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Keywords

structural vector autoregressions, small open economies, monetary policy shocks, exchange rates, delayed overshooting, uncovered interest rate parity, spillovers of US monetary policy

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Dornbusch overshooting, UIP, and the systematic components of domestic and foreign monetary policy in SVARs*

Nicolas Groshenny[†] Naveed Javed[‡]

January 26, 2025

Abstract

We highlight the importance of *jointly* identifying domestic and foreign monetary shocks in SVAR-based evaluations of Dornbusch exchange rate overshooting and uncovered interest parity (UIP) in small open economies (SOEs). We estimate SVAR models for six developed SOEs to understand the effects of SOE and US monetary policy shocks on bilateral SOE/US exchange rates. Our novel identification strategy features block exogeneity combined with sign restrictions imposed on the coefficients of the SOE and US monetary policy rules. Crucially, the response of the exchange rate to monetary shocks is not restricted, and the SOE policy rate and the exchange rate are allowed to interact instantaneously. Exchange rate dynamics triggered by monetary shocks are found to be broadly in line with Dornbusch overshooting and UIP. We demonstrate that the few cases in which US monetary shocks trigger delayed overshooting in specific samples may not be inconsistent with UIP, but rather an artifact of SOE endogenous monetary policy responses to the US policy rate.

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1 Introduction

How does the exchange rate react to monetary policy shocks? This classic question in international finance largely remains unsettled. On the one hand, the benchmark theoretical result of Dornbusch (1976) predicts that a surprise monetary tightening will cause the exchange rate to overshoot *on impact*, displaying an *instantaneous* appreciation followed by a gradual depreciation. To this day, Dornbusch's overshooting model remains at the core of a wide range of open-economy Dynamic Stochastic General Equilibrium models featuring rational expectations, uncovered interest parity (UIP) and price stickiness (Lane, 2001; Corsetti, 2008). On the other hand, a vast empirical literature estimating Structural Vector Autoregressions (SVARs) often finds evidence of a *gradual and persistent* appreciation, typically lasting for more than a year, in response to a contractionary domestic monetary shock. Such hump-shaped, rather than immediate, empirical responses of the real exchange rate (RER) are referred to as the 'delayed overshooting puzzle'.[1,2] While delayed overshooting is associated with the *timing* of the peak response of the exchange rate, UIP is related to the *quantitative* dynamics of the exchange rate explained by interest rate differentials. Faust and Rogers (2003) assert that "*the UIP element of the Dornbusch model is most problematic empirically*". Deviations from UIP conditional on monetary shocks (dubbed the 'forward discount puzzle') and their relation with delayed overshooting has been discussed extensively in the literature.[3] We build on this branch of literature by demonstrating that our *joint* identification of domestic (SOE) and foreign (US) monetary shocks is crucial to interpret the link between delayed overshooting and UIP.

Specifically, we contribute towards the debate sparked by Scholl and Uhlig (2008): the forward discount puzzle is a "*twin appearance*" of delayed overshooting. Their influential finding that "*the forward discount puzzle is robust even without delayed overshooting*" has been recently challenged by Rüth and Van der Veken (2023) who

[1] See Eichenbaum and Evans (1995), Cushman and Zha (1997), Bagliano et al. (1999), Kim and Roubini (2000) and Kim et al. (2017) among others.

[2] The pervasiveness of delayed overshooting has convinced some DSGE modelers to alter the UIP condition to obtain hump-shaped responses of exchange rate to monetary shocks, e.g. Adolfson et al. (2008).

[3] See Faust and Rogers (2003), Scholl and Uhlig (2008), Rüth and Van der Veken (2023) and Müller et al. (2024).

report that "*UIP may be intact even if foreign exchange rates overshoot*".[4] Both Scholl and Uhlig (2008) and Rüth and Van der Veken (2023) are studies based on US data. In contrast, we examine conditional UIP in SOEs, and also differentiate between exchange rate dynamics in periods before and including the Global Financial Crisis (GFC). Our findings for SOEs in the pre-GFC sample are in line with those of Rüth and Van der Veken (2023) for the US: UIP is not violated even when the exchange rate overshooting is *mildly* delayed. However, we go a step further and demonstrate that our *joint* identification scheme plays a key role in highlighting that the slight delay in overshooting can be attributed to the endogenous responses of SOE central banks' to US monetary shocks. To sum up, delayed overshooting continues to be consistent with UIP, and the delay in the exchange rate response may not be a puzzle.

Our novel findings rest on three features that make our econometric strategy especially suitable to investigate the robustness of delayed overshooting and forward discount puzzle. First, we *jointly* identify SOE and US monetary policy shocks, and investigate whether the responses of exchange rate to monetary shocks differ according to the origin of the shocks.[5] We live in a 'dollar world' (Gourinchas, 2021), and the Federal Reserve is the main driver of global funding costs (Miranda-Agrippino and Rey, 2020). This fundamental unevenness in the global financial system reflects in the asymmetries between the monetary policy reaction functions of the Federal Reserve and the SOE central bank. Our approach recognizes that a SOE central bank reacts endogenously to US monetary shocks, so that the SOE-US policy rate differential may be affected differently by SOE and US monetary shocks.[6] In fact, UIP implies that the entire expected path of the SOE-US interest rate differential determines the response of the bilateral exchange rate to a monetary shock (Engel, 2014; Galí, 2020). Second, our strategy allows for instantaneous interactions between the exchange rate and policy rate. This is important as Faust and Rogers (2003) and Bjørnland (2009) argue that

[4] Kim et al. (2017) revisit Scholl and Uhlig (2008) and find that delayed overshooting occurs in conjunction with UIP failure during Volcker's era.

[5] Several papers focus on US monetary shocks (Eichenbaum and Evans, 1995; Faust and Rogers, 2003; Scholl and Uhlig, 2008; Kim et al., 2017; Rüth, 2020; Castelnuovo et al., 2022; Rüth and Van der Veken, 2023). Others study the impacts of non-US monetary shocks (Cushman and Zha, 1997; Kim and Roubini, 2000; Bjørnland, 2009; Bjørnland and Halvorsen, 2014; Kim and Lim, 2018; Terrell et al., 2023). Earlier papers consider relative money shocks without taking a stance on the origin of disturbances (Clarida and Gali, 1994; Rogers, 1999).

[6] See also Davis and Zlate (2019).

delayed overshooting may be an artifact of recursive short-run restrictions that do not allow for simultaneous interactions between the exchange rate and money market rates.[7] Third, our approach is agnostic, in the sense that it leaves the response of the exchange rate to monetary shocks unrestricted at all horizons.

We estimate Bayesian SVAR models for six advanced SOEs with floating exchange rates and inflation targeting central banks (Australia, Canada, New Zealand, Norway, Sweden, and the United Kingdom) to measure the effects of SOE and US monetary shocks on bilateral SOE/US real exchange rates. Our identification scheme combines two key ingredients. First, our SVAR model features a block-exogenous structure, meaning that we classify the variables of our model into a US block and a SOE block. The US block influences the SOE block both contemporaneously and over time, whereas the SOE block has no effect on the US block (Cushman and Zha, 1997). Second, we *jointly* identify the policy rules of the Federal Reserve and the SOE central bank using a set of sign restrictions imposed directly on the policy parameters (Arias et al., 2019).

We characterize the systematic component of US monetary policy by imposing that the Federal funds rate responds positively to US output and inflation, and negatively to the Baa credit spread (Baumeister and Hamilton, 2018; Caldara and Herbst, 2019; Rüth and Van der Veken, 2023).[8] For the SOE central bank, in line with Taylor (2001), we assume that it follows an augmented Taylor-type rule that reacts positively to output, inflation and the real SOE/US exchange rate.[9] Hence, we require that the SOE monetary authority does not usually exacerbate exchange rate fluctuations by raising its policy rate in response to an appreciation. This restriction embodies policymakers' rules of thumb that monetary policy should generally lean against the real exchange rate, and real appreciations are opportunities to ease monetary conditions (Obstfeld and Rogoff, 1995). This restriction is supported by the findings of Bjørnland (2009) and Bjørnland and Halvorsen (2014) based on SVAR models. It is also consistent with some estimates of Taylor-type rules in SOE-DSGE models (Lubik and Schorfheide,

[7] Rüth (2020) stresses the importance of a full contemporaneous interplay between all variables in the SVAR system.

[8] Curdia and Woodford (2010, 2016) present normative analyses that justify a systematic easing of monetary policy in response to tighter credit conditions.

[9] Calvo and Reinhart (2002), Reinhart and Rogoff (2004), Obstfeld (2013) and Ilzetzki et al. (2019) provide evidence that central banks react to movements in the bilateral dollar exchange rate.

2007; Justiniano and Preston, 2010). In line with Braig et al. (2024), we demonstrate that this restriction is essential for exchange rate overshooting.

Our baseline findings, based on the period 1992:Q1-2019:Q4, provide little evidence of delayed overshooting. In the six SOEs, a contractionary monetary shock triggers a strong and immediate exchange rate appreciation followed by a gradual depreciation. Symmetrically, a tightening of US monetary policy causes an instantaneous exchange rate depreciation followed by a gradual appreciation. In almost all cases, the peak response of the exchange rate occurs on impact or shortly after. Furthermore, we do not observe any compelling evidence of the forward discount puzzle either: conditional on SOE or US shocks, exchange rate dynamics broadly offset the SOE-US interest rate differentials. Thus, our baseline findings contrast sharply with Faust and Rogers (2003) and Scholl and Uhlig (2008) who, like us, consider set-identified SVARs, but find large deviations from UIP conditional on US shocks, and argue that the forward discount puzzle is robust with or without delayed overshooting. Instead, our results are consistent with the findings of Bjørnland (2009) regarding the propagation of SOE monetary disturbances, and Rüth (2020) and Rüth and Van der Veken (2023) in terms of the open-economy implications of US monetary shocks.

We then distinguish between the exchange rate impacts of conventional and unconventional monetary policy, both by the Federal Reserve and some SOE central banks. To do so, we split the sample into two sub-periods, 1992:Q1-2008:Q3 and 2008:Q4-2019:Q4, and re-estimate the six SOE-SVARs over each sub-period. We obtain the following results: First of all, conditional on US or SOE monetary shocks, UIP broadly holds in both sub-periods. Second, conditional on SOE shocks, exchange rates display overshooting à la Dornbusch in both sub-periods. Third, conditional on US shocks, we observe *mildly* delayed overshooting in the pre-GFC sample, but not in the sample following the GFC. We attribute these findings to a change, across the two periods, in the endogenous responses of SOE central banks to US monetary shocks.[10] Put differently, SOE central banks reacted differently to US conventional and unconventional monetary shocks.

[10] Cushman and Zha (1997) show that in the SOE setting, explicitly accounting for the endogenous policy reactions is essential for UIP to hold. In the same vein, McCallum (1994) demonstrates that serious UIP violations occur in regressions that ignore endogenous monetary policy reactions to interest rates and exchange rates.

For our pre-GFC sample, we observed that SOE central banks tended to mimic the Fed's actions. Such behavior contributed to keep the SOE-US policy rate differentials narrow in the aftermath of US conventional monetary shocks. As a result, in accordance with UIP, instantaneous reactions of the exchange rates to US shocks appear muted. This conveys the impression of slightly delayed overshooting. On the other hand, during 2008:Q4 to 2019:Q4 when the Fed switched to unconventional monetary policy, we observed that SOE central banks tended to move in the opposite direction to the Fed. Hence, US unconventional monetary shocks generated larger SOE-US policy rate differentials and, in line with UIP, exchange rate displayed clear-cut instantaneous overshooting. Summing up, the estimations over these sub-periods highlight that UIP may be intact even if the exchange rate does not overshoot instantaneously (Rüth and Van der Veken, 2023; Müller et al., 2024).

The rest of the paper is structured as follows. The next section places the contributions of the paper against the backdrop of the related literature. Section 3 describes the data, the identification scheme and the Bayesian estimation. Sections 4 and 5 present our main results, Section 6 contains robustness checks while Section 7 concludes.

2 Related literature

Our paper lies at the interface of two distinct strands of the SVAR literature on the effects of monetary policy on the exchange rate. One strand of the SVAR literature focuses on the US. Eichenbaum and Evans (1995) document the delayed overshooting puzzle. They employ a recursive identification scheme and find evidence of a gradual and persistent appreciation in both the nominal and real US exchange rates in response to a contractionary US monetary policy shock.[11] Their findings contradict the fundamental prediction of the Dornbusch hypothesis; that the exchange rate would overshoot instantaneously. Further studies by Faust and Rogers (2003) and Scholl and Uhlig (2008) replace the controversial recursive identification scheme with sign restrictions on the impulse response functions. However, these studies again document puzzling responses with delays lasting around 3 years.

[11]With nominal rigidities, the responses of the real and nominal exchange rates are similar in the short run.

In contrast, Kim et al. (2017) using sign restrictions similar to Scholl and Uhlig (2008) report findings consistent with Dornbusch's prediction except during Volcker's tenure as Federal Reserve Chairman. Rüth (2020) uses surprises in Federal funds futures around policy announcements as external instruments to estimate a proxy-SVAR model and measure the effects of US monetary policy shocks on various measures of US exchange rates. His findings are consistent with Dornbusch's predictions, including during Volcker's tenure. Castelnuovo et al. (2022) identify their SVAR model by applying restrictions on IRFs and structural parameters of the systematic component of US monetary policy. They find no evidence of the delayed overshooting puzzle. Our study complements that of Castelnuovo et al. (2022), but identifies SOE monetary policy shocks in addition to US monetary policy shocks by implementing block exogeneity. Moreover, while Castelnuovo et al. (2022) impose a combination of zero and sign restrictions on both IRFs and policy coefficients, we impose restrictions only on policy coefficients. Rüth and Van der Veken (2023) employ a hybrid identification scheme by combining sign restrictions on IRFs and structural parameters with narrative sign restrictions in order to examine overshooting hypothesis for US during the pre-GFC period. Although they find evidence of delayed overshooting, little empirical support for forward discount premia is reported.[12]

Another strand of the SVAR literature focuses on SOEs. Cushman and Zha (1997) and Kim and Roubini (2000) apply non-recursive zero restrictions to implement block exogeneity and identify monetary policy shocks. Bjørnland (2009) uses data from four SOEs to estimate an SVAR model combining short-run and long-run zero restrictions. Her identification scheme allows for simultaneous interactions between monetary policy and the exchange rate while requiring that monetary shocks have no impact on the real exchange rate in the long run. She finds no evidence of delayed overshooting, suggesting that Dornbusch was right after all. Recently, Terrell et al. (2023) estimate a time-varying SVAR model with stochastic volatility using the same data and identification scheme as Bjørnland (2009). Their results are in line with Bjørnland (2009). Other studies applying agnostic identification procedures to analyse the exchange-rate response to monetary shocks in SOEs include Bjørnland and Halvorsen (2014)

[12] Also see Yang et al. (2024), Müller et al. (2024) and Braig et al. (2024).

and Kim and Lim (2018).[13] Bjørnland and Halvorsen (2014) consider six SOEs and identify monetary disturbances by imposing a combination of sign and exclusion restrictions on IRFs. Unlike us, they impose a sign restriction on the impact response of the exchange rate, forcing an instantaneous appreciation, and thus ruling out the so-called exchange rate puzzle by construction. They do not find evidence of delayed overshooting. Kim and Lim (2018) consider four SOEs and achieve identification by imposing sign restrictions on IRFs. They confirm the findings of Bjørnland and Halvorsen (2014).

Our contribution lies in *jointly* identifying the systematic components of US and SOE monetary policy rules through the block exogeneity structure (Cushman and Zha, 1997) and selected sign restrictions imposed directly on the policy parameters (Arias et al., 2019). The imposition of block exogeneity, a key building block of the *joint* identification strategy, hinges on our novel adaptation of the methodology of Arias et al. (2019) in two ways. First, unlike Arias et al. (2019), we are less constrained by the number of exclusion restrictions as we use highly information priors centered at zero on the lagged SOE variables in the US block (Dieppe et al., 2016). Second, while Arias et al. (2019) use Natural Conjugate Normal Inverse Wishart ($\mathcal{NIW}$) priors, which are not suitable to impose exclusion restrictions on selected set of parameters, we employ Independent $\mathcal{NIW}$ priors, which enables us to impose the block exogeneity structure (Dieppe et al., 2016; Koop et al., 2010). Thus, we put in place a US-SOE framework using our innovative identification scheme, which allows us to reconcile delayed overshooting with UIP. Specifically, we observe *mildly* delayed overshooting in the pre-GFC sample in response to US monetary shocks only, while UIP remains intact. Our identification approach helps us resolve this puzzling result by taking into account the endogenous reactions of SOE central banks to Fed funds rate surprises. We show that delayed overshooting may not be a puzzle, but a consequence of a narrow gap between US and SOE policy rates during the pre-GFC period. Hence, our *joint* identification scheme plays a pivotal role in demonstrating that delayed overshooting and UIP may coexist.

[13] Also see Jääskelä and Jennings (2011), Read (2023) and Fisher and Huh (2023) for related studies focusing on Australia.

3 Econometric strategy

3.1 Data

We consider six advanced small-open economies, namely Australia, Canada, New Zealand, Norway, Sweden, and the UK. We use quarterly data from 1992:Q1 to 2019:Q4.[14] The starting date corresponds broadly to the adoption of inflation targeting by the six SOEs considered here (Kim and Lim, 2018). The end date marks the onset of COVID-19 pandemic, mitigating the risk that the sequence of extreme observations may substantially affect the parameter estimates (Lenza and Primiceri, 2022). Moreover, estimating a SVAR over a stable monetary policy regime helps in solving the delayed overshooting puzzle.[15] Following Cushman and Zha (1997), we organize the variables into two blocks, a domestic one and a foreign one. The domestic block represents the SOE, while the foreign block stands for the US economy. The domestic block includes real GDP (y), inflation (π) measured as the annualized quarterly rate of change in the consumer price index, the policy rate (r) proxied by the 3-month interbank rate, and the bilateral SOE/US real exchange rate (e). For Sweden and the UK, we use shadow rates constructed by De Rezende and Ristiniemi (2023) and Wu and Xia (2016), respectively. The foreign block consists of four variables: US real GDP (y^*), US inflation (π^*), Moody's Baa corporate credit spread (cs^*), and the US shadow rate (r^*) constructed by Wu and Xia (2016). All variables are expressed in log levels except the credit spread, inflation rates and policy rates, which are expressed in percentage points.

3.2 Model

Our structural model is given by:

$$\mathbf{y_t'A_0} = \sum_{l=1}^{p} \mathbf{y}_{t-l}'\mathbf{A}_l + \mathbf{c}' + \epsilon_\mathbf{t}' \ , \qquad \text{for } 1 \leq \mathbf{t} \leq \mathbf{T} \tag{1}$$

where $\mathbf{y_t'} = \left[\mathbf{y_{1t}'} \quad \mathbf{y_{2t}'}\right]$, $\mathbf{y_{1t}'} = [y_t^*, \pi_t^*, cs_t^*, r_t^*]$ and $\mathbf{y_{2t}'} = [y_t, \pi_t, r_t, e_t]$. $\mathbf{y_{1t}}$ is a $(n_1 \times 1)$ vector of US variables and $\mathbf{y_{2t}}$ is a $(n_2 \times 1)$ vector of SOE variables, with $n = n_1 + n_2$ denoting the total number of variables. Similarly, the vector of structural shocks $\epsilon_\mathbf{t}$

[14] Appendix A describes the data and the data sources.

[15] See Kim and Lim (2018), Kim et al. (2017) and Castelnuovo et al. (2022).

is divided into two blocks, $\epsilon_{\mathbf{t}}{}' = \begin{bmatrix} \epsilon'_{1t} & \epsilon'_{2t} \end{bmatrix}$. $\mathbf{A_i}$, for $0 \leq i \leq p$, are $(n \times n)$ matrices of structural parameters, with $\mathbf{A_0}$ invertible. $\mathbf{c}$ is a $(n \times 1)$ vector of constants, p is the lag length, and $\mathbf{T}$ is the sample size. Conditional on past information and initial conditions $\mathbf{y}_0, ..., \mathbf{y}_{1-p}$, the vector $\epsilon_{\mathbf{t}}$ is Gaussian with mean zero and covariance matrix $\mathbb{I}_n$. Following Rubio-Ramirez et al. (2010), we can write the SVAR in compact form:

$$\mathbf{y'_t A_0} = \mathbf{x'_t A_+} + \epsilon'_t \ , \tag{2}$$

where $\mathbf{x'_t} = [\mathbf{y'_{t-1}} \dots \mathbf{y'_{t-p}} \ \ 1]$. $\mathbf{A_0}$ and $\mathbf{A_+} = [\mathbf{A'_1} \dots \mathbf{A'_p} \ \ \mathbf{c'}]$ are matrices of structural parameters.

Post-multiplying Equation (2) by $\mathbf{A_0^{-1}}$, we obtain the reduced-form VAR model:

$$\mathbf{y'_t} = \mathbf{x'_t B} + u'_t \ , \tag{3}$$

where $\mathbf{B} = \mathbf{A_+ A_0^{-1}}, u'_t = \epsilon'_t \mathbf{A_0^{-1}}$ and $\mathbf{E}[u_t u'_t] = \mathbf{\Sigma} = (\mathbf{A_0 A'_0})^{-1}$. $\mathbf{B}$ is the matrix of reduced-form coefficients and $\mathbf{\Sigma}$ is the residual variance-covariance matrix.

