Figure E.7: Impulse responses of all variables to the residual shock to the GFC



NOTES: The solid lines in the figure show median impulse responses of all variables (in rows) from all countries (in columns) to a residual shock to the global financial cycle over 20 quarters. The shaded areas (dotted lines) show the 68% (95%) posterior credibility sets. The shocks have size of one unit (i.e., one percentage point).