3.3 Identification

Our strategy *jointly* identifies the SOE and US monetary shocks by bringing together two distinct approaches: sign restrictions on the parameters governing the monetary policy rules as in Arias et al. (2019) and imposing block exogeneity on the foreign (US) block as in Cushman and Zha (1997).

The first procedure, sign restrictions on policy parameters, offers an approach to identify the systematic component of monetary policy, and thereby monetary policy shocks.[16] The appeal of this method stems from its agnosticism and robustness as it hinges solely on a few qualitative and fairly uncontroversial restrictions on the structural coefficients of the monetary policy rule. This method only achieves set-identification. A caveat inherent in identification schemes based on sign restrictions is the so-called multiple shocks problem: different/distinct shocks may satisfy the same set of sign restrictions (Fry and Pagan, 2011), and the econometrician may end up identifying a shock different from her object of interest.[17] The fact that we combine a set

[16] Leeper et al. (1996) make explicit the link between identifying the systematic component of monetary policy and identifying monetary policy shocks.

[17] Also see Wolf (2020) for an explanation of the 'masquerading shocks' problem in sign-restricted models.

of exclusion restrictions with sign restrictions imposed simultaneously on both US and SOE policy parameters should in principle help to alleviate this partial-identification problem.

The second procedure, block exogeneity, is the hallmark of any SOE model: the SOE is influenced by foreign factors and has no impact on the Rest of the World. In practice, block exogeneity consists of imposing a set of non-recursive zero restrictions. In our setting, block exogeneity complements the minimal set of sign restrictions on policy parameters and augments the information content of our identification scheme. In other words, block exogeneity strengthens the identification of US and SOE monetary shocks through a set of highly plausible zero restrictions.

Our goal in this paper is to assess the robustness of the delayed overshooting and forward discount puzzles. Importantly for our purpose, our identification strategy allows for a simultaneous relationship between the exchange rate and the SOE policy rate and leaves the response of the exchange rate to SOE and US monetary shocks unrestricted at all horizons (Faust and Rogers, 2003; Bjørnland, 2009). We first present details about the way we implement block exogeneity. We then explain the identification of the systematic component of monetary policy in the US and in SOE through sign restrictions on structural parameters.

3.3.1 Block exogeneity

We adapt the methodology of Arias et al. (2019) to incorporate block exogeneity (Cushman and Zha, 1997). Given the partition of $\mathbf{y_t'} = \begin{bmatrix}\mathbf{y_{1t}'} & \mathbf{y_{2t}'}\end{bmatrix}$, the matrix of contemporaneous relationships, $\mathbf{A_0}$, has the following structure:

$$\mathbf{A_0} = \begin{bmatrix} A_{0,11} & A_{0,12} \\ A_{0,21} & A_{0,22} \end{bmatrix},$$

where $A_{0,11}$ is $(n_1 \times n_1)$, $A_{0,12}$ is $(n_1 \times n_2)$, $A_{0,21}$ is $(n_2 \times n_1)$, $A_{0,22}$ is $(n_2 \times n_2)$. To ensure that SOE variables in $\mathbf{y_{2t}}$ do not influence US variables in $\mathbf{y_{1t}}$ contemporaneously, we apply zero-restrictions on the block $A_{0,21}$:

$$\mathbf{A_0} = \begin{bmatrix} A_{0,11} & A_{0,12} \\ 0 & A_{0,22} \end{bmatrix}.$$

We should also prevent SOE variables from influencing US variables in a dynamic fashion. Put differently, in line with Cushman and Zha (1997), we should impose a block of zero-restrictions on each lag matrix $\mathbf{A}_l$, $1 \leq l \leq p$, in Equation (1), so that:

$$\mathbf{A}_l = \begin{bmatrix} A_{11,l} & A_{12,l} \\ A_{21,l} & A_{22,l} \end{bmatrix} = \begin{bmatrix} A_{11,l} & A_{12,l} \\ 0 & A_{22,l} \end{bmatrix},$$

where $A_{11,l}$ is $(n_1 \times n_1)$, $A_{12,l}$ is $(n_1 \times n_2)$, $A_{21,l}$ is $(n_2 \times n_1)$, $A_{22,l}$ is $(n_2 \times n_2)$.

The procedure of Arias et al. (2019) only allows us to impose a maximum of $(n - k)$ zero restrictions per equation, where $k = 1, ..., n$, denotes the order of the k^{th} equation in the system. As a result, we cannot impose $A_{21,l} = \mathbf{0}$. We bypass this issue by formulating a variant of Minnesota priors on the reduced-form VAR, where the priors for the coefficients governing the influence of lagged SOE variables on US variables are concentrated tightly around zero.

Moreover, Arias et al. (2019) specify Natural Conjugate Normal Inverse Wishart ($\mathcal{NIW}$) priors for the reduced-form parameters. Such Natural Conjugate priors are ill-fitted for our purpose: they feature a Kronecker structure for the variance-covariance matrix of the reduced-form parameters, so that variances are proportional to one another. Moreover, the Kronecker structure implies that every equation has the same set of explanatory variables, meaning that if we removed a variable in one equation, that variable would be removed from all equations. Imposing block exogeneity on one equation would then impose it on all equations (Dieppe et al., 2016; Koop et al., 2010). However, the techniques developed by Arias et al. (2018) work for any prior distributions, and hence we employ Independent $\mathcal{NIW}$ priors that allow us to impose block exogeneity structure on a set of equations.

Our specification of the Independent $\mathcal{NIW}$ priors for $\beta = vec(\mathbf{B})$, the vector of reduced-form coefficients, and $\mathbf{\Sigma}$, the residual variance-covariance matrix is:

$$\beta \sim \mathcal{N}(\beta_0, \Omega_0) \tag{4}$$

$$\mathbf{\Sigma} \sim \mathcal{IW}(S_0, \alpha_0).^{18} \tag{5}$$

[18] We set the hyperparameters of the inverse Wishart distribution in a conventional way: $\alpha_0 = n+1$ and $S_0 = \mathbb{I}_n$ (Dieppe et al., 2016).

We center the distribution of every first-order auto-regressive coefficient at 1, and 0 otherwise, as in standard Minnesota priors. The variance-covariance matrix Ω_0 contains the hyper-parameters that control the tightness of the distributions of reduced-form coefficients.[19] The elements of Ω_0 take the following form:

$$\sigma_{c_i}^2 = \sigma_i^2(\lambda_1\lambda_4)^2 \qquad \text{if constant} \tag{6}$$

$$\sigma_{ii}^2 = (\lambda_1/L^{\lambda_3})^2 \qquad \text{if } i = j \tag{7}$$

$$\sigma_{ij}^2 = (\sigma_i/\sigma_j)^2(\lambda_1\lambda_2/L^{\lambda_3})^2 \qquad \text{if } i \neq j \tag{8}$$

$$\sigma_{US_{ij}}^2 = (\sigma_i/\sigma_j)^2(\lambda_1\lambda_2\lambda_5/L^{\lambda_3})^2 \qquad \text{if } i \neq j \text{ and } n_1 < j \leq n \tag{9}$$

where σ_i^2 and σ_j^2 denote the variances of OLS residuals of the auto-regressive models estimated for variables i and j. L is the lag on the coefficient. λ_1 controls the overall tightness of the distribution. λ_4 is the variance parameter of constants. λ_2 controls the tightness of cross-variable distributions. λ_3 is a decaying parameter that controls the speed at which coefficients of variable's own lags (Equation 7) and cross-variable lags (Equation 8), greater than 1 converge to 0. Equations (6), (7) and (8) constitute the standard Minnesota priors. Equation (9) is key to implement block exogeneity: it only applies to the US block and features the additional hyper-parameter λ_5, which controls the tightness of the distributions of coefficients of SOE variables in the US block (Dieppe et al., 2016). Specifically, in Equation (9), $\sigma_{US_{ij}}^2$ are the diagonal elements of Ω_0 corresponding to the domestic coefficients in the equations of foreign variables.[20] This corresponds to the same cross-variables as in Equation (8) but applies only on the domestic variables in the foreign block. This is controlled by the variable range $n_1 < j \leq n$, where n_1 denotes the number of foreign variables.

We set $\lambda_5 = 1e\,{-8}$ to obtain highly informative priors concentrated around zero. We select standard prior variances for the rest of the parameters ($\lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = 1$). To sum up, we implement block exogeneity through a combination of two ingredients: i) we impose exclusion restrictions on the block $A_{0,21}$ in the structural-form matrix of contemporaneous relationships; ii) we apply a special case of Independent $\mathcal{NIW}$ priors

[19] Unlike with Natural Conjugate $\mathcal{NIW}$ priors, Ω_0 is independent of $\mathbf{\Sigma}$.

[20] Ω_0 is expanded in Appendix B.

for the reduced-form VAR, where the reduced-form coefficients follow Minnesota priors with an additional hyper-parameter for implementing block exogeneity on lagged matrices (Dieppe et al., 2016).

3.3.2 Sign restrictions on monetary policy parameters

Our identification scheme builds on that of Arias et al. (2019). Using the techniques developed by Arias et al. (2018), they impose sign and exclusion restrictions on the coefficients of the Federal Reserve's interest-rate rule.[21] As mentioned above, we adapt their methodology to a SOE context in two ways. First, we impose a block-exogenous structure on the SVAR model, meaning that for each SOE, the variables are classified into a US block and a SOE block, where the SOE block has no effect on the US block. Second, for each SOE, we identify simultaneously the interest-rate rule followed by the Federal Reserve and the SOE central bank. Like in Arias et al. (2019), our identification concentrates on the contemporaneous structural parameters. We abstract from the constant term and lags of the structural model in Equation (1) and expand $\mathbf{y_t'}$, $\mathbf{A_0}$ and ϵ_t' as follows:

$$\begin{bmatrix} y_t^* & \pi_t^* & cs_t^* & r_t^* & y_t & \pi_t & r_t & e_t \end{bmatrix} \begin{bmatrix} a_{0,11} & a_{0,12} & a_{0,13} & a_{0,14} & a_{0,15} \cdots a_{0,18} \\ a_{0,21} & a_{0,22} & a_{0,23} & a_{0,24} & a_{0,25} \cdots a_{0,28} \\ a_{0,31} & a_{0,32} & a_{0,33} & a_{0,34} & a_{0,35} \cdots a_{0,38} \\ a_{0,41} & a_{0,42} & a_{0,43} & a_{0,44} & a_{0,45} \cdots a_{0,48} \\ 0 & 0 & 0 & 0 & a_{0,55} \cdots a_{0,58} \\ \vdots & \ddots & \ddots & \ddots & \ddots \quad \vdots \\ 0 & 0 & 0 & 0 & a_{0,85} \cdots a_{0,88} \end{bmatrix} = \ldots + \begin{bmatrix} \epsilon_{1,t} \\ \epsilon_{2,t} \\ \epsilon_{3,t} \\ \epsilon_{4,t} \\ \epsilon_{5,t} \\ \epsilon_{6,t} \\ \epsilon_{7,t} \\ \epsilon_{8,t} \end{bmatrix}'$$

We identify the first and the fifth shock in the SVAR model as the US and the SOE monetary policy shocks, respectively:

$$y_t^* a_{0,11} + \pi_t^* a_{0,21} + cs_t^* a_{0,31} + r_t^* a_{0,41} = \epsilon_{1,t} \tag{10}$$

$$y_t^* a_{0,15} + \pi_t^* a_{0,25} + cs_t^* a_{0,35} + r_t^* a_{0,45} + y_t a_{0,55} + \pi_t a_{0,65} + r_t a_{0,75} + e_t a_{0,85} = \epsilon_{5,t} \tag{11}$$

The systematic component of US monetary policy: The US monetary policy rule, from Equation (10), is given by:

[21] Also see Binning (2013).

$$r_t^* = -a_{0,41}^{-1}a_{0,11}y_t^* - a_{0,41}^{-1}a_{0,21}\pi_t^* - a_{0,41}^{-1}a_{0,31}cs_t^* + a_{0,41}^{-1}\epsilon_{1,t} \tag{12}$$

where $-a_{0,41}^{-1}a_{0,11} = \psi_{y^*}$, $-a_{0,41}^{-1}a_{0,21} = \psi_{\pi^*}$, $-a_{0,41}^{-1}a_{0,31} = \psi_{cs^*}$ and $a_{0,41}^{-1} = \sigma^*$.

To characterize the systematic component of US monetary policy, we impose the following two restrictions.

Restriction 1. The contemporaneous response of the US policy rate to US output and US inflation is positive: $\psi_{y^*} > 0$ and $\psi_{\pi^*} > 0$.

Restriction 2. The contemporaneous reaction of the US policy rate to the Baa corporate credit spread is negative: $\psi_{cs^*} < 0$.

Restriction 1 is motivated by Taylor (1993) and a large DSGE literature.[22] Restriction 2 is consistent with the restrictions imposed by Baumeister and Hamilton (2018) and Rüth and Van der Veken (2023) as well as the SVAR evidence provided by Caldara and Herbst (2019).[23] Combining Restrictions 1 and 2, we obtain the following characterization of US monetary policy:

$$r_t^* = \underbrace{-a_{0,41}^{-1}a_{0,11}}_{\psi_{y^*}>0} y_t^* \underbrace{-a_{0,41}^{-1}a_{0,21}}_{\psi_{\pi^*}>0} \pi_t^* \underbrace{-a_{0,41}^{-1}a_{0,31}}_{\psi_{cs^*}<0} cs_t^* \underbrace{+a_{0,41}^{-1}}_{\sigma^*} \epsilon_{1,t} \tag{13}$$

The systematic component of monetary policy in SOEs: The SOE monetary policy rule, from Equation (11), is given by:

$$\begin{aligned} r_t = &-a_{0,75}^{-1}a_{0,15}y_t^* - a_{0,75}^{-1}a_{0,25}\pi_t^* - a_{0,75}^{-1}a_{0,35}cs_t^* - a_{0,75}^{-1}a_{0,45}r_t^* \\ &-a_{0,75}^{-1}a_{0,55}y_t - a_{0,75}^{-1}a_{0,65}\pi_t - a_{0,75}^{-1}a_{0,85}e_t + a_{0,75}^{-1}\epsilon_{5,t} \end{aligned} \tag{14}$$

where $-a_{0,75}^{-1}a_{0,55} = \psi_y$, $-a_{0,75}^{-1}a_{0,65} = \psi_\pi$, $-a_{0,75}^{-1}a_{0,85} = \psi_e$ and $a_{0,75}^{-1} = \sigma$.

To identify the systematic component of SOE monetary policy, we impose the following two restrictions.

[22] Regarding the timing assumption implied by Restriction 1, where the policy rate, output and inflation interact simultaneously, we follow the argument of Arias et al. (2019): monetary authorities crunch a large battery of real-time indicators to nowcast the current state of the economy.

[23] Curdia and Woodford (2010, 2016) present DSGE-based analysis justifying a negative systematic response of monetary policy to a worsening of credit conditions.

Restriction 3. The contemporaneous reaction of the SOE policy rate to domestic output and inflation is positive: $\psi_y > 0$ and $\psi_\pi > 0$.

Restriction 4: The contemporaneous reaction of the SOE policy rate to the real bilateral SOE/US exchange rate is positive: $\psi_e > 0$.

Restrictions 3 and 4 leave the reaction of the SOE central bank to foreign variables unrestricted as in Cushman and Zha (1997). Restriction 4 means that the SOE central bank usually leans against the real SOE/US exchange rate, cutting its policy rate in response to an appreciation of the domestic currency, and increasing it in response to a depreciation. Restriction 4 is consistent with findings based on SVARs (Bjørnland, 2009; Bjørnland and Halvorsen, 2014) and Taylor-type rules embedded in DSGE models (Lubik and Schorfheide, 2007; Kam et al., 2009; Justiniano and Preston, 2010).[24] Taken together, Restrictions 3 and 4 imply that the SOE central bank follows a Taylor-type rule in line with Taylor (2001):

$$\begin{aligned} r_t = & \underbrace{-a_{0,75}^{-1} a_{0,15}}_{unrestricted} y_t^* \underbrace{-a_{0,75}^{-1} a_{0,25}}_{unrestricted} \pi_t^* \underbrace{-a_{0,75}^{-1} a_{0,35}}_{unrestricted} cs_t^* \underbrace{-a_{0,75}^{-1} a_{0,45}}_{unrestricted} r_t^* \\ & \underbrace{-a_{0,75}^{-1} a_{0,55}}_{\psi_y > 0} y_t \underbrace{-a_{0,75}^{-1} a_{0,65}}_{\psi_\pi > 0} \pi_t \underbrace{-a_{0,75}^{-1} a_{0,85}}_{\psi_e > 0} e_t \underbrace{+a_{0,75}^{-1}}_{\sigma} \epsilon_{5,t} \end{aligned} \tag{15}$$

Equation (15) shows that the coefficients on the US variables are left unrestricted for the SOE monetary policy rule, which captures the spillover effects of the US economy on the SOE policy rate. On the other hand, Equation (13) shows that SOE variables have no impact on the US policy rate, which captures the closed economy feature of the US.

[24] Calvo and Reinhart (2002), Reinhart and Rogoff (2004), Obstfeld (2013) and Ilzetzki et al. (2019) find that many central banks react to the dollar exchange rate. Gopinath et al. (2020) and Gourinchas et al. (2019) document the central role of the dollar in international monetary and financial system. Egorov and Mukhin (2023) show that, when prices are invoiced and sticky in dollars, it can be desirable for non-US central banks to stabilize the dollar exchange rate.

4 Results

This section presents our main results.[25] First, we discuss the impulse response functions (IRFs) to a tightening of monetary policy in the SOE and in the US. Second, we investigate the importance of Restriction 4 (*leaning against the RER*) and Restriction 2 (*leaning against credit frictions*) for the identification of, respectively, SOE and US monetary shocks.

4.1 IRFs to a SOE contractionary monetary shock

Figure 1 plots the IRFs of domestic variables for the six SOEs to one standard deviation contractionary domestic monetary policy shock. The blue solid lines depict the point-wise posterior median responses while the blue shaded bands correspond to the 68% equal-tailed point-wise posterior probability bands. For all SOEs, we observe that the policy rate jumps on impact within a range of 15 to 40 basis points, and reaches its peak over the first three quarters. Except for Canada, the policy rate increase remains significant for several quarters. For all SOEs, the posterior median response of output displays an instantaneous contraction and stays below trend for several years after the shock. Except for the UK, the decline in output remains significant for several quarters after the shock. For the UK, the response of output is insignificant, although the bulk of the 68% probability bands lies in the negative region, suggesting that output contracts at least in the short run. For Canada and New Zealand, the response of output stays significantly below trend throughout the entire five-year horizon. Hence, we do not observe any evidence of the output puzzle (Uhlig, 2005).

For all SOEs, inflation falls instantaneously and reverts back to its steady state quickly. The negative impact response of inflation is either significant or borderline significant across all SOEs.

Turning to our variable of interest, in all SOEs, the RER appreciates sharply and significantly on impact: we do not find any evidence of the exchange rate puzzle (an immediate depreciation after the tightening of domestic monetary policy). For Canada, New Zealand and the UK, the instantaneous appreciation is immediately

[25] We set the lag order $p = 2$. All results are based on 1 million draws from the posterior distributions.

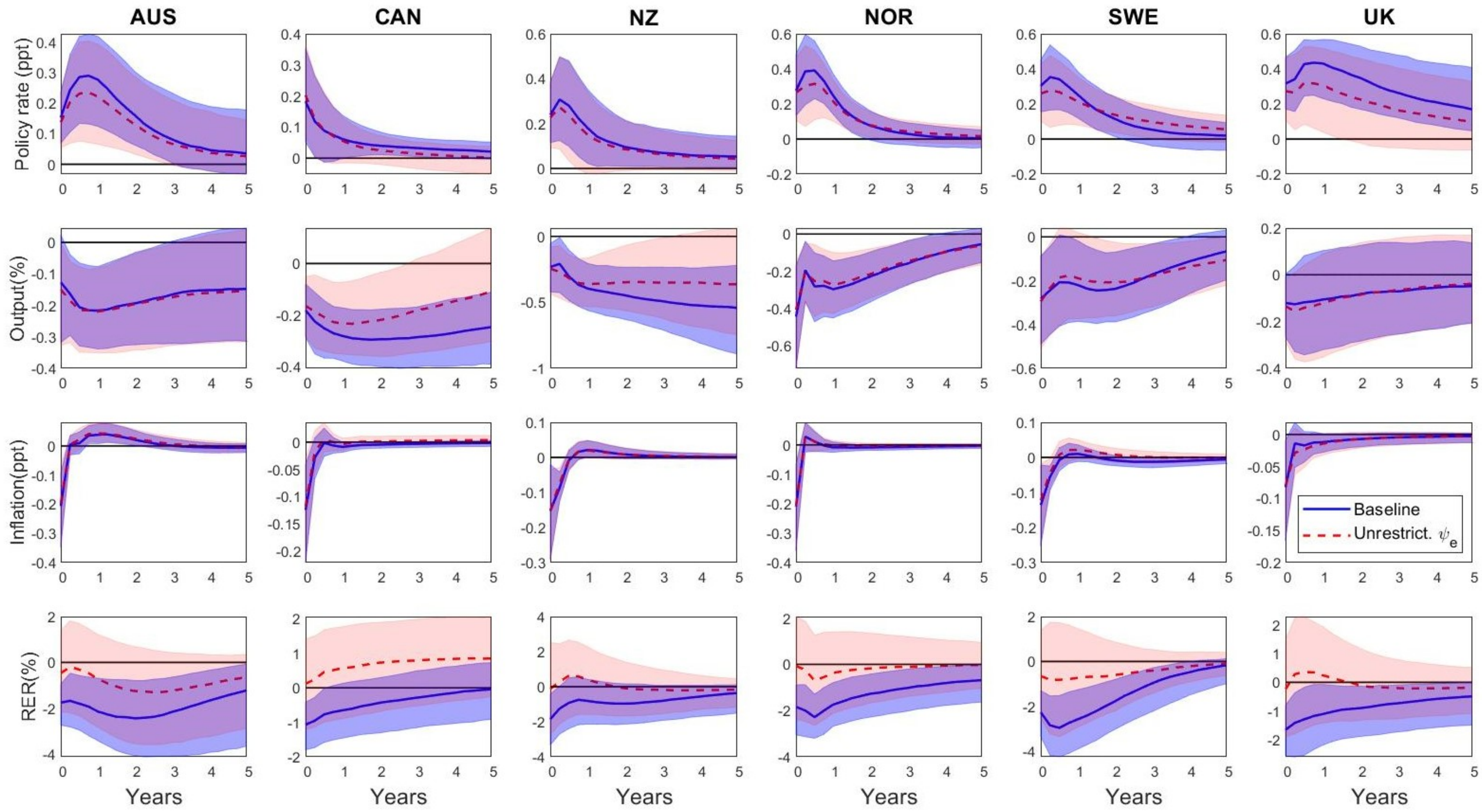

Figure 1: IRFs of SOE variables to a one standard deviation SOE contractionary monetary shock. *Note:* The blue solid lines are the point-wise posterior median responses under block exogeneity and Restrictions 1 to 4. The red dashed lines are the posterior median responses, after relaxing Restriction 4 ($\psi_e > 0$). The shaded areas represent the respective 68% equal-tailed posterior probability bands.

followed by a gradual and persistent depreciation. For Norway and Sweden, the RER appreciation reaches its peak two quarters after the monetary tightening. On the other hand, the AUD/US RER displays a hump-shaped response, with the peak appreciation occurring two years after the shock. Thus, except for Australia, we observe little evidence of the delayed overshooting puzzle (a gradual and persistent appreciation that reaches its peak roughly two years after the shock). The IRFs of the bilateral SOE/US RER to a SOE monetary shock appear broadly consistent with the Dornbusch overshooting hypothesis. Our findings are in line with the SOE-SVAR studies by Bjørnland (2009), Bjørnland and Halvorsen (2014), Kim and Lim (2018) and Terrell et al. (2023). Our findings reinforce the view that the exchange rate puzzle and the delayed overshooting puzzle may be artifacts caused by dubious identifying restrictions that hinder the simultaneous interactions between monetary policy and the exchange rate.

4.2 IRFs to a US contractionary monetary shock

Figure 2 shows the responses of SOE variables to a one standard deviation US contractionary monetary shock (blue solid line). The responses of US variables are reported in Appendix E, which are almost identical across the six SOE-SVARs due to the block-exogeneous structure of the model.

Looking at the RER responses across the six SOEs (blue solid lines in Figure 2), we observe that the US dollar appreciates significantly on impact in response to the US monetary tightening (i.e. no exchange rate puzzle). Moreover, the US dollar reaches its peak appreciation within the first quarter after the shock, and gradually depreciates afterwards (i.e. no delayed overshooting). Similar findings are reported in the US SVAR monetary policy literature (Kim et al., 2017; Rüth, 2020; Castelnuovo et al., 2022).[26] A distinguishing feature of our study, however, is the inclusion of the ZLB period.[27]

Looking at the response of SOE output to US monetary tightening, we observe a

[26] On the other hand, Eichenbaum and Evans (1995), Faust and Rogers (2003) and Rüth and Van der Veken (2023) find evidence of delayed overshooting puzzle while Scholl and Uhlig (2008) rule this puzzle out by construction.

[27] Our results are robust to using a shorter sample period from 1992:Q1 to 2008:Q3 that excludes the zero-lower-bound (ZLB) period. See the section on robustness checks below.

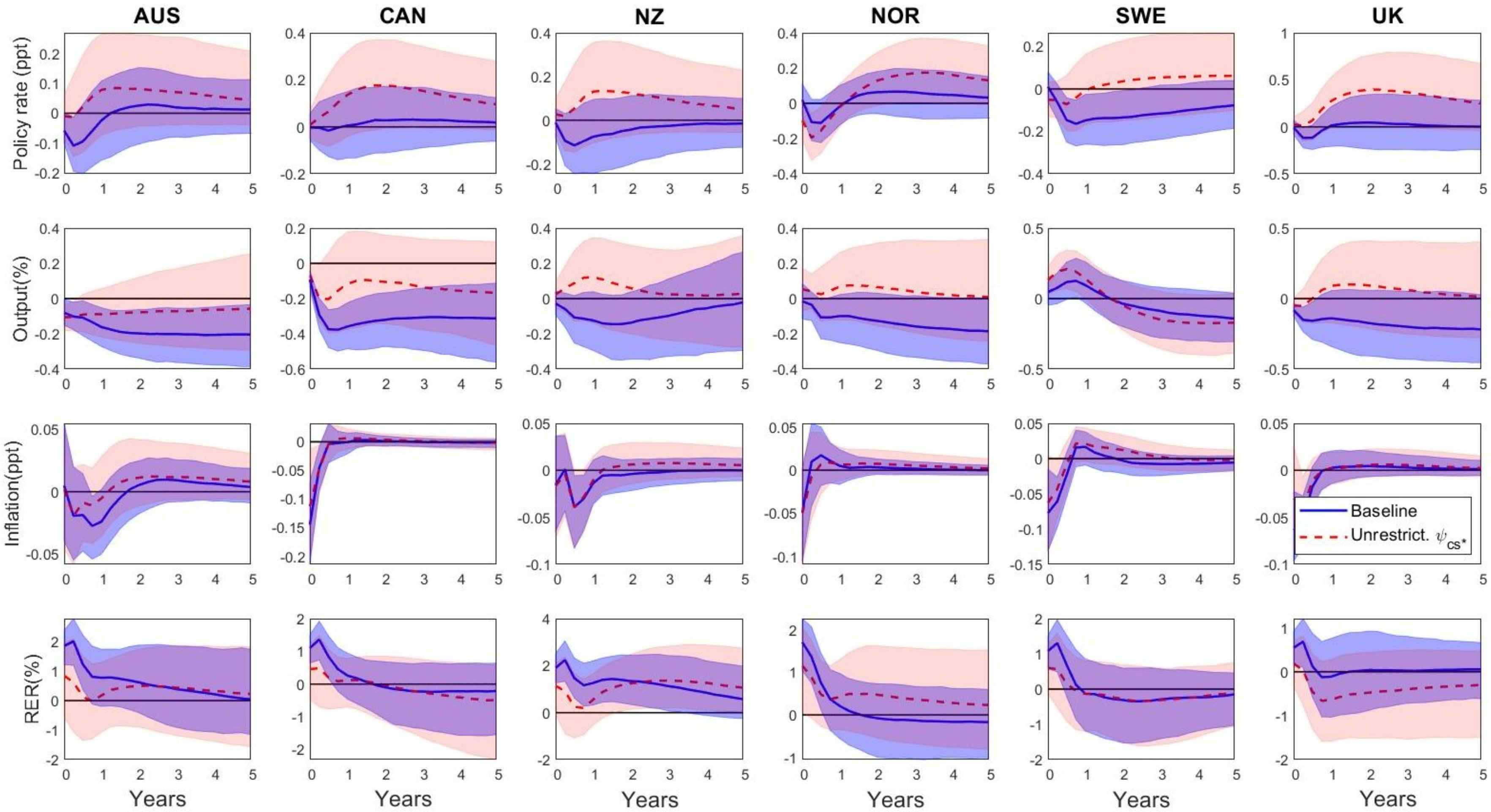

Figure 2: IRFs of SOE variables to a one standard deviation US contractionary monetary shock. *Note:* The blue solid lines are the point-wise posterior median responses under block exogeneity and Restrictions 1 to 4. The red dashed lines are the posterior median responses, after relaxing Restriction 2 ($\psi_{cs^*} < 0$). The shaded areas represent the respective 68% equal-tailed posterior probability bands.

protracted contraction for Australia, Canada, the UK and to a lesser extent Norway. The decline in output is particularly strong in Canada and Australia. The response of output in New Zealand and Sweden is more muted. Following the US monetary tightening, inflation falls in all SOEs. Except for Canada, SOE central banks lower their policy rate slightly to mitigate the negative spillovers of the US monetary tightening. The monetary easing is most visible in Sweden and Australia.[28]

4.3 Relaxing Restriction 4: Importance of the SOE monetary response to the RER

We now perform a sensitivity analysis to shed light on the importance of Restriction 4 ($\psi_e > 0$) in our identification scheme. We re-estimate the six SOE-SVAR models without Restriction 4 while keeping everything else unchanged. As a result, the contemporaneous response of the SOE policy rate to the real exchange rate is now left unrestricted. The systematic component of SOE monetary policy takes the following form:

$$\begin{aligned} r_t = & \underbrace{-a_{0,75}^{-1}a_{0,15}}_{unrestricted} y_t^* \underbrace{-a_{0,75}^{-1}a_{0,25}}_{unrestricted} \pi_t^* \underbrace{-a_{0,75}^{-1}a_{0,35}}_{unrestricted} cs_t^* \underbrace{-a_{0,75}^{-1}a_{0,45}}_{unrestricted} r_t^* \\ & \underbrace{-a_{0,75}^{-1}a_{0,55}}_{\psi_y>0} y_t \underbrace{-a_{0,75}^{-1}a_{0,65}}_{\psi_\pi>0} \pi_t \underbrace{-a_{0,75}^{-1}a_{0,85}}_{unrestricted} e_t \underbrace{+a_{0,75}^{-1}}_{\sigma} \epsilon_{5,t} \end{aligned} \tag{16}$$

Figure 1 compares the IRFs (blue solid lines) of SOE variables to an SOE monetary shock with Restriction 4 imposed as in Equation (15) to the IRFs (red dashed lines) without Restriction 4, as in Equation (16). The shaded regions are the respective 68% equal-tailed point-wise posterior probability bands. Figure 1 shows that relaxing Restriction 4 has virtually no effect on the IRFs of the policy rate, output and inflation. Instead, relaxing Restriction 4 greatly alters the IRFs of the RER. Most remarkable is the fact that, for all SOEs, the effects of monetary shocks on the RER are now insignificant, even in the short run. This finding clearly goes against the consensus view that monetary policy plays a role in accounting for the elevated short-run volatility typically observed in exchange rates. Strikingly, we also notice the re-emergence of several exchange rate puzzles when the monetary policy response to the exchange rate

[28]Similar findings are reported by Camara et al. (2024) for 10 advanced economies (AEs) using a panel VAR.

is left unrestricted. For Canada, the posterior median response clearly indicates that the exchange rate depreciates instead of appreciating, consistent with the exchange rate puzzle. Taken together, these findings contradict the Dornbusch (1976) overshooting hypothesis according to which a surprise tightening of monetary policy at home causes an instantaneous appreciation of the domestic currency, immediately followed by a gradual depreciation back to the steady state. Considering the puzzling evidence obtained when relaxing Restriction 4 and the assorted motivations for imposing Restriction 4 found in various strands of the literature (Taylor, 2001; Bjørnland, 2009; Lubik and Schorfheide, 2007), we conclude that imposing Restriction 4 contributes usefully to a proper identification of the systematic behavior of SOE central banks.[29]

4.4 Relaxing Restriction 2: The RER implications of the Fed's leaning against the credit spread

We now emphasize the importance of restricting the coefficient on Baa spread in the US monetary policy rule. Specifically, we assess the implications of Restriction 2 ($\psi_{cs^*} < 0$) for the RER. Unlike previous studies that focus on the importance of restricting the US Baa spread (Baumeister and Hamilton, 2018) and the excess bond premium (Rüth and Van der Veken, 2023) for the US only, we focus on the importance of restricting the US Baa spread for the SOEs.[30] We re-estimate the six SOE-SVARs without Restriction 2, keeping everything else unchanged. The contemporaneous response of the US policy rate to the credit spread is left unrestricted:

$$r_t^* = \underbrace{-a_{0,41}^{-1}a_{0,11}}_{\psi_{y^*}>0} y_t^* \underbrace{-a_{0,41}^{-1}a_{0,21}}_{\psi_{\pi^*}>0} \pi_t^* \underbrace{-a_{0,41}^{-1}a_{0,31}}_{unrestricted} cs_t^* \underbrace{+a_{0,41}^{-1}}_{\sigma^*} \epsilon_{1,t} \tag{17}$$

Figure 2 compares the IRFs of SOE variables to a US contractionary monetary policy shock with (blue solid lines) and without (red dashed lines) Restriction 2. We see that the main effect of relaxing Restriction 2 results in insignificant responses of RER for all the six SOEs (red dashed lines). This finding, which suggests that US monetary policy shocks have no material effects on exchange rates even in the short run, goes against the conventional wisdom on the contribution of monetary disturbances to exchange rate

[29]The IRFs to a US monetary shock with and without Restriction 4 are reported in Appendix F. They show that Restriction 4 is irrelevant for identifying US monetary shocks.

[30]The responses of US variables with/without $\psi_{cs^*} < 0$ are provided in Appendix G.

volatility. Moreover, the fact that US monetary policy is perceived as the main driver of the global financial cycle (Rey, 2015; Miranda-Agrippino and Rey, 2020), makes this finding look somewhat implausible. We generally observe that relaxing Restriction 2 shifts the posterior probability bands towards negative territory, meaning that the indicative evidence of a depreciation of the US dollar (instead of an appreciation, as we would have expected) builds up. In other words, relaxing Restriction 2 makes the exchange rate puzzle more visible (see in particular the IRFs of the GBP/USD, CAD/USD and SEK/USD). Overall, these dubious phenomena emphasize the added value of imposing Restriction 2 in order to obtain plausible responses of the RER and other SOE variables through the correct identification of the systematic component of US monetary shocks.[31] Rüth and Van der Veken (2023) show that imposing a similar restriction on excess bond premium for the US results in (i) exchange rate overshooting with less delay and (ii) little deviations from UIP.

5 Revisiting UIP in the context of exchange rate overshooting

The uncovered interest rate parity (UIP) condition is one of the key building blocks underpinning Dornbusch's overshooting hypothesis (Rüth, 2020), and more generally the New Keynesian DSGE models (Lane, 2001). UIP postulates that a rise in the interest rate differential between the domestic and foreign policy rates has to be quantitatively offset by an expected fall in the value of the home currency, i.e. a depreciation of the nominal exchange rate one period ahead.

In this section, we proceed in two stages. First, we understand how excess interest rate returns, a common metric to validate UIP, behave in our set of SOEs conditional on domestic and US monetary policy shocks. Then, we delve deeper into the link between UIP and exchange rate overshooting, highlighting key results obtained by estimating the SOE-SVARs on shorter samples. A delay in exchange rate overshooting is not necessarily a violation of UIP. The *joint* identification of the SOE and US monetary policy rules helps us establish that the delay in overshooting may in fact be necessary for the UIP to hold.

[31] The full set of IRFs to a US monetary shock with and without Restriction 2 are in Appendix G.

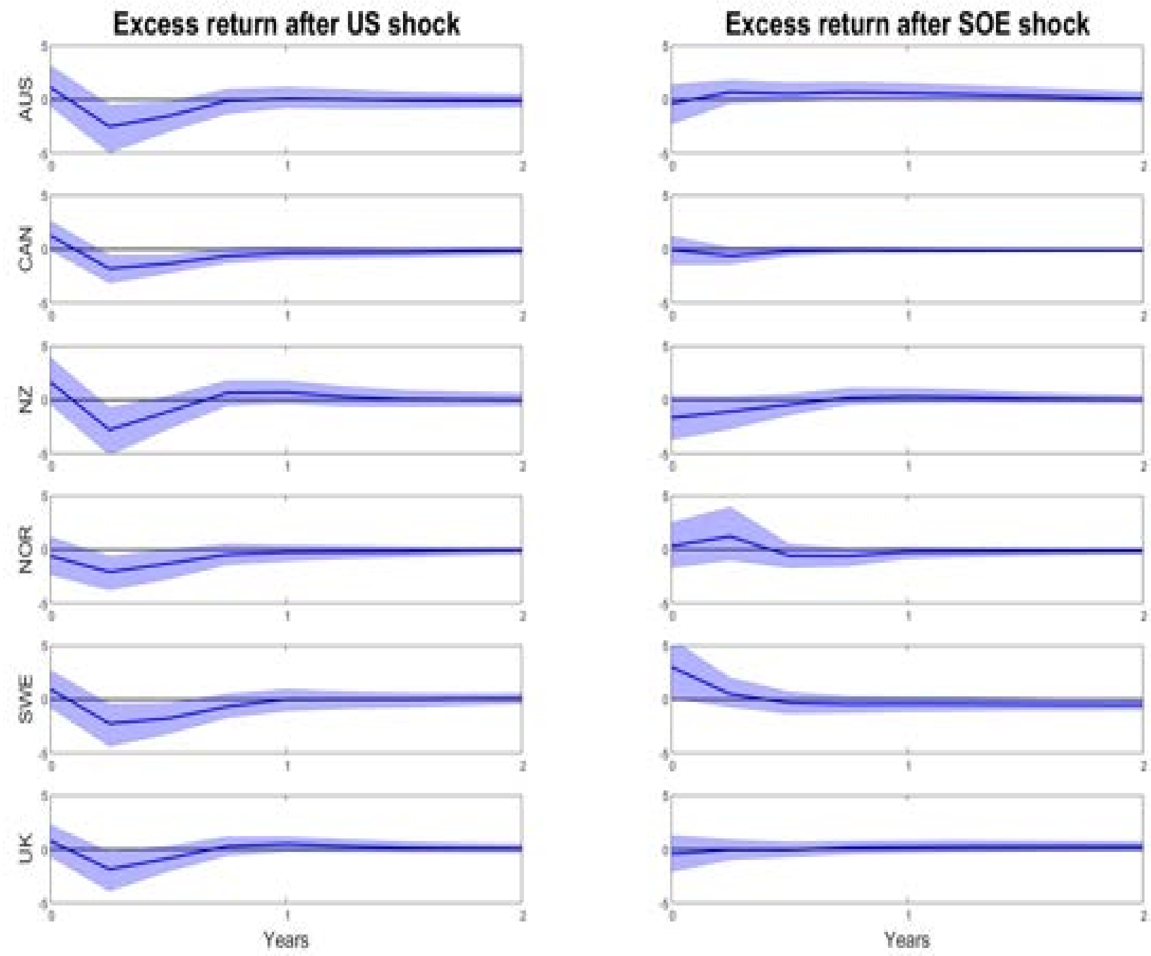

Figure 3: Deviations from UIP conditional on US (left) and SOE (right) monetary policy shocks. *Note:* The solid lines represent the point-wise posterior median estimates of excess returns. The shaded areas are the 68% posterior probability intervals.

5.1 Deviations from UIP triggered by monetary policy shocks

Following Eichenbaum and Evans (1995) and Bjørnland (2009), we compute the excess returns measured in USD and SOE currency as Λ_t^{US} and Λ_t^{SOE}, respectively:

$$\Lambda_t^{US} = r_t^* - r_t + 4 \times (\mathbb{E}_t\{s_{t+1}\} - s_t) \tag{18}$$

$$\Lambda_t^{SOE} = r_t - r_t^* - 4 \times (\mathbb{E}_t\{s_{t+1}\} - s_t) \tag{19}$$

where s_t is the nominal exchange rate and a fall in s_t is an appreciation of the SOE currency.[32] $\mathbb{E}$ denotes the conditional expectations operator.

[32]It is straightforward to compute the excess returns directly from RER(e) impulse and US and SOE inflation and policy rate impulses. For example, for the US, $\Lambda_t^{US} = r_t^* - r_t + 4*(e_{t+1} - e_t) - \pi_{t+1}^* + \pi_{t+1}$.

According to UIP, excess returns should be zero at all horizons:

$$\mathbb{E}_t\{\Lambda_{t+j}\} = 0 \qquad \text{for all } j \geqslant 0$$

for simplicity, we abstract from the superscript on Λ.

Figure 3 reports the point-wise posterior median estimates of excess returns conditional on US (left panel) and SOE (right panel) monetary policy shocks, along with the 68% posterior probability intervals. We do not find any evidence of UIP violations in response to SOE monetary shocks: excess returns triggered by SOE disturbances are quantitatively modest and insignificant at all horizons. Deviations from UIP generated by US policy shocks are also moderate and largely insignificant. Thus, overall, the conditional dynamics of exchange rates following US and SOE monetary disturbances appear to be largely consistent with UIP. Our results are in line with Bjørnland (2009), who reports exchange rate movements broadly consistent with UIP conditional on SOE monetary disturbances, and with Rüth (2020) who finds little evidence of UIP violations conditional on US monetary shocks. Instead, Eichenbaum and Evans (1995), Faust and Rogers (2003) and Scholl and Uhlig (2008) report evidence of the forward discount puzzle, i.e. large and significant deviations from UIP, conditional on US monetary shocks.

5.2 Monetary policy, UIP and delayed overshooting: what a split sample brings to light

The previous section provided empirical support in favour of UIP in our set of SOEs, conditional on SOE and US shocks. Here, we focus on the insights that emerge from an exercise splitting the sample into two: (i) 1992:Q1-2008:Q3 that excludes episodes of unconventional monetary policy and the binding zero lower bound and (ii) 2008:Q4-2019:Q4 that includes these episodes.[33] Specifically, we examine the questions posed by Scholl and Uhlig (2008): (i) is delayed overshooting consistently observed in response to monetary policy shocks? (ii) is the forward discount puzzle a 'twin appearance' of delayed overshooting? We further investigate: what mechanism drives

[33] We set the lag order $p = 1$ for the two small samples. For the period 1992:Q1-2008:Q3, we do not need any shadow rates to measure the stance of SOE and US monetary policies, while we use shadow rates for the period 2008:Q4-2019:Q4. US shadow rate falls to around -4%, which is synonymous to "powerful non-conventional stimuli" (Lhuissier et al., 2020).

the exchange rate dynamics when UIP holds even with delayed overshooting? Since the exchange rate dynamics triggered by SOE monetary policy shocks are similar across the samples, we consider a contractionary monetary policy shock in the US for both samples.[34] Figures 4 and 5 report the distinct dynamic responses triggered by a surprise monetary contraction during the periods of US conventional (92:Q1-08:Q3) and unconventional (08:Q4-19:Q4) monetary policy, respectively.[35,36]

5.2.1 Endogenous SOE monetary policy responses and *mildly* delayed overshooting

The spillover effects of US monetary shocks on SOE policy rates have evolved over time. For the first period, i.e. US conventional monetary stance, we find a narrow gap between US and SOE policy rates to US monetary surprises (gap between IRFs of rows 1 and 2 in Figure 4, as shown in row 3). As a result, the magnitude of the interest rate differentials, $r^* - r$, is *small*. It is pertinent to note that the median responses of the interest rate differentials mostly fall in *negative* territory for the six SOEs, however, as discussed below, a *positive* interest rate differential is essential for the exchange rate to depreciate after an impact appreciation (Scholl and Uhlig, 2008). The narrow interest rate differential is the reason we find evidence of *mildly* delayed overshooting for most of the economies in response to US shocks (Figure 4, row 5). During this period, we find that SOE central banks mimic the Fed policy rate in an effort to stabilize the exchange rates. On the other hand, for the period of US unconventional monetary policy, interest rate differentials become *positive* due to the widening policy rate gaps (Figure 5, row 3). Thus, we see an immediate appreciation of USD followed by a mean reversion (Figure 5, row 5). During this period, SOE central banks try to mitigate the contractionary impact of the US Fed by responding with monetary easing. Significantly large responses of SOE policy rates may be attributed to the un-

[34]IRFs in response to SOE monetary shocks for the two sub-samples are provided in Appendix H.

[35]Given that we do not have price levels in our model, we approximate the IFRs of nominal exchange rate (s) using RER, US and SOE inflation impulses, such that $s_t \approx e_t - \pi^* + \pi$. Ideally, we should cumulate the impulse responses of Δs_t, where $\Delta s_t = \Delta e_t - \frac{1}{4}\pi_t^* + \frac{1}{4}\pi_t$. However, we do not find accurate IRFs of s by cumulating Δs_t. Virtually identical responses of s and RER are not uncommon in the literature (see, e.g., Eichenbaum and Evans (1995) and Scholl and Uhlig (2008)).

[36]Identical responses of r^* in first rows of Figures 4 and 5 are due to the block exogeneity structure.

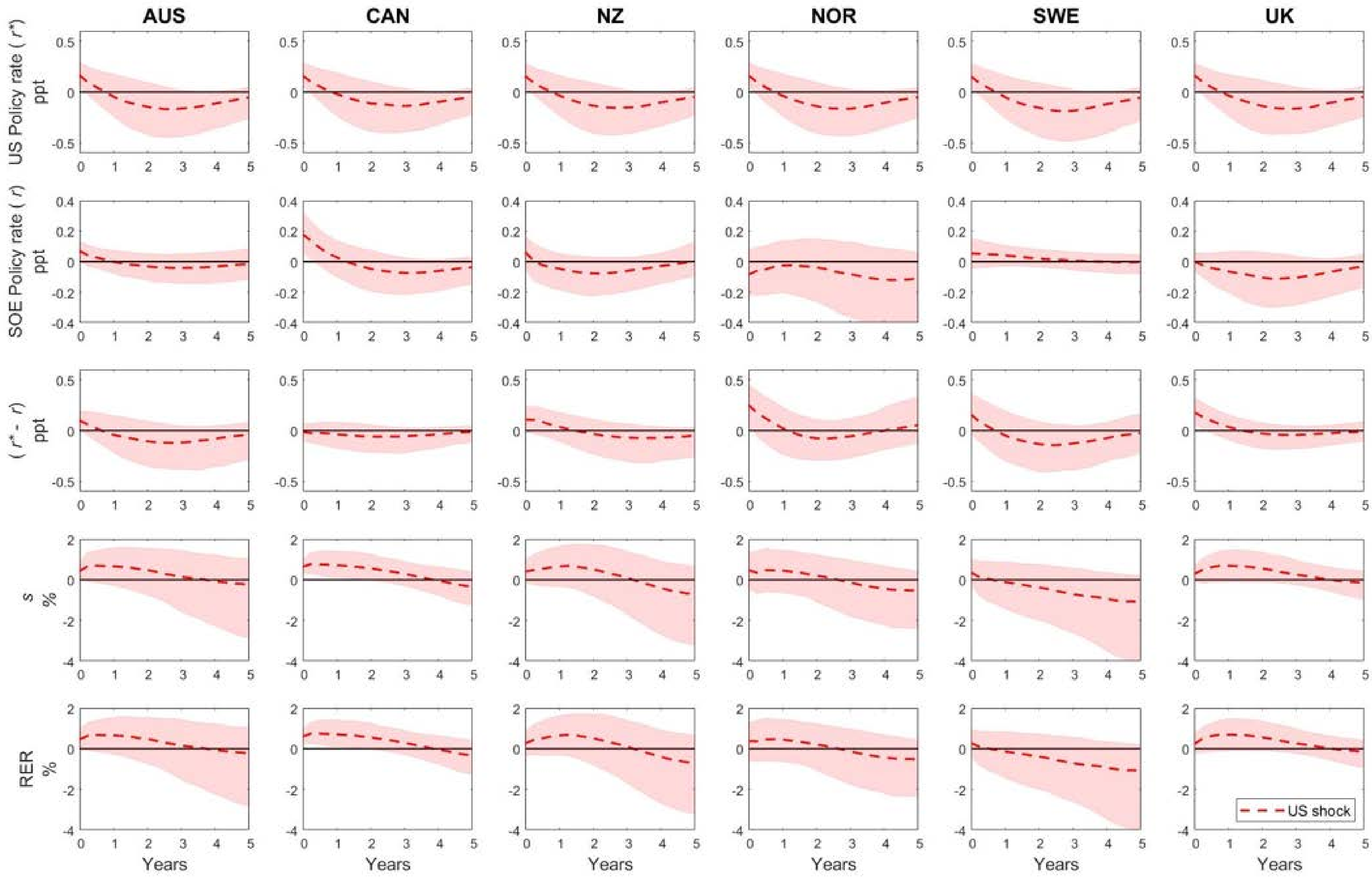

Figure 4: IRFs of US and SOE policy rates and nominal (s) and real exchange rate (RER) to a one standard deviation US contractionary monetary shocks identified using block exogeneity and Restrictions 1 to 4 for sample period 1992:Q1 to 2008:Q3. Row 3 shows the interest rate differentials between the US (r^*) and SOE (r) policy rates to US shocks. *Note:* The red dashed lines are the point-wise posterior median responses from US contractionary monetary shocks. The shaded areas represent the respective 68% equal-tailed posterior probability bands.

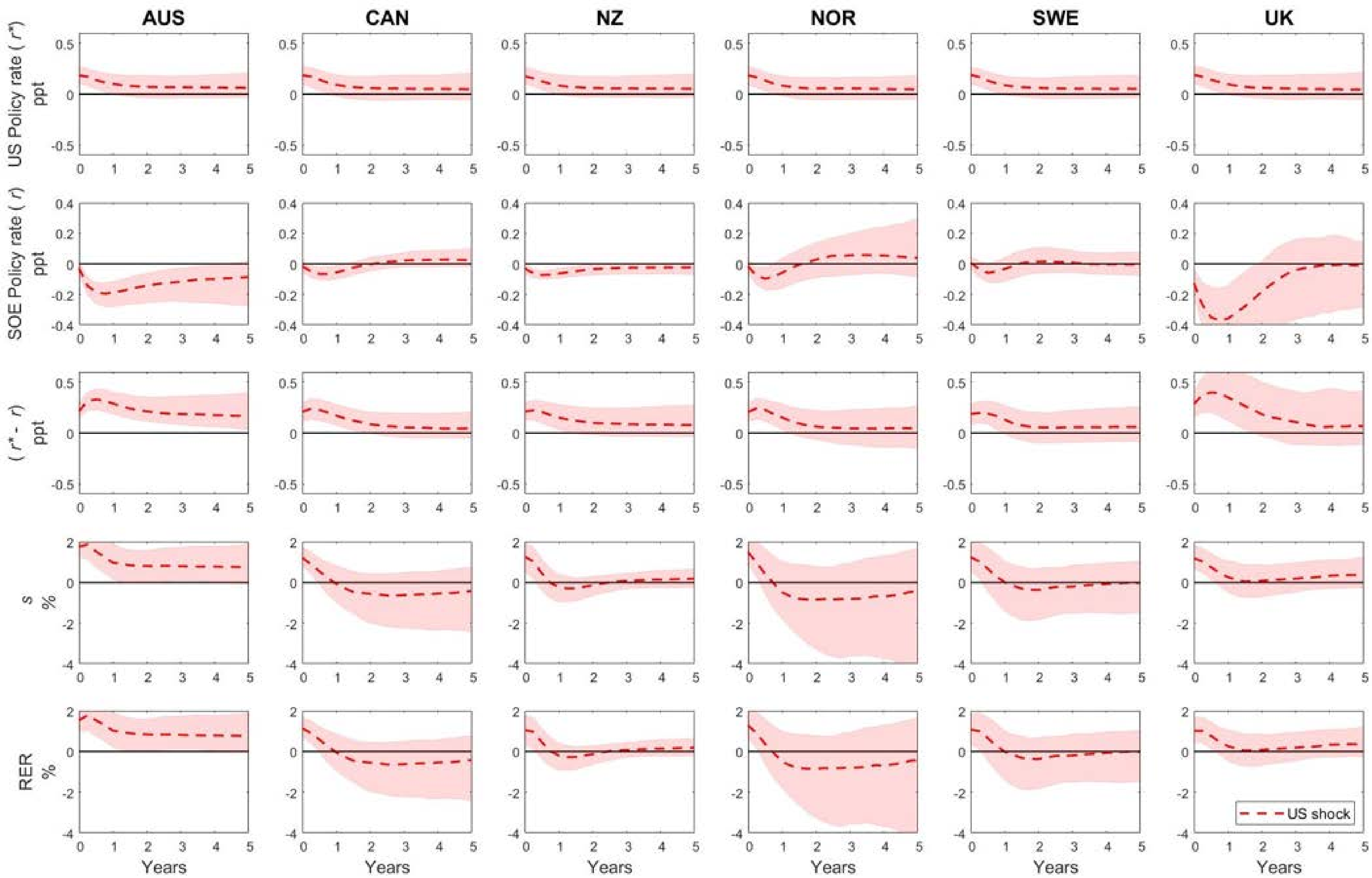

Figure 5: IRFs of US and SOE policy rates and nominal (s) and real exchange rate (RER) to a one standard deviation US contractionary monetary shocks identified using block exogeneity and Restrictions 1 to 4 for sample period 2008:Q4 to 2019:Q4. Row 3 shows the interest rate differentials between the US (r^*) and SOE (r) policy rates to US shocks. *Note:* The red dashed lines are the point-wise posterior median responses from US contractionary monetary shocks. The shaded areas represent the respective 68% equal-tailed posterior probability bands.

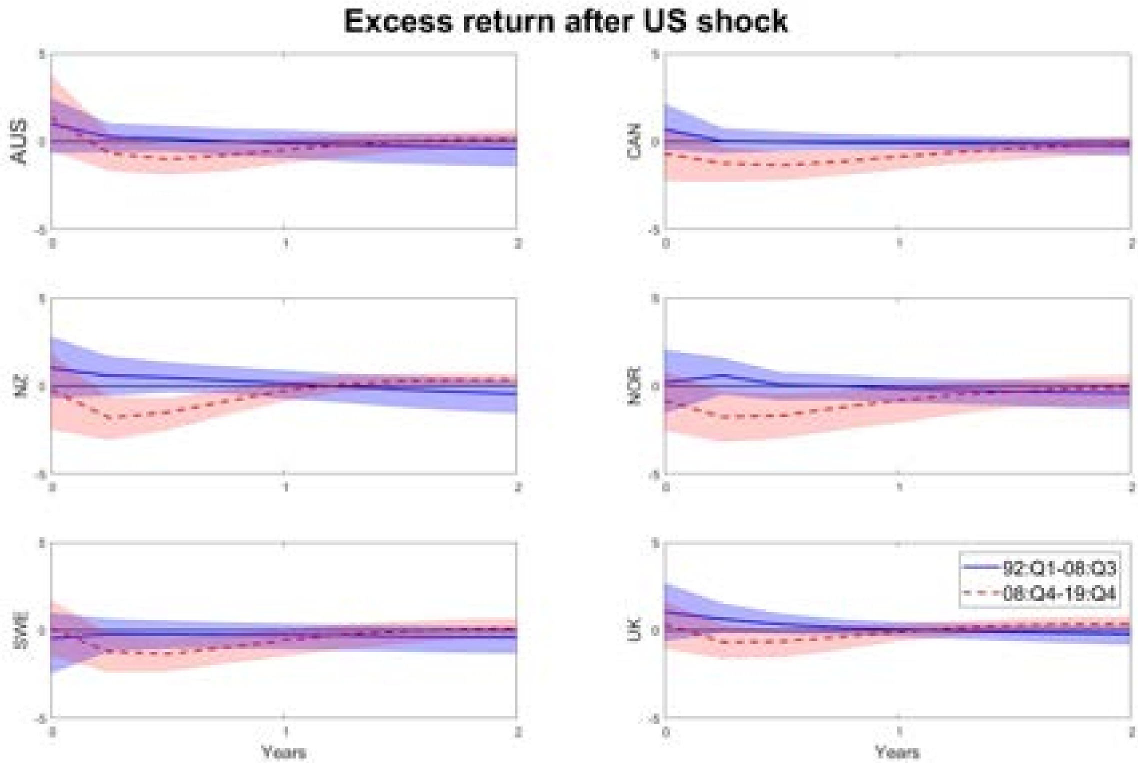

Figure 6: Deviations from UIP conditional on US monetary policy shocks. *Note:* The blue solid (red dashed) lines depict the point-wise posterior median responses from 92:Q1-08:Q3 (08:Q4-19:Q4). The shaded areas are the 68% posterior probability intervals.

conventional policy stance adopted by the US.[37,38] Hence, we argue that SOE central banks may have responded differently to US conventional vs unconventional monetary surprises resulting in an evolution of US-SOE interest rate differentials. What follows is a discussion in detail how the exchange rate dynamics adjusts to these interest rate differentials through UIP.

5.2.2 UIP holds with/without delayed overshooting

Figure 6 reports that UIP conditional on US shocks largely holds for the two samples, except for some borderline cases.[39] However, we do find evidence of *mildly* delayed

[37] Neely (2015) reports substantial international spillovers through long-term foreign bond yields when US adopted unconventional measures during the GFC.

[38] Curcuru et al. (2023) state that the evidence of spillover effects of unconventional monetary policy is scant because it is difficult to estimate the effects of QE. While the popular view associates greater international spillovers from quantitative easing than conventional monetary shocks, Curcuru et al. (2023) find the opposite. See Bhattarai et al. (2016) for a comprehensive survey on the empirical literature on US unconventional monetary policy.

[39] UIP deviations conditional on US and SOE shocks for the two sub-samples are in Appendix H.

overshooting to US conventional monetary shocks.[40] We argue that delayed overshooting is not inconsistent with UIP. Recall that from Equation (18), UIP in US currency requires:[41]

$$s_t - \mathbb{E}_t\{s_{t+1}\} \approx r_t^* - r_t - \Lambda_t^{US} \tag{20}$$

where $\mathbb{E}_t\{\Lambda_{t+j}^{US}\} = 0$, for all $j \geqslant 0$.

Scholl and Uhlig (2008) state that exchange rate movements depend on the policy rules of both the domestic and foreign countries, where the interest rate differential $r_t^* - r_t > 0$ is necessary for the US dollar to depreciate after an initial appreciation, $|s_0| > 0$.[42] We argue that it is the magnitude of SOE policy rate responses to US shocks that determines the *sign* of the interest rate differentials. If the SOE central banks mimic the Fed's policy rate, these differentials become *negative* (Figure 4, row 3), which when coincide with delayed overshooting results in no excess returns. Put differently, a narrow gap between US and SOE policy rates is the source of hump-shaped dynamics of the exchange rate, which 'ensures' that UIP is not violated. On the other hand, UIP remains intact even when we do not see any evidence of delayed overshooting, as the gap widens and interest rate differentials become *positive* (Figure 5, row 3). We, thus, revise the inference drawn in Scholl and Uhlig (2008) by saying that a combination of *negative* (*positive*) interest rate differential and delayed overshooting (no delayed overshooting) is in fact essential for the UIP to hold. Recent study by Müller et al. (2024) demonstrate that a combination of *positive* interest rate differential and delayed overshooting results in excess returns. For our pre-2008:Q3 case, a combination of *negative* interest rate differential and delayed overshooting offsets the excess returns. As an explanation, consider a US investor who borrows USD for k periods at US interest rate after a US monetary shock, exchanges USD to SOE currency and invests at SOE interest rate, gains excess returns due to higher SOE interest rate (${r_t}^* - r_t < 0$) and exchanges it back to USD after k periods. In order to offset the excess gains made by the investor, USD must overshoot further after the initial appreciation. It

[40] Rüth and Van der Veken (2023) also find little evidence of forward discount premia once they impose an additional IRF restriction on excess bond premium to account for monetary policy propagation through US financial conditions. However, delayed overshooting persists.

[41] See Faust and Rogers (2003) and Scholl and Uhlig (2008) for further discussion on UIP.

[42] Scholl and Uhlig (2008) remove delayed overshooting puzzle by restricting interest rate differential to be positive for the first six months.

implies that delayed overshooting is essential for the UIP to hold, which is reflected in our findings from the pre-2008:Q3 sample. In other words, UIP implies that current exchange rate adjusts depending upon investor's expectations of the future interest rates. As we discuss below, the link between anticipated interest rates and exchange rates further explains the coexistence of UIP and delayed overshooting.

Solving forward for s_t in Equation (20), we obtain:[43]

$$s_t \approx \sum_{j=0}^{\infty} \mathbb{E}_t(r^*_{t+j} - r_{t+j}) + \lim_{j \to \infty} \mathbb{E}_t\{s_{t+j+1}\} \tag{21}$$

From Equation (21), UIP implies that the current exchange rate response depends on the sum of expected future path of interest rate differentials (Engel, 2014; Galí, 2020). From Figure 4, it is not hard to see that in response to US conventional monetary shocks, the sum of interest rate differentials over the horizon is *negative* and *small* for most of the SOEs. Recall from our discussion above that *negative* interest rate differentials may result in excess returns. Therefore, in order to offset the investor's gains from his/her future expectations of interest rate differentials, exchange rate must adjust by overshooting in short to medium term resulting in a *mildly* delayed overshooting that we observe in our pre-2008:Q3 sample. Comparing these results with the period of US unconventional monetary stance in Figure 5, we clearly see that the sum of expected future interest rate differentials is *positive* and *large*. Recall again from the discussion above that *positive* interest rate differentials are essential for the overshooting hypothesis to hold. Therefore, exchange rate immediately responds with a *large* appreciation in order to offset the investor's future expectations of the interest rate differentials.[44,45] As the investor's future expectations of interest rate differentials become narrower in the subsequent horizons, exchange rates respond by depreciating. Note that the appreciation of the currency on impact in response to US conventional monetary shocks is *small* (around 0.4%) and that to US unconventional monetary

[43] See Engel (2014) for the derivation.

[44] Curcuru et al. (2023) and Glick and Leduc (2018) show that USD is more sensitive and has a larger response to a US unconventional monetary policy surprise due to quantitative easing (QE) and forward guidance, compared to that of a conventional monetary shock.

[45] Dedola et al. (2021) also find that QE measures have large and persistent effects on the exchange rate. They further report contribution of the "signalling" channel of QE to exchange rate response through changes in expectations about the future monetary policy stance.

shocks is relatively *large* (around 1.5%).[46] UIP implies that these impact responses are derived by the sum of expected future interest rate differentials, which are narrow (wide) during the times of US conventional (unconventional) monetary policy stance. Summing up, the *mildly* delayed overshooting that we see in response to US conventional monetary shocks is a characteristic of *small negative* sum of expected future interest rate differentials. On the other hand, we find no evidence of delayed overshooting in response to US unconventional monetary shocks where the sum of expected future interest rate differentials is *large* and *positive*. These findings further substantiate our claim that delayed overshooting may not be inconsistent with UIP, rather an artifact of SOE central bank's response that triggers an adjustment of investor's future expectations of interest rate differentials.

The results in this section have conveyed that (i) delayed exchange rate overshooting is not robust across samples, (ii) when exchange rate overshooting is indeed observed to be delayed, UIP still holds, and (iii) delayed overshooting is essential for the UIP to hold when interest rate differentials are *negative*. We assert that delayed overshooting may neither be evidence against UIP nor a puzzle. Rather, it may merely be an artifact of the endogenous responses of SOE policy rates to US conventional monetary surprises. That UIP and delayed exchange rate overshooting can in fact be in harmony has also been pointed out by Rüth and Van der Veken (2023), who show that UIP is fulfilled even with delayed overshooting. Also, Müller et al. (2024) find that delayed overshooting may not be a failure of UIP, and may just reflect a sluggish adjustment of exchange rate expectations by market participants. Neither paper focuses on the relevance of both domestic and foreign monetary policy behavior in the context of these exchange rate puzzles. However, the findings of this strand of the literature that confirm the comfortable coexistence of UIP and delayed overshooting, stand in stark contrast to the early findings of Scholl and Uhlig (2008). They had found empirical support for the forward discount puzzle - equivalently, the failure of UIP - with or without delayed overshooting.

[46] Glick and Leduc (2018) report roughly three to four times larger impact on USD to monetary policy surprises since the Fed lowered its policy rate to the effective lower bound.

6 Robustness Checks

We perform two robustness checks: (1) excluding ZLB episodes and (2) including an SOE-specific "Commodity Terms of Trade".[47]

6.1 Excluding ZLB episodes

Figure 7 (blue solid lines) shows RER responses of SOE variables to a one standard deviation SOE contractionary monetary policy surprise using the sample period 1992:Q1 - 2008:Q3, while the shaded bands are the associated 68% posterior probability bands.[48] Figure 8 (blue solid lines) plots the deviations from UIP based on the estimation period 1992:Q1 - 2008:Q3. We conclude that our main results are qualitatively and quantitatively robust to excluding the ZLB episodes.[49]

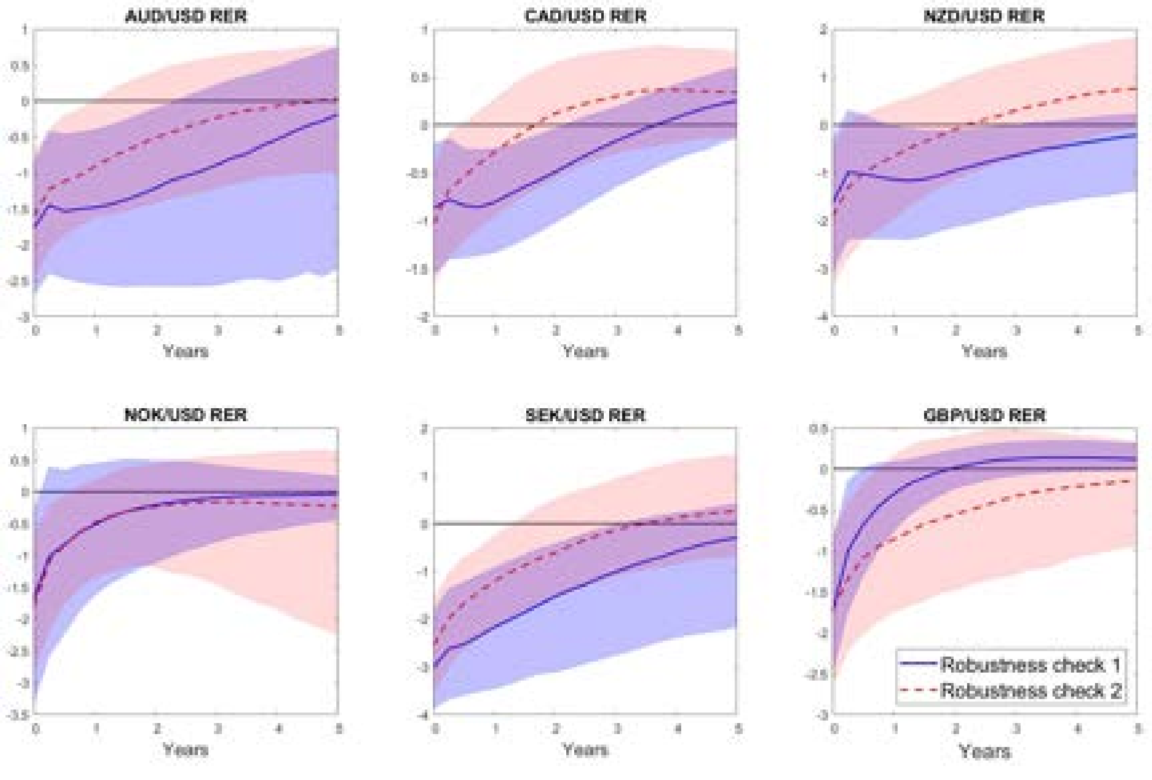

Figure 7: IRFs of SOE/USD RER to a SOE contractionary monetary shock identified using block exogeneity and Restrictions 1 to 4. *Note:* Solid (dashed) lines depict point-wise posterior median IRFs for robustnessness check 1 (2). The shaded regions represent the associated 68% posterior probability bands.

[47]We set the lag order $p = 1$ for both robustness checks.
[48]For the US, Sweden and the UK, we replace the shadow rate with the 3-month interbank rate.
[49]IRFs of SOE variables to SOE shocks for pre-GFC sample without Restriction 4 ($\psi_e > 0$) are provided in Appendix I.

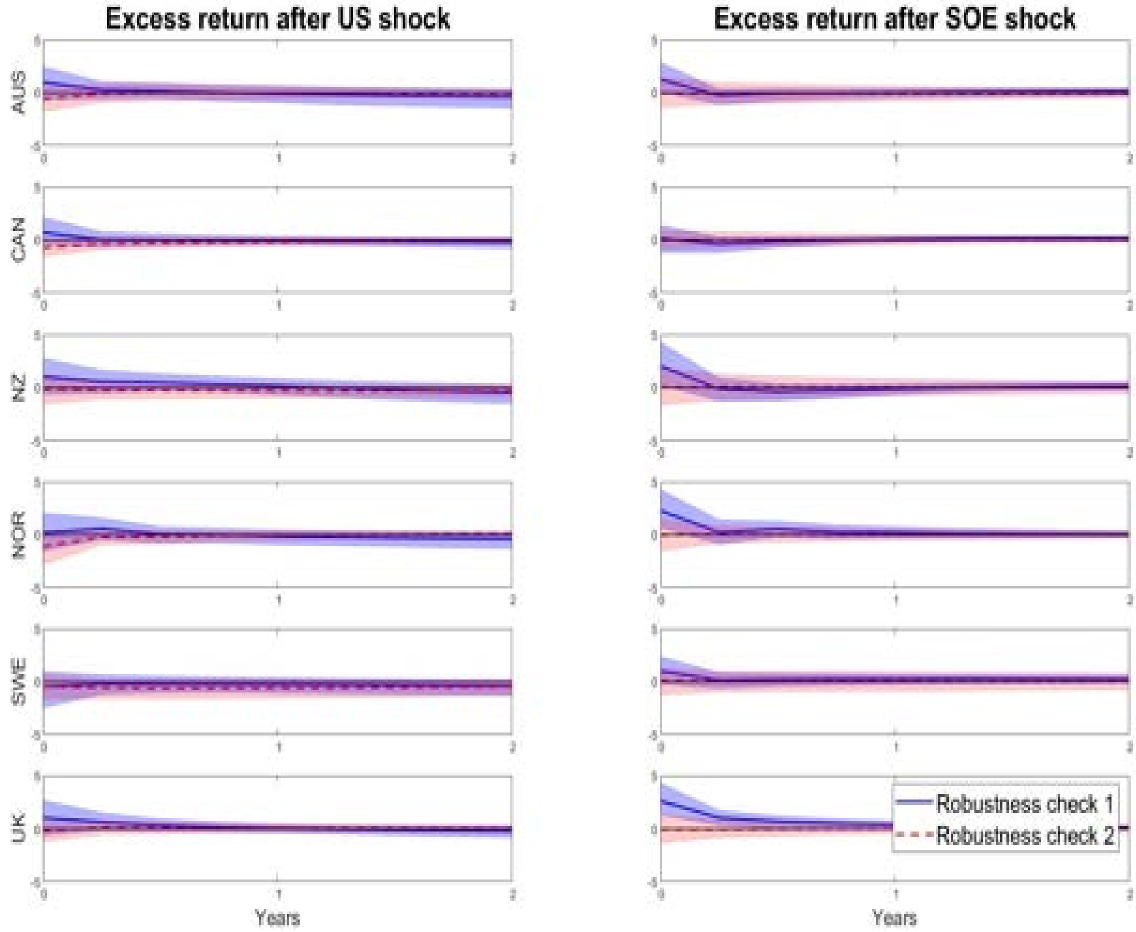

Figure 8: Deviations from UIP conditional on US (left) and SOE (right) monetary policy shocks. *Note:* The blue solid (red dashed) lines depict the point-wise posterior median responses of robustness check 1 (2). The shaded areas are the 68% posterior probability intervals.

6.2 Including an SOE Commodity Terms of Trade

The second check aims to test the robustness of our baseline results with those of a 9-variable SOE-SVAR model for the six SOEs using the full sample. This larger model includes an SOE-specific "Commodity Terms of Trade" in the US block. We draw motivation for this robustness check as four of our SOEs are major commodity exporters (Australia, Canada, Norway, and the UK). We extend Equation (13) by including the commodity Terms of Trade (cp^*), however we do not apply any further restrictions:

$$r_t^* = \underbrace{-a_{0,51}^{-1}a_{0,11}}_{\psi_{y^*}>0} y_t^* \underbrace{-a_{0,51}^{-1}a_{0,21}}_{\psi_{\pi^*}>0} \pi_t^* \underbrace{-a_{0,51}^{-1}a_{0,31}}_{\psi_{cs^*}<0} cs_t^* \underbrace{-a_{0,51}^{-1}a_{0,41}}_{unrestricted} cp_t^* \underbrace{+a_{0,51}^{-1}}_{\sigma^*} \epsilon_{1,t} \tag{22}$$

Figure 7 (red dashed lines) plots the IRFs of SOE/USD RER variables to a one standard deviation SOE contractionary monetary policy surprises. We find no evidence of delayed overshooting for the six SOEs - an appreciation on impact is followed by an immediate depreciation. Similarly, Figure 8 (red dashed lines) reports no evidence of UIP violations in our 9-variable SOE-SVAR models for the six economies. Interestingly, including the "Commodity Terms of Trade" helps in addressing the 2 year overshoot of AUD/USD RER and the borderline cases of UIP deviations in our baseline results. We conclude that our results are robust to including the SOE-specific commodity TOT in the US block.[50,51]

7 Conclusion

The extant SVAR literature has focused on domestic monetary policy shocks in the context of the exchange rate overshooting hypothesis of Dornbusch (1976), and its crucial building block, uncovered interest rate parity (UIP). In contrast, this paper tests these theories of the exchange rate by estimating six empirical *two-country* models identified *jointly* through the systematic components of SOE and US monetary policy rules. The principal innovation of our framework involves combining the block exogeneity structure of traditional SOE-SVARs with the imposition of sign restrictions on the structural monetary policy rules in *both* countries. The latter part of the identification strategy unravels the endogenous monetary policy responses of SOE central banks to US monetary policy shocks.

Our baseline results for the sample 1992:Q1-2019:Q4 demonstrates that exchange rate dynamics are typically in line with the Dornbusch hypothesis and UIP: a contractionary domestic monetary policy shock triggers an immediate appreciation of the domestic currency followed by a depreciation. We demonstrate that two aspects of the joint identification strategy are crucial for this result: the endogenous response of the SOE central bank to the exchange rate, and the response of the Federal Reserve to the US credit spread.

[50] Our results are robust for the 9-variable SOE-SVAR models excluding the zero lower bound (ZLB) period.

[51] Our results are robust when we replace commodity TOT with Spot Crude Oil Price: West Texas Intermediate (https://fred.stlouisfed.org/series/WTISPLC), following Terrell et al. (2023).

However, an anomaly emerges when we estimate the models on samples ending with and following the global financial crisis. Conditional on US monetary shocks, exchange rate overshooting appears to be delayed in several SOE-SVARs estimated over 1992:Q1-2008:Q3. Nevertheless, we emphasize that delayed overshooting is not a violation of UIP but rather an artifact of a narrow interest rate gap between the SOE and the US. After a contractionary US monetary shock, central banks in SOEs also raise their policy rates, likely reacting to the depreciation of their currencies. This narrows the interest rate gaps between the two countries. For UIP to hold in this context, exchange rate overshooting is delayed. It is our novel joint identification scheme that brings to the forefront the exchange rate effects of the interplay between domestic and foreign monetary policy.

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Appendix

A Data Sources

The dataset spans from 1992:Q1 to 2019:Q4. Where required, series are seasonally adjusted using Eviews Census X-13.

A.1 United States

- **Real Gross Domestic Product** (Billions of Chained 2012 Dollars, Seasonally Adjusted Annual Rate)
  - Source: FRED Economic Data (GDPC1)
- **Consumer Price Index: All Items for the United States** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (USACPIALLMINMEI)
- **Federal Funds Effective Rate** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (FEDFUNDS)
- **Shadow rate** (Percent)
  - Source: Wu and Xia (2016)
- **Moody's Seasoned Baa Corporate Bond Yield Relative to Yield on 10-Year Treasury Constant Maturity** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (BAA10YM)

A.2 Australia

- **Real Gross Domestic Product for Australia** (Domestic Currency, Seasonally Adjusted)
  - Source: FRED Economic Data (NGDPRSAXDCAUQ)
- **Consumer Price Index: All Items: Total: Total for Australia** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (AUSCPIALLQINMEI)
- **3-Month or 90-day Rates and Yields: Interbank Rates for Australia** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (IR3TIB01AUQ156N)

- **Exchange rate** **
  - Source: Reserve Bank of Australia (Refinitiv Datastream, AUUSDSP)
- **Commodity Terms of Trade**
  - Description: Commodity Export Price Index, Individual Commodites Weighted by Ratio of Exports to GDP.
  - Source: International Monetary Fund (Country code: 193)
    - https://www.imf.org/en/Research/commodity-prices

A.3 Canada

- **Real Gross Domestic Product for Canada** (Domestic Currency, Seasonally Adjusted)
  - Source: FRED Economic Data (NGDPRSAXDCCAQ)
- **Consumer Price Index: All Items: Total: Total for Canada** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (CANCPIALLQINMEI)
- **3-Month or 90-day Rates and Yields: Interbank Rates for Canada** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (IR3TIB01CAQ156N)
- **Exchange rate**
  - Source: Bank of Canada (Refinitiv Datastream, CNXRUSD)
- **Commodity Terms of Trade**
  - Description: Commodity Export Price Index, Individual Commodites Weighted by Ratio of Exports to GDP.
  - Source: International Monetary Fund (Country code: 156)
    - https://www.imf.org/en/Research/commodity-prices

**For each SOE, we calculate real exchange rate from the nominal exchange rate and the US and domestic price levels, such that, $ln(e_t) = ln(s_t) + ln(p_t^*) - ln(p_t)$; where $ln(e_t)$ and $ln(s_t)$ are the logs of real and nominal exchange rates, respectively and $ln(p_t^*)$ and $ln(p_t)$ are the logs of US and domestic consumer price indices, respectively.

A.4 New Zealand

- **Production-based gross domestic product (GDP)** (Real, NZD, Seasonally Adjusted)
  - Source: Statistics New Zealand (GDP06.Q.QT0.rs)
- **Consumer Price Index: All Items for New Zealand** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (NZLCPIALLQINMEI)
- **3-Month or 90-day Rates and Yields: Interbank Rates for New Zealand** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (IR3TIB01NZQ156N)
- **Exchange rate**
  - Source: Reserve Bank of New Zealand (EXR.MS11.D06)
- **Commodity Terms of Trade**
  - Description: Commodity Export Price Index, Individual Commodites Weighted by Ratio of Exports to GDP.
  - Source: International Monetary Fund (Country code: 196)
    - https://www.imf.org/en/Research/commodity-prices

A.5 Norway

- **Real Gross Domestic Product for Norway** (Millions of Chained 2010 National Currency, Seasonally Adjusted)
  - Source: FRED Economic Data (CLVMNACSCAB1GQNO)
- **Consumer Price Index: All Items for Norway** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (NORCPIALLQINMEI)
- **3-Month or 90-day Rates and Yields: Interbank Rates for Norway** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (IR3TIB01NOQ156N)
- **Exchange rate**
  - Source: Norges Bank (Refinitiv Datastream, NWXRUSD)
- **Commodity Terms of Trade**
  - Description: Commodity Export Price Index, Individual Commodites Weighted by Ratio of Exports to GDP.

- Source: International Monetary Fund (Country code: 142)
  - https://www.imf.org/en/Research/commodity-prices

A.6 Sweden

- **Real Gross Domestic Product for Sweden** (Millions of Chained 2010 National Currency, Seasonally Adjusted)
  - Source: FRED Economic Data (CLVMNACSCAB1GQSE)
- **Consumer Price Index: All Items for Sweden** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (SWECPIALLQINMEI)
- **3-Month or 90-day Rates and Yields: Interbank Rates for Sweden** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (IR3TIB01SEQ156N)
- **Shadow rate** (Percent)
  - Source: De Rezende and Ristiniemi (2023)
- **Exchange rate**
  - Source: Sveriges Riksbank Bank (Refinitiv Datastream, SDXRUSD)
- **Commodity Terms of Trade**
  - Description: Commodity Export Price Index, Individual Commodites Weighted by Ratio of Exports to GDP.
  - Source: International Monetary Fund (Country code: 144)
    - https://www.imf.org/en/Research/commodity-prices

A.7 United Kingdom

- **Real Gross Domestic Product for United Kingdom** (Millions of Chained 2010 National Currency, Seasonally Adjusted)
  - Source: FRED Economic Data (CLVMNACSCAB1GQUK)
- **Consumer Price Index of All Items in the United Kingdom** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (GBRCPIALLQINMEI)
- **3-Month or 90-day Rates and Yields: Interbank Rates for the United Kingdom** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (IR3TIB01GBQ156N)

- **Shadow rate** (Percent)
  - Source: https://sites.google.com/view/jingcynthiawu/shadow-rates
- **Exchange rate**
  - Source: Bank of England (Refinitiv Datastream, UKXRUSD)
- **Commodity Terms of Trade**
  - Description: Commodity Export Price Index, Individual Commodites Weighted by Ratio of Exports to GDP.
  - Source: International Monetary Fund (Country code: 112)
    - https://www.imf.org/en/Research/commodity-prices

B Technical Appendix (Not for Publication)

B.1 The Model

Our methodology draws heavily on Arias et al. (2018) and Arias et al. (2019). First, consider the SVAR model:

$$\mathbf{y_t'A_0} = \sum_{l=1}^{p} \mathbf{y}_{t-l}'\mathbf{A}_l + \mathbf{c}' + \epsilon_{\mathbf{t}}' \ , \qquad \text{for } 1 \leq \mathbf{t} \leq \mathbf{T} \tag{23}$$

where $\mathbf{y_t'} = \begin{bmatrix} \mathbf{y_{1t}'} & \mathbf{y_{2t}'} \end{bmatrix}$. $\mathbf{y_{1t}}$ is a $(n_1 \times 1)$ vector of US variables and $\mathbf{y_{2t}}$ is a $(n_2 \times 1)$ vector of SOE variables, with $n = n_1 + n_2$ denoting the total number of variables. Similarly, the vector of structural shocks $\epsilon_{\mathbf{t}}$ is divided into two blocks, $\epsilon_{\mathbf{t}}{}' = \begin{bmatrix} \epsilon_{1t}' & \epsilon_{2t}' \end{bmatrix}$. $\mathbf{A_i}$, for $0 \leq i \leq p$, are $(n \times n)$ matrices of structural parameters, with $\mathbf{A_0}$ invertible. $\mathbf{c}$ is a $(n \times 1)$ vector of constants, p is the lag length, and $\mathbf{T}$ is the sample size. Conditional on past information and initial conditions $\mathbf{y}_0, ..., \mathbf{y}_{1-p}$, the vector $\epsilon_{\mathbf{t}}$ is Gaussian with mean zero and covariance matrix $\mathbb{I}_n$. Following Rubio-Ramirez et al. (2010), we can write the SVAR in compact form:

$$\mathbf{y_t'A_0} = \mathbf{x_t'A_+} + \epsilon_t' \ , \tag{24}$$

where $\mathbf{x_t'} = [\mathbf{y_{t-1}'} \dots \mathbf{y_{t-p}'} \ \ 1]$. $\mathbf{A_0}$ and $\mathbf{A_+} = [\mathbf{A_1'} \dots \mathbf{A_p'} \ \ \mathbf{c}']$ are matrices of structural parameters.

Post-multiplying Equation (24) by $\mathbf{A_0^{-1}}$, we obtain the reduced form VAR model:

$$\mathbf{y_t'} = \mathbf{x_t'B} + u_t' \ , \tag{25}$$

where $\mathbf{B} = \mathbf{A_+A_0^{-1}}, u_t' = \epsilon_t'\mathbf{A_0^{-1}}$ and $\mathbb{E}[u_t u_t'] = \mathbf{\Sigma} = (\mathbf{A_0A_0'})^{-1}$. $\mathbf{B}$ is the matrix of reduced-form coefficients and $\mathbf{\Sigma}$ is the residual variance-covariance matrix.

B.2 Impulse response functions

The impulse response functions (IRFs) are defined as follows:

Definition 1. *Let* $(\mathbf{A_0}, \mathbf{A}_+)$ *be any value of structural parameters: The IRF of the* i*-th variable to the* j*-th structural shock at finite horizon* h *corresponds to the element in row* i *and column* j *of the matrix*

$$\mathbf{L}_h(\mathbf{A_0}, \mathbf{A}_+) = (\mathbf{A_0}^{-1}\mathbf{J}'\mathbf{F}^h\mathbf{J})', \text{ where } \mathbf{F} = \begin{bmatrix} \mathbf{A}_1\mathbf{A_0}^{-1} & \mathbb{I}_n & \dots & \mathbf{0} \\ \vdots & \vdots & \ddots & \vdots \\ \mathbf{A}_{p-1}\mathbf{A_0}^{-1} & \mathbf{0} & \dots & \mathbb{I}_n \\ \mathbf{A}_p\mathbf{A_0}^{-1} & \mathbf{0} & \dots & \mathbf{0} \end{bmatrix} \text{ and } \mathbf{J} = \begin{bmatrix} \mathbb{I}_n \\ \mathbf{0} \\ \vdots \\ \mathbf{0} \end{bmatrix}$$

Importantly note that $\mathbf{L}_h(\mathbf{A}_0\mathbf{Q}, \mathbf{A}_+\mathbf{Q}) = \mathbf{L}_h(\mathbf{A}_0, \mathbf{A}_+)\mathbf{Q}$ for $0 \leq h \leq \infty$ and $\mathbf{Q} \in O(n)$, where $O(n)$ denotes the set of all orthogonal matrices with dimensions n x n.

B.3 Set Identification by Sign and Zero Restrictions on $\mathbf{A}_0$

We follow closely the work of Arias et al. (2019) to achieve set identification by imposing sign and exclusion restrictions directly on the structural parameters in $\mathbf{A}_0$ matrix. Further, we impose two normalizing restrictions on the IRFs of US and SOE policy rates.

B.3.1 Sign restrictions

First, we impose sign restrictions on some elements of $\mathbf{A}_0$ and on IRFs of US and SOE policy rates at horizon 0. Following Arias et al. (2019), we stack $\mathbf{A}_0$ matrix and the IRFs at horizon 0 into a single matrix of dimension $k \times n$, which is denoted by $\mathbf{F}(\mathbf{A_0}, \mathbf{A}_+)$:

$$\mathbf{F}(\mathbf{A_0}, \mathbf{A}_+) = \begin{bmatrix} \mathbf{A}_0 \\ \mathbf{L_0}(\mathbf{A_0}, \mathbf{A}_+) \end{bmatrix}_{k \times n} \text{ where } k = 2n \text{ in our case.}$$

Sign restrictions on $\mathbf{F}(\mathbf{A_0}, \mathbf{A}_+)$ are represented by matrices $\mathbf{S}_j$ for $1 \leq j \leq n$, where the number of columns in $\mathbf{S}_j$ is equal to the number of rows in $\mathbf{F}(\mathbf{A_0}, \mathbf{A}_+)$. $\mathbf{S}_j$ is a selection matrix with only one non-zero entry in each row. If the rank of $\mathbf{S}_j$ is s_j, then s_j is the number of sign restrictions imposed to identify the j-th structural shock. The total number of sign restrictions are $s = \sum_{j=1}^{n} s_j$.[52]

Definition 2. *Let* $(\mathbf{A_0},\ \mathbf{A}_+)$ *be any value of structural parameters. These parameters satisfy the sign restrictions if and only if*

$$\mathbf{S}_j\mathbf{F}(\mathbf{A_0}, \mathbf{A}_+)\mathbf{e}_j > \mathbf{0}, \text{ for } 1 \leq j \leq n$$

where $\mathbf{e}_j$ denotes the j-th column of an identity matrix $\mathbb{I}_n$ of dimensions $n \times n$.

Algorithm 1: The following algorithm independently draws from the normal-generalized-normal $\mathcal{NGN}(\alpha,\ \beta,\ S,\ \Omega)$ distribution over the structural parameterization conditional on the sign restrictions.

1. Draw reduced-form parameters $(\mathbf{B}, \mathbf{\Sigma})$ independently from the $\mathcal{NIW}(\alpha,\ \beta,\ S,\ \Omega)$ distribution.
2. Draw $\mathbf{Q}$ independently from the uniform distribution over $O(n)$ using Theorem 1 (below), where $\mathbf{Q}$ is an $n \times n$ orthogonal matrix.
3. Keep $(\mathbf{A_0}, \mathbf{A}_+) = f_{Chol}^{-1}(\mathbf{B}, \mathbf{\Sigma}, \mathbf{Q})$ if the sign restrictions are satisfied.[53] This step is equivalent to keeping the draw if $\mathbf{S}_j\mathbf{F}(\mathbf{U}^{-1}\mathbf{Q}, \mathbf{B}\mathbf{U}^{-1}\mathbf{Q})\mathbf{e}_j > \mathbf{0}$, for $1 \leq j \leq n$, where $\mathbf{U}$ is upper triangular with positive diagonal and $\mathbf{\Sigma} = \mathbf{U}'\mathbf{U}$ is the Cholesky decomposition of $\mathbf{\Sigma}$.
4. Return to Step 1 until the required number of draws has been obtained.

To efficiently implement the first step of Algorithm 1 where each draw is serially uncorrelated, we follow Arias et al. (2018) and exploit the fact that the space of all structural parameters is equivalent to the product of the space of all reduced-form parameters and $O(n)$. This mapping is given by $(\mathbf{B}, \mathbf{\Sigma}, \mathbf{Q}) \rightarrow (\mathbf{A_0}, \mathbf{A}_+) = (\mathbf{U}^{-1}\mathbf{Q}, \mathbf{B}\mathbf{U}^{-1}\mathbf{Q})$,

[52] See Arias et al. (2013) for specific examples on imposing sign restrictions using $\mathbf{S}_j$ matrices.
[53] f_{Chol} defines the mapping from $(\mathbf{A_0}, \mathbf{A}_+)$ to $(\mathbf{B}, \mathbf{\Sigma}, \mathbf{Q})$ as:

$f_{Chol}(\mathbf{A_0}, \mathbf{A}_+) = (\underbrace{\mathbf{A}_+\mathbf{A_0}^{-1}}_{\mathbf{B}}, \underbrace{(\mathbf{A_0}\mathbf{A}_0')^{-1}}_{\mathbf{\Sigma}}, \underbrace{Chol((\mathbf{A_0}\mathbf{A}_0')^{-1})\mathbf{A_0}}_{\mathbf{Q}})$; where *Chol* is the Cholesky decomposition. See Arias et al. (2018) for details.

where $\mathbf{\Sigma} = \mathbf{U}'\mathbf{U}$ is the Cholesky decomposition of $\mathbf{\Sigma}$ such that $\mathbf{U}$ is upper triangular with positive diagonal and $\mathbf{Q}$ is an element of $O(n)$. The likelihood is flat over the space of orthogonal matrices because $(\mathbf{U}^{-1}, \mathbf{B}\mathbf{U}^{-1})$ is observationally equivalent to $(\mathbf{U}^{-1}\mathbf{Q}, \mathbf{B}\mathbf{U}^{-1}\mathbf{Q})$. Also, the likelihood at the reduced-form parameters $(\mathbf{B}, \mathbf{\Sigma})$ will be equal to the likelihood at the structural parameters $(\mathbf{U}^{-1}, \mathbf{B}\mathbf{U}^{-1})$. A prior on the reduced-form parameters, together with the uniform distribution, induces a prior on the structural parameters via the above mapping. Thus, if we independently draw $(\mathbf{B}, \mathbf{\Sigma}, \mathbf{Q})$ from a uniform-independent-normal-inverse-Wishart distribution over the orthogonal reduced-form parameterization with parameters α, β, S, Ω and then transform the draws to $(\mathbf{A_0}, \mathbf{A}_+)$ using a mapping function, f_{Chol}^{-1}. We, then, can independently draw from a normal-generalized-normal distribution over the structural parameterization, denoted by $\mathcal{NGN}(\alpha, \beta, S, \Omega)$, where its density over the structural parameterization induced by the uniform-normal-inverse-Wishart density over the orthogonal reduced-form parameterization is denoted by:[54]

$$\begin{aligned}\mathcal{NGN}_{(\alpha,\beta,S,\Omega)}(\mathbf{A_0}, \mathbf{A}_+) &= \mathcal{UNIW}_{(\alpha,\beta,S,\Omega)}(f_{Chol}(\mathbf{A_0}, \mathbf{A}_+))\vartheta_{f_{Chol}}(\mathbf{A_0}, \mathbf{A}_+) \\ &\propto \underbrace{|\det(\mathbf{A_0})|^{\nu-n} e^{-\frac{1}{2}\mathrm{vec}(\mathbf{A_0})'(\mathbb{I}_n \bigotimes \beta)\mathrm{vec}(\mathbf{A_0})}}_{\mathrm{generalized-normal}} \\ &\times \underbrace{e^{-\frac{1}{2}\mathrm{vec}(\mathbf{A}_+ - S\mathbf{A_0})'(\mathbb{I}_n \bigotimes \Omega)^{-1}\mathrm{vec}(\mathbf{A}_+ - S\mathbf{A_0})}}_{\mathrm{conditionally-normal}}\end{aligned} \tag{26}$$

where $\vartheta_{f_{Chol}}(\mathbf{A_0}, \mathbf{A}_+) = 2^{\frac{n(n+1)}{2}}|\det(\mathbf{A_0})|^{-(2n+m+1)}$ is the volume element of f_{Chol} at $(\mathbf{A}_0, \mathbf{A}_+)$.

It is easy to independently draw from the uniform-normal-inverse-Wishart distribution. Matlab has routines for making independent draws from both the inverse-Wishart distribution and the normal distribution. There are efficient algorithms for making independent draws from the uniform distribution over $O(n)$. Canova and De Nicolo (2002), Uhlig (2005), and Rubio-Ramirez et al. (2010) all proposed algorithms to do this. The algorithm of Rubio-Ramirez et al. (2010) is the most efficient, particularly for larger SVAR systems (e.g., $n > 4$).[55] Rubio-Ramirez

[54]See Arias et al. (2018) and Arias et al. (2013) for details.
[55]See Rubio-Ramirez et al. (2010) for details.

et al. (2010) results are based on the following well-known theorem (corresponds to theorem 4 in Arias et al. (2018)).

Theorem 1: *Let* $\mathbf{X}$ *be an* $n \times n$ *random matrix with each element having an independent standard normal distribution. Let* $\mathbf{X} = \mathbf{QR}$ *be the* QR *decomposition of* $\mathbf{X}$ *with the diagonal of* $\mathbf{R}$ *normalized to be positive. The random matrix* $\mathbf{Q}$ *is orthogonal and is a draw from the uniform distribution over* $O(n)$*.*[56]

B.3.2 Zero restrictions

Second, we impose zero restrictions on some elements of $\mathbf{A_0}$. For the sake of consistency, we use the function $\mathbf{F}(\mathbf{A_0}, \mathbf{A_+}) = \mathbf{A_0}$ to impose zero restrictions on $\mathbf{A_0}$ matrix. In our case we do not impose any zero restrictions on the IRFs but this can easily be extended by stacking $\mathbf{L_0}(\mathbf{A_0}, \mathbf{A_+})$ in $\mathbf{F}(\mathbf{A_0}, \mathbf{A_+})$, similar to the case of imposing sign restrictions. Zero restrictions can be represented by matrices $\mathbf{Z}_j$ for $1 \leq j \leq n$, where the number of columns in $\mathbf{Z}_j$ is equal to the number of rows in $\mathbf{F}(\mathbf{A_0}, \mathbf{A_+})$. If the rank of $\mathbf{Z}_j$ is z_j, then z_j is the number of zero restrictions associated with the j-th structural shock. The total number of zero restrictions is $z = \sum_{j=1}^{n} z_j$.[57]

Definition 2. *Let* $(\mathbf{A_0},\, \mathbf{A_+})$ *be any value of structural parameters. These parameters satisfy the zero restrictions if and only if*

$$\mathbf{Z}_j \mathbf{F}(\mathbf{A_0}, \mathbf{A_+})\mathbf{e}_j = \mathbf{0}, \text{ for } 1 \leq j \leq n \ .$$

Similar to sign restrictions, $\mathbf{e}_j$ denotes the j-th column of $\mathbb{I}_n$, where $\mathbb{I}_n$ is the identity matrix of dimension $n \times n$.

From the definition of f_{Chol} and the fact that $\mathbf{F}(\mathbf{A}_0\mathbf{Q}, \mathbf{A}_+\mathbf{Q}) = \mathbf{F}(\mathbf{A}_0, \mathbf{A}_+)\mathbf{Q}$, the zero restrictions in the orthogonal reduced-form parameterization are:

$$\mathbf{Z}_j \mathbf{F}(f^{-1}_{Chol}(\mathbf{B}, \mathbf{\Sigma}, \mathbf{Q}))\mathbf{e}_j = \mathbf{Z}_j \mathbf{F}(f^{-1}_{Chol}(\mathbf{B}, \mathbf{\Sigma}, \mathbb{I}_n))\mathbf{Q}\mathbf{e}_j = 0 \quad \text{for } 1 \leq j \leq n$$

[56] For proof see Stewart (1980).

[57] See Arias et al. (2013) for specific examples on imposing zero-restrictions using $\mathbf{Z}_j$ matrices.

This means that the zero restrictions in the orthogonal reduced-form parameterization are really just linear restrictions on each column of the orthogonal matrix $\mathbf{Q}$, conditional on the reduced-form parameters $(\mathbf{B}, \mathbf{\Sigma})$. It is this observation that is key to being able to make independent draws from the set of all structural parameters satisfying the zero restrictions (Arias et al., 2018).

Algorithm 2: The following algorithm makes independent draws from a distribution over the structural parameterization conditional on the zero restrictions. We use it to impose zero restrictions on $\mathbf{A_0}$ in order to implement block exogeneity.

1. Draw $(\mathbf{B}, \mathbf{\Sigma})$ independently from the $\mathcal{NIW}(\alpha,\ \beta,\ S,\ \Omega)$ distribution.
2. For $1 \leq j \leq n$, draw $\mathbf{x}_j \in \mathbb{R}^{n+1-j-z_j}$ independently from a standard normal distribution and set $\mathbf{w}_j = \mathbf{x}_j/||\mathbf{x}_j||$.
3. Define $\mathbf{Q} = [\mathbf{q}_1 \dots \mathbf{q}_n]$ recursively by $\mathbf{q}_j = \mathbf{K}_j\mathbf{w}_j$ for any matrix $\mathbf{K}_j$ whose columns form an orthonormal basis for the null space of the $(j-1+z_j) \times n$ matrix.[58]

$$\mathbf{M}_j = [\mathbf{q}_1 \dots \mathbf{q}_{j-1} \quad (\mathbf{Z}_j\mathbf{F}(f^{-1}_{Chol}(\mathbf{B}, \mathbf{\Sigma}, \mathbb{I}_n)))']'.$$

By construction, $\mathbf{q}_j$ is perpendicular to the rows of $\mathbf{M}_j$ and $||\mathbf{x}_j|| = ||\mathbf{w}_j|| = 1$. Thus, the matrix $\mathbf{Q}$ obtained through Steps 2 and 3 is orthogonal and

$$\mathbf{Z}_j\mathbf{F}(f^{-1}_{Chol}(\mathbf{B}, \mathbf{\Sigma}, \mathbb{I}_n))\mathbf{Q}\mathbf{e}_j = 0 \quad \text{for } 1 \leq j \leq n$$

4. Set $(\mathbf{A_0}, \mathbf{A}_+) = f^{-1}_{Chol}(\mathbf{B}, \mathbf{\Sigma}, \mathbf{Q})$.
5. Return to Step 1 until the required number of draws has been obtained.

B.3.3 Implementing block exogeneity

From Equation 23:[59]

$$\mathbf{y_t'} = \begin{bmatrix} \mathbf{y_{1t}'} & \mathbf{y_{2t}'} \end{bmatrix}, \quad \mathbf{A}_l = \begin{bmatrix} A_{11,l} & A_{12,l} \\ \\ A_{21,l} & A_{22,l} \end{bmatrix}, \quad \epsilon_\mathbf{t}' = \begin{bmatrix} \epsilon_{1t}' & \epsilon_{2t}' \end{bmatrix}$$

[58] The computation of each $\mathbf{K}_j$ requires a single QR-decompostion of an $n \times n$ invertible matrix. In Matlab $\mathbf{K}_j = \text{null}(\mathbf{M}_j)$.

[59] For simplicity, we omit constant terms in the exposition.

where $\mathbf{y}'_{1t}$ is $1 \times n_1$, $\mathbf{y}'_{2t}$ is $1 \times n_2$, $n_1 + n_2 = n$, $A_{11,l}$ is $n_1 \times n_1$, $A_{12,l}$ is $n_1 \times n_2$, $A_{21,l}$ is $n_2 \times n_1$, $A_{22,l}$ is $n_2 \times n_2$, ϵ'_{1t} is $1 \times n_1$ and ϵ'_{2t} is $1 \times n_2$.

Cushman and Zha (1997) implement block exogeneity by restricting $A_{21,l} = 0$ for $l = 0, 1, ..., p$. These restrictions mean that the domestic block ($\mathbf{y}'_{2t}$) does not influence the foreign block ($\mathbf{y}'_{1t}$), neither contemporaneously, nor in a dynamic way through lags. However, when using the methodology of Arias et al. (2018), one can only impose $(n - k)$ zero restrictions per equation, where k corresponds to the k^{th} equation in the system.[60] As a result, we can only impose zero restrictions on the contemporaneous structural matrix $\mathbf{A}_0$: $A_{21,l} = 0$, for $l = 0$, and are unable to impose $A_{21,l} = 0$ for $l = 1, ..., p$. To circumvent this technical obstacle, we implement block exogeneity on the lagged matrices by using extremely tight priors with distribution centered at zero for all the reduced-form coefficients of domestic variables appearing in the foreign block. Specifically, we use a special case of Independent Normal Inverse Wishart ($\mathcal{NIW}$) priors where we adopt conventional Minnesota priors for the reduced-form VAR coefficients, β.[61]

B.3.4 Independent $\mathcal{NIW}$ priors and posteriors

Arias et al. (2019) use Natural Conjugate $\mathcal{NIW}$ priors for the reduced form parameters, $\mathbf{B}$ and $\mathbf{\Sigma}$:

$$\beta \sim \mathcal{N}(\beta_0, \mathbf{\Sigma} \otimes \Phi_0), \tag{27}$$

$$\mathbf{\Sigma} \sim \mathcal{IW}(S_0, \alpha_0), \tag{28}$$

where matrix Φ_0 contains the hyperparameters that control the tightness of distribution of reduced form parameters, the residual variance-covariance matrix $\mathbf{\Sigma}$ has an inverse Wishart distribution with scale matrix S_0 and degrees of freedom α_0.

However, these priors are not suited for the purpose of implementing block exogeneity. Natural Conjugate $\mathcal{NIW}$ priors employ Kronecker structure for the variance-covariance matrix of the reduced-form parameters. Hence, the variances are proportional to one another. As a result, imposing block exogeneity on one equation would impose it on all equations (Dieppe et al., 2016; Koop et al., 2010). Moreover, the

[60] See also Arias et al. (2021) page # 92 and Kilian and Lütkepohl (2017) page # 481.

[61] $\beta = vec(\mathbf{B})$.

structure implied by the Kronecker product requires that every equation has the same set of explanatory variables (Koop et al., 2010), meaning that if we remove a variable in one equation, that variable would be removed from all equations.

The techniques developed by Arias et al. (2018) can be used for any prior distributions. Hence, instead of using Natural Conjugate $\mathcal{NIW}$ priors which are problematic for our purpose of implementing block exogeneity, we use Independent $\mathcal{NIW}$ priors over the reduced-form parameters. The prior distributions of the reduced-form parameters from Equation (25) take the form:

$$\beta \sim \mathcal{N}(\beta_0, \Omega_0), \tag{29}$$

$$\mathbf{\Sigma} \sim \mathcal{IW}(S_0, \alpha_0), \tag{30}$$

where the residual variance-covariance matrix $\mathbf{\Sigma}$ has an inverse Wishart distribution with scale matrix S_0 and degrees of freedom α_0.[62]

We follow standard Minnesota priors for the prior distribution of the reduced form coefficients β in Equation (29), where the distribution is centered at 1 for the coefficients of own 1st lags, and for the rest, the distributions are centered at 0, including the constants (Kilian and Lütkepohl, 2017; Dieppe et al., 2016). Ω_0 is the prior variance-covariance matrix of β. Unlike the case of the Natural Conjugate $\mathcal{NIW}$ priors, here Ω_0 is independent of the residual variance-covariance matrix $\mathbf{\Sigma}$. Moreover, Ω_0 is a diagonal matrix containing the hyperparameters that control the variance of the distributions. Ω_0 takes the following form, where for simplicity we only consider the first lag of the first equation and the constant:

[62] We set the hyperparameters of inverse Wishart distribution in a conventional way: $\alpha_0 = n + 1$ and $S_0 = \mathbb{I}_n$ (Dieppe et al., 2016).

$$\Omega_0 = \begin{bmatrix} \sigma_1^2(\lambda_1\lambda_4)^2 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & \dots \\ 0 & (\lambda_1/L^{\lambda_3})^2 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & \dots \\ 0 & 0 & (\sigma_1/\sigma_2)^2(\lambda_1\lambda_2/L^{\lambda_3})^2 & 0 & 0 & 0 & 0 & 0 & 0 & \dots \\ 0 & 0 & 0 & (\sigma_1/\sigma_3)^2(\lambda_1\lambda_2/L^{\lambda_3})^2 & 0 & 0 & 0 & 0 & 0 & \dots \\ 0 & 0 & 0 & 0 & (\sigma_1/\sigma_4)^2(\lambda_1\lambda_2/L^{\lambda_3})^2 & 0 & 0 & 0 & 0 & \dots \\ 0 & 0 & 0 & 0 & 0 & (\sigma_1/\sigma_5)^2(\lambda_1\lambda_2\lambda_5/L^{\lambda_3})^2 & 0 & 0 & 0 & \dots \\ 0 & 0 & 0 & 0 & 0 & 0 & (\sigma_1/\sigma_6)^2(\lambda_1\lambda_2\lambda_5/L^{\lambda_3})^2 & 0 & 0 & \dots \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & (\sigma_1/\sigma_7)^2(\lambda_1\lambda_2\lambda_5/L^{\lambda_3})^2 & 0 & \dots \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & (\sigma_1/\sigma_8)^2(\lambda_1\lambda_2\lambda_5/L^{\lambda_3})^2 & \dots \\ \ddots & \ddots & \ddots & \ddots & \ddots & \ddots & \ddots & \ddots & \ddots & \ddots \\ \dots & \dots & \dots & \dots & \dots & \dots & \dots & \dots & \dots & \dots \end{bmatrix}$$

The diagonal elements of Ω_0 can be written in the following form:

$$\sigma_{c_i}^2 = \sigma_i^2(\lambda_1\lambda_4)^2 \qquad \text{if constant} \tag{31}$$

$$\sigma_{ii}^2 = (\lambda_1/L^{\lambda_3})^2 \qquad \text{if } i = j \tag{32}$$

$$\sigma_{ij}^2 = (\sigma_i/\sigma_j)^2(\lambda_1\lambda_2/L^{\lambda_3})^2 \qquad \text{if } i \neq j \tag{33}$$

$$\sigma_{for_{ij}}^2 = (\sigma_i/\sigma_j)^2(\lambda_1\lambda_2\lambda_5/L^{\lambda_3})^2 \qquad \text{if } i \neq j \text{ and } n_1 < j \leq n \tag{34}$$

where σ_i^2 and σ_j^2 denote the OLS residual variance of the auto-regressive models estimated for variables i and j. n_1 denotes the number of foreign variables.

Equations (31), (32) and (33) are the standard Minnesota priors. We follow Dieppe et al. (2016) and introduce hyperparamter (λ_5) in Equation (34). λ_5 is applied only on the foreign block and controls the tightness of the distributions of domestic variables in the foreign block. We set $\lambda_5 = 1e-8$. This small value imposes highly informative priors on the parameters of the domestic variables in the foreign block, where the distribution is centered at zero. We select standard prior variances for the rest of the parameters ($\lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = 1$). Thus, we implement block exogeneity through a special case of Independent $\mathcal{NIW}$ priors, where the reduced-form coefficients follow Minnesota priors with an additional hyperparameter. Finally, if the prior distribution over the reduced-form parameters is $\mathcal{NIW}(\alpha_0, \beta_0, S_0, \Omega_0)$, then the posterior distribution over the reduced-form parameters is $\mathcal{NIW}(\tilde{\alpha}, \tilde{\beta}, \tilde{S}, \tilde{\Omega})$, where

$$\begin{aligned}
\tilde{\alpha} &= \mathbf{T} + \alpha_0, \\
\tilde{\Omega} &= [\Omega_0^{-1} + Z'VZ]^{-1}, \\
\tilde{\beta} &= \tilde{\Omega}[\Omega_0^{-1}\beta_0 + Z'V\mathbf{Y}] \\
\tilde{S} &= S_0 + (\mathbf{Y} - \mathbf{X}\hat{\mathbf{B}})'(\mathbf{Y} - \mathbf{X}\hat{\mathbf{B}}),
\end{aligned}$$

where $\mathbf{Y} = [\mathbf{y_1} \cdots \mathbf{y_T}]'$, $\mathbf{X} = [\mathbf{x_1} \cdots \mathbf{x_T}]'$, $V = \mathbf{\Sigma}^{-1} \otimes \mathbb{I}_\mathbf{T}$, $Z = \mathbb{I}_n \otimes \mathbf{X}$ and $\hat{\mathbf{B}} = \tilde{\beta} + chol(\tilde{\Omega})' * RAND((n*p+1)*n, 1)$.

B.3.5 The Importance Sampler

Algorithm 2 draws independently from a distribution over the structural parameterization conditional on the zero restrictions that is not equal to the $\mathcal{NGN}(\alpha,\ \beta,\ S,\ \Omega)$ distribution conditional on the zero restrictions. Since the objective is to independently draw from the $\mathcal{NGN}(\alpha,\ \beta,\ S,\ \Omega)$ distribution over the structural parameterization conditional on the zero restrictions, we employ the following importance sampler algorithm.[63]

Algorithm 3: The following algorithm independently draws from the $\mathcal{NGN}(\alpha,\ \beta,\ S,\ \Omega)$ distribution over the structural parameterization conditional on the sign and zero restrictions.

1. Use Algorithm 2 to independently draw $(\mathbf{A}_0, \mathbf{A}_+)$.
2. If $(\mathbf{A}_0, \mathbf{A}_+)$ satisfies the sign restrictions, then set its importance weight to

$$\frac{\mathcal{NGN}_{(\alpha,\beta,S,\Omega)}(\mathbf{A}_0, \mathbf{A}_+)}{\mathcal{NIW}_{(\alpha,\beta,S,\Omega)}(\mathbf{B}, \mathbf{\Sigma})\ \vartheta_{(g \circ f_{Chol})|\mathcal{Z}}(\mathbf{A}_0, \mathbf{A}_+)} \propto \frac{|\det \mathbf{A}_0|^{-(2n+m+1)}}{\vartheta_{(g \circ f_{Chol})|\mathcal{Z}}(\mathbf{A}_0, \mathbf{A}_+)},$$

where $(\mathbf{B}, \mathbf{\Sigma}, \mathbf{Q}) = f_{Chol}(\mathbf{A}_0, \mathbf{A}_+)$ and $\mathcal{Z}$ denotes the set of all structural parameters that satisfy the zero restrictions. Otherwise, set its importance weight to zero.
3. Return to Step 1 until the required number of draws has been obtained.
4. Re-sample with replacement using the importance weights.

[63] see Arias et al. (2018) for details.

Non-Technical Appendix (Not for Publication)

C Baseline

C.1 IRFs to a one standard deviation SOE monetary policy shocks

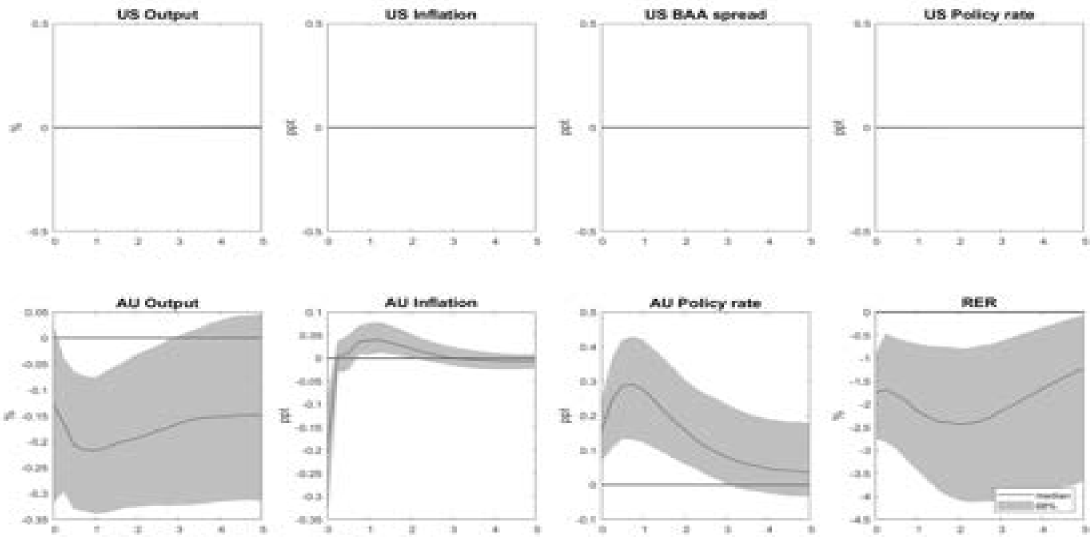

Figure 9: Australia - IRFs to a one standard deviation contractionary monetary policy shock

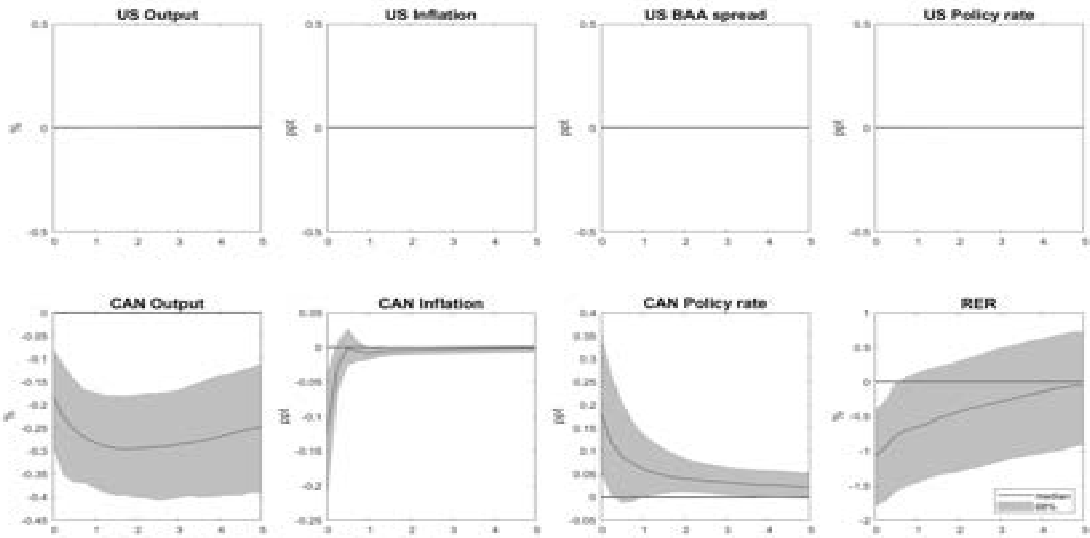

Figure 10: Canada - IRFs to a one standard deviation contractionary monetary policy shock

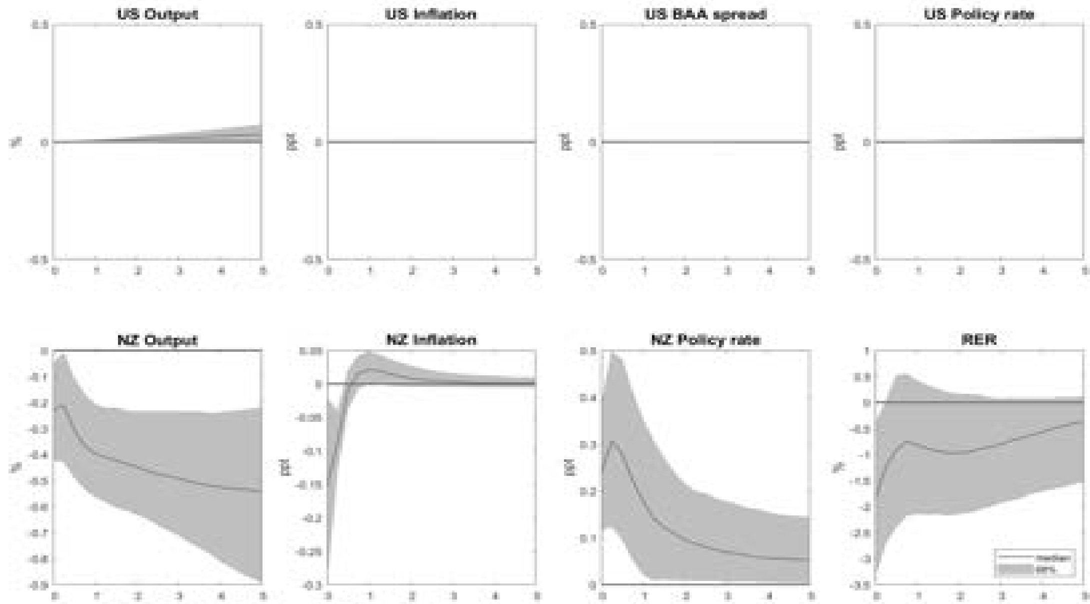

Figure 11: New Zealand - IRFs to a one standard deviation contractionary monetary policy shock

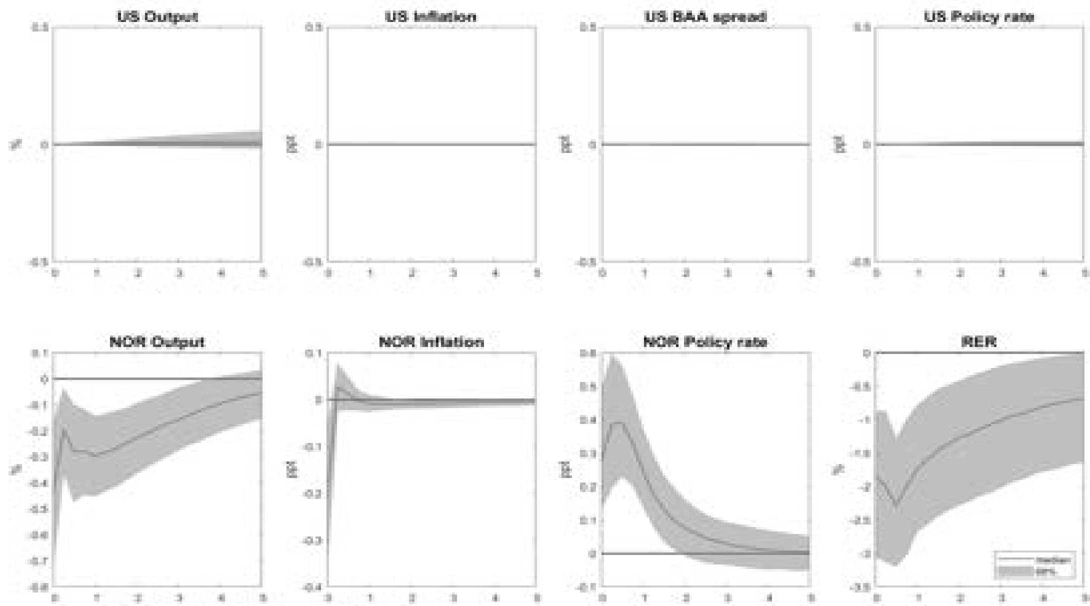

Figure 12: Norway - IRFs to a one standard deviation contractionary monetary policy shock

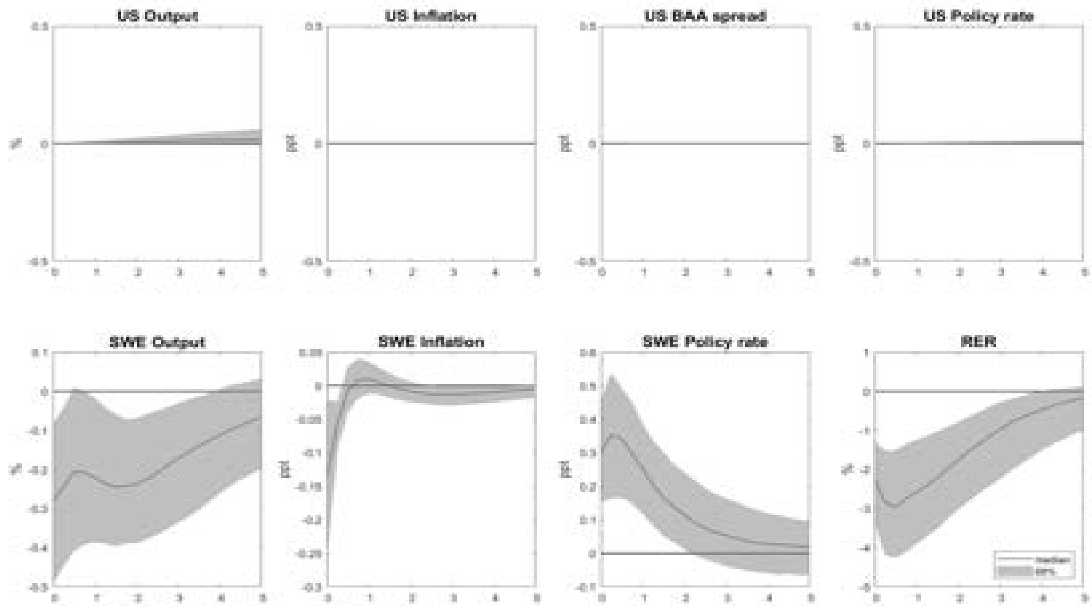

Figure 13: Sweden - IRFs to a one standard deviation contractionary monetary policy shock

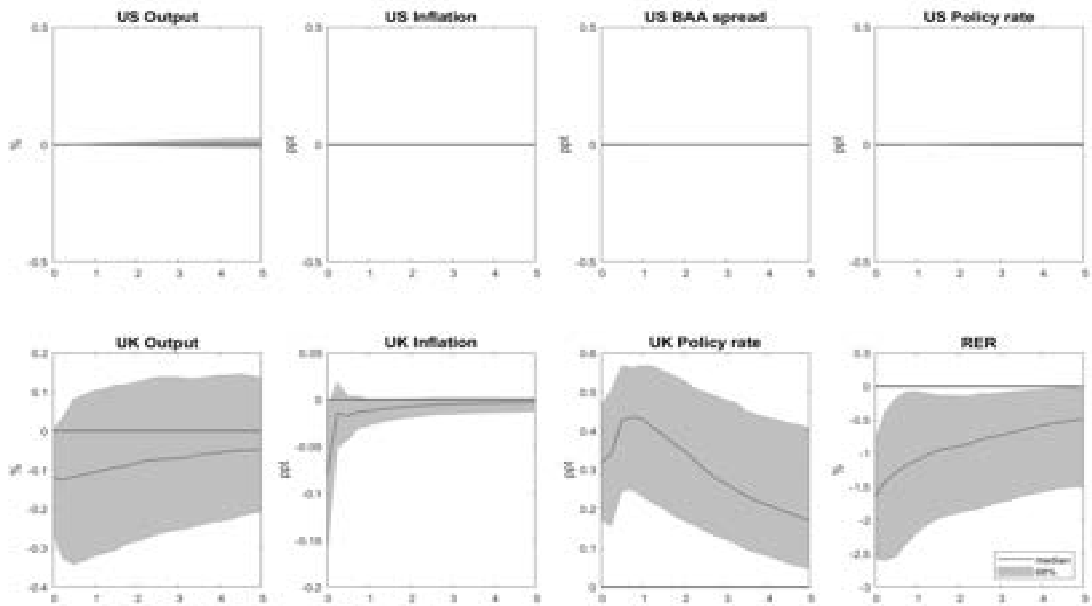

Figure 14: UK - IRFs to a one standard deviation contractionary monetary policy shock

D Baseline

D.1 IRFs to a one standard deviation US monetary policy shocks

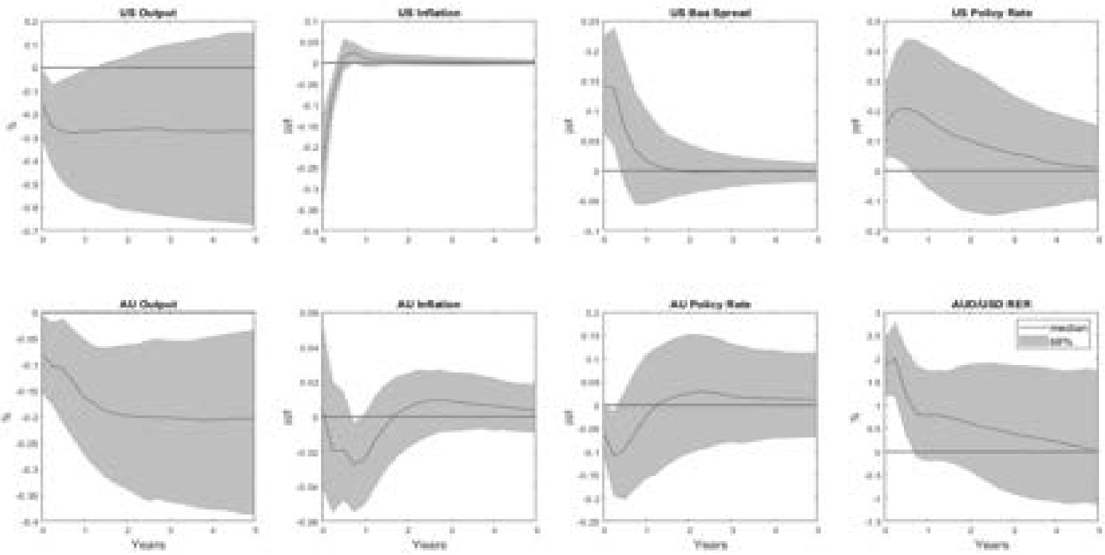

Figure 15: Australia - IRFs to a one standard deviation US contractionary monetary policy shock

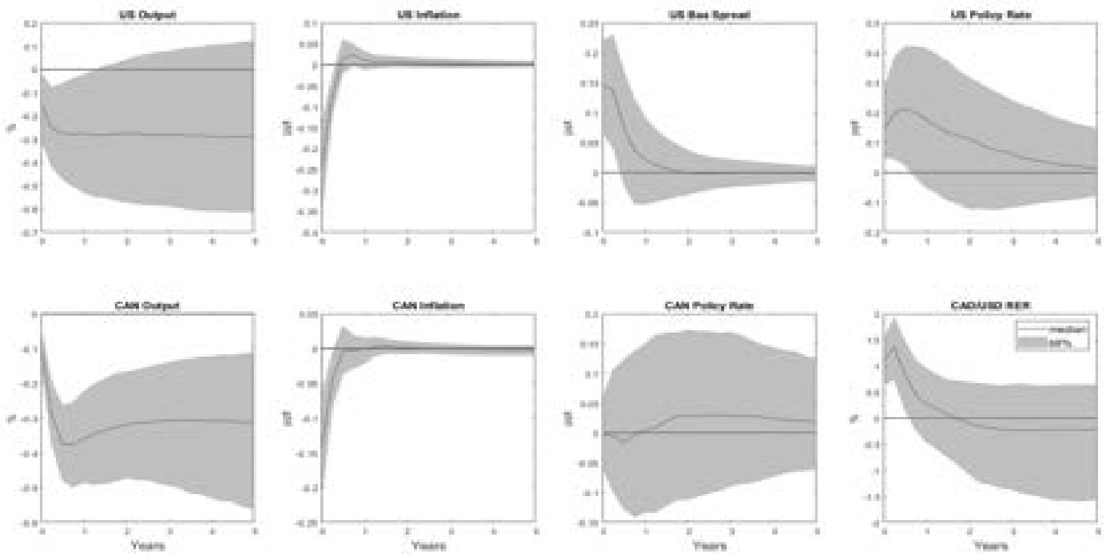

Figure 16: Canada - IRFs to a one standard deviation US contractionary monetary policy shock

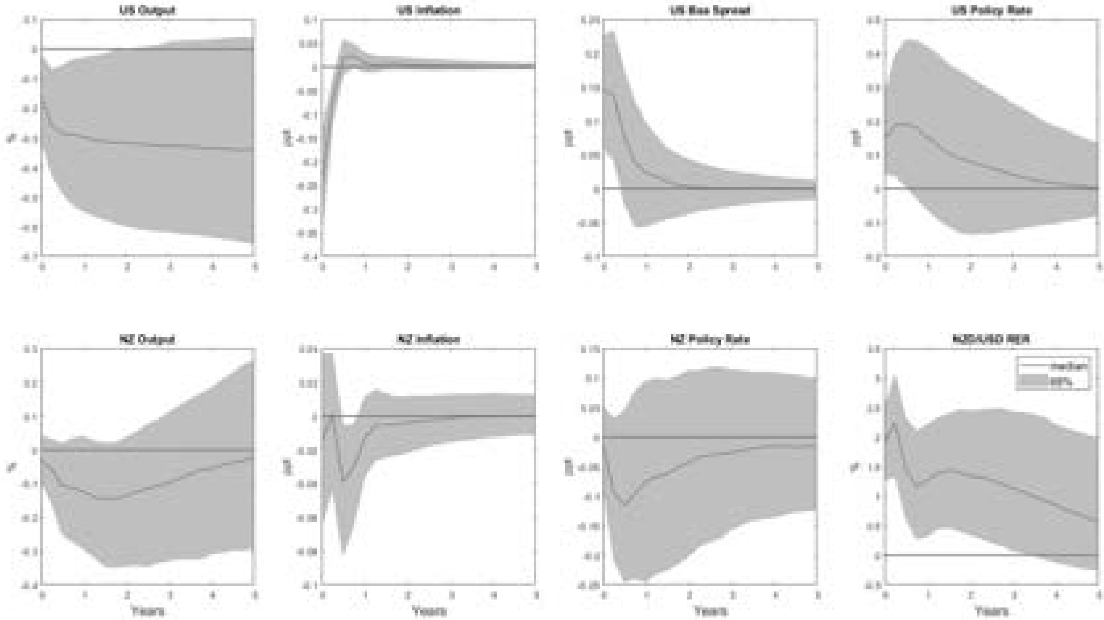

Figure 17: New Zealand - IRFs to a one standard deviation US contractionary monetary policy shock

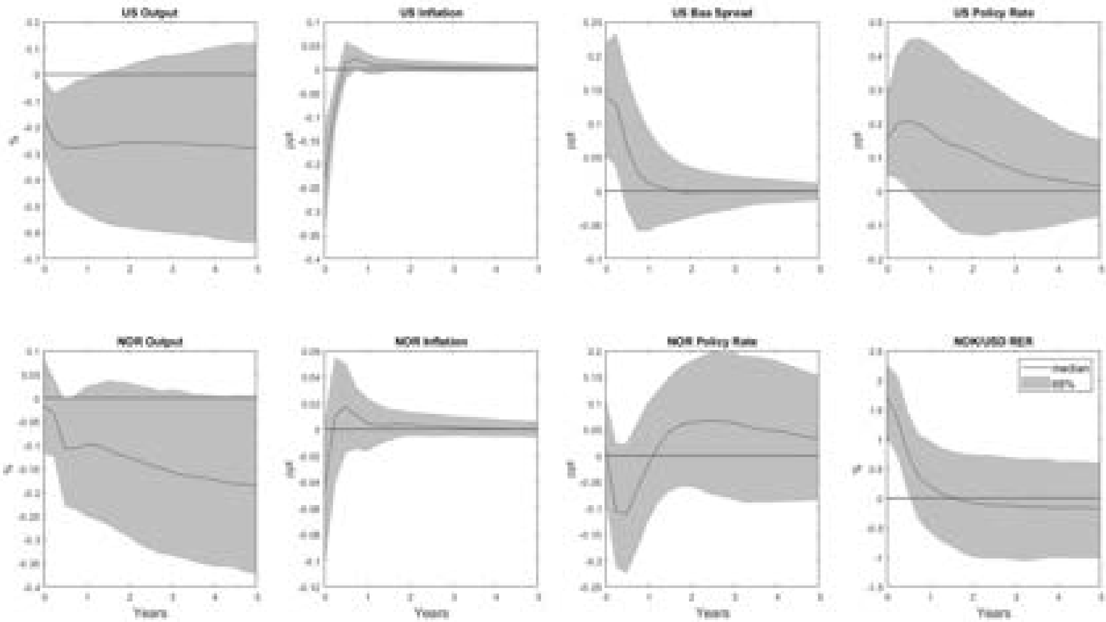

Figure 18: Norway - IRFs to a one standard deviation US contractionary monetary policy shock

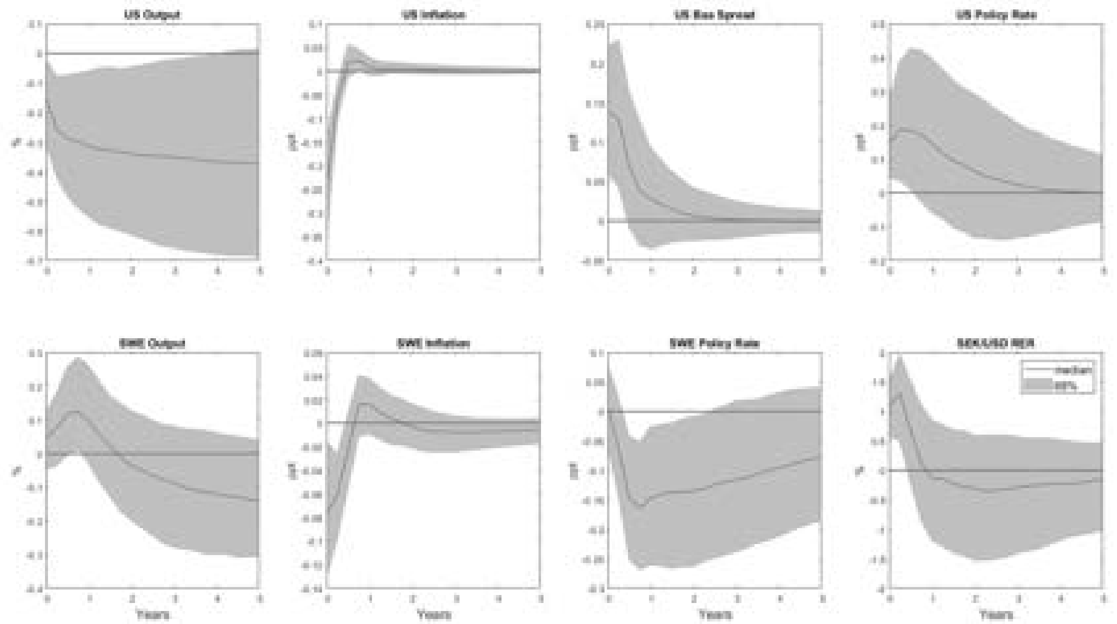

Figure 19: Sweden - IRFs to a one standard deviation US contractionary monetary policy shock

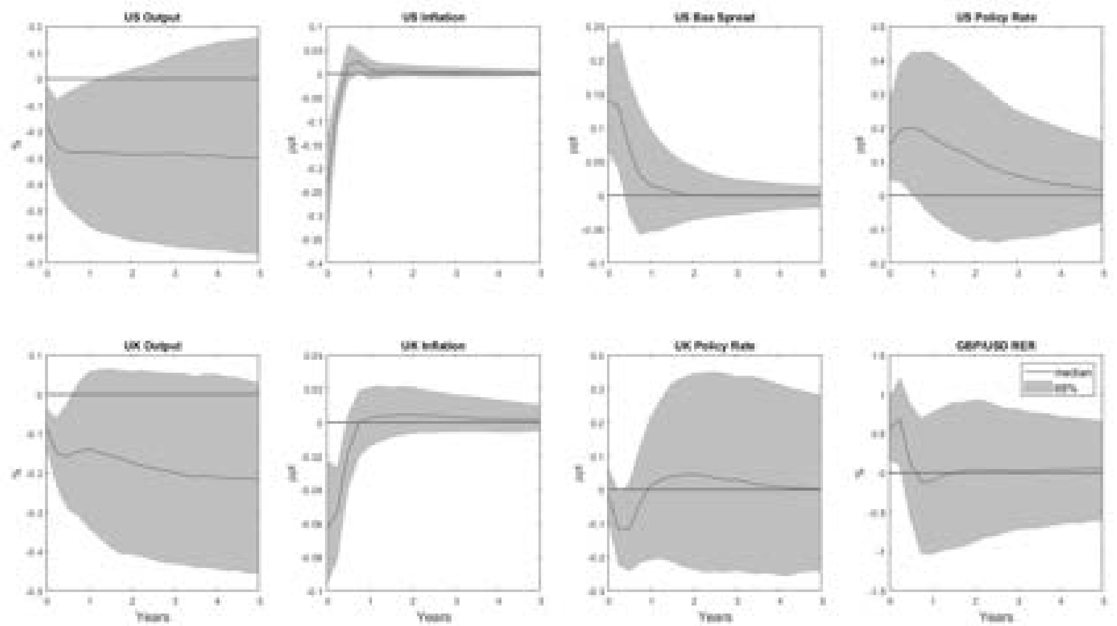

Figure 20: UK - IRFs to a one standard deviation US contractionary monetary policy shock

E Baseline

E.1 Responses of US variables to US shocks

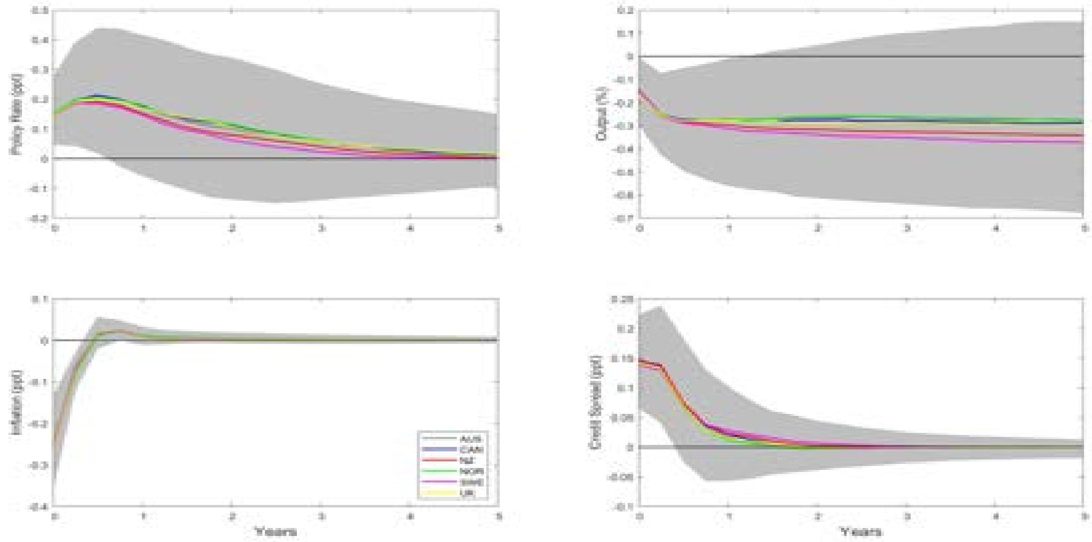

Figure 21: IRFs to a one standard deviation US contractionary monetary policy shock for six SOEs

F Baseline

F.1 IRFs to a one standard deviation US monetary policy shocks with/ out $\psi_e > 0$

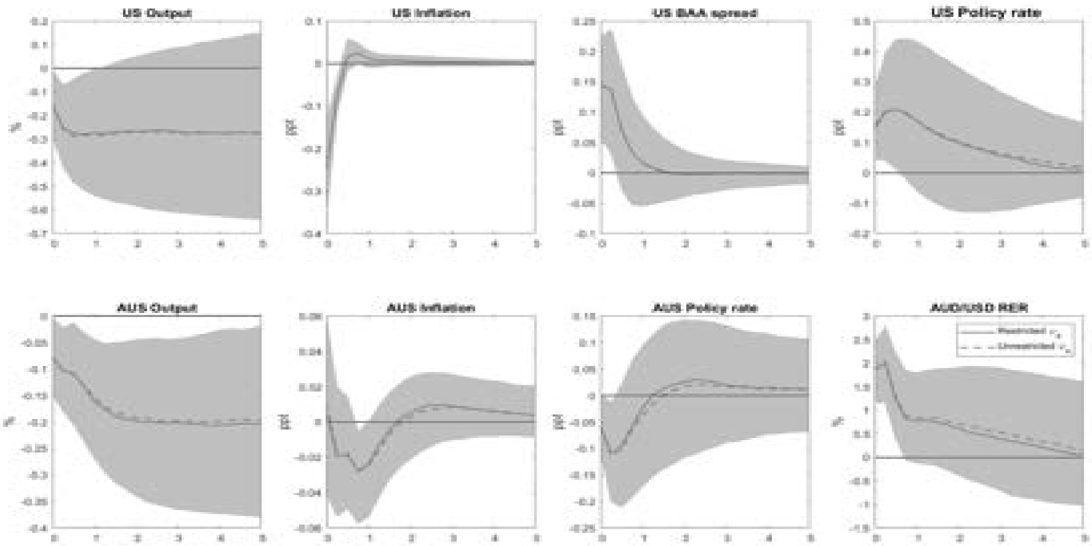

Figure 22: Australia - IRFs to a one standard deviation contractionary monetary policy shock

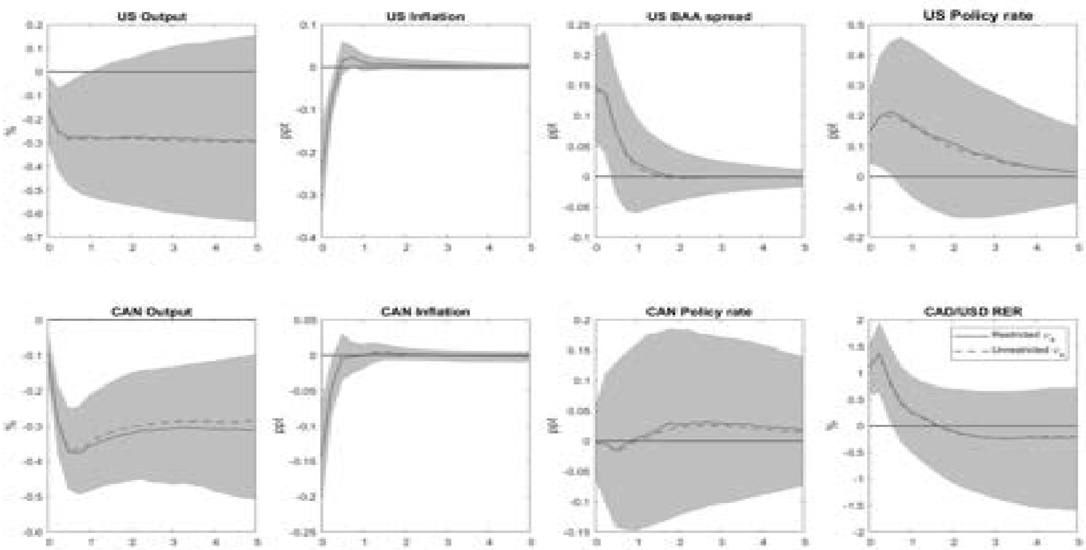

Figure 23: Canada - IRFs to a one standard deviation contractionary monetary policy shock

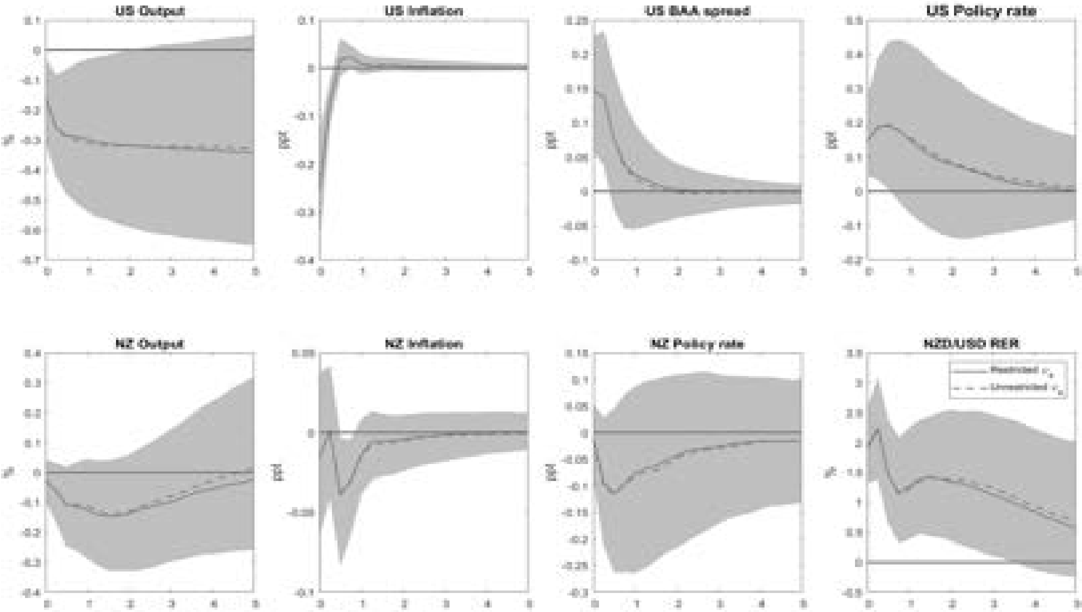

Figure 24: New Zealand - IRFs to a one standard deviation contractionary monetary policy shock

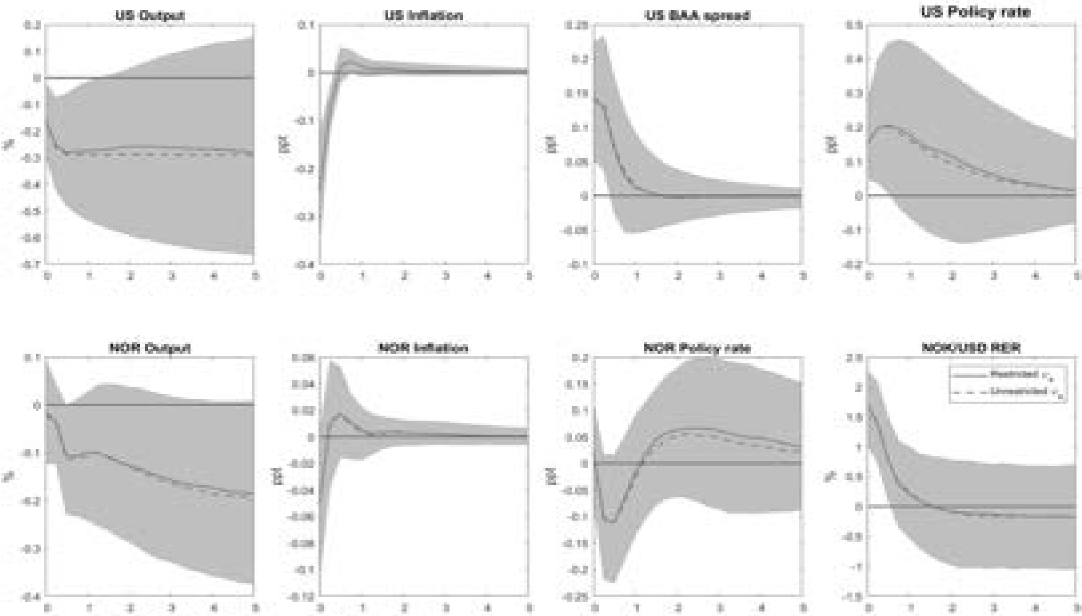

Figure 25: Norway - IRFs to a one standard deviation contractionary monetary policy shock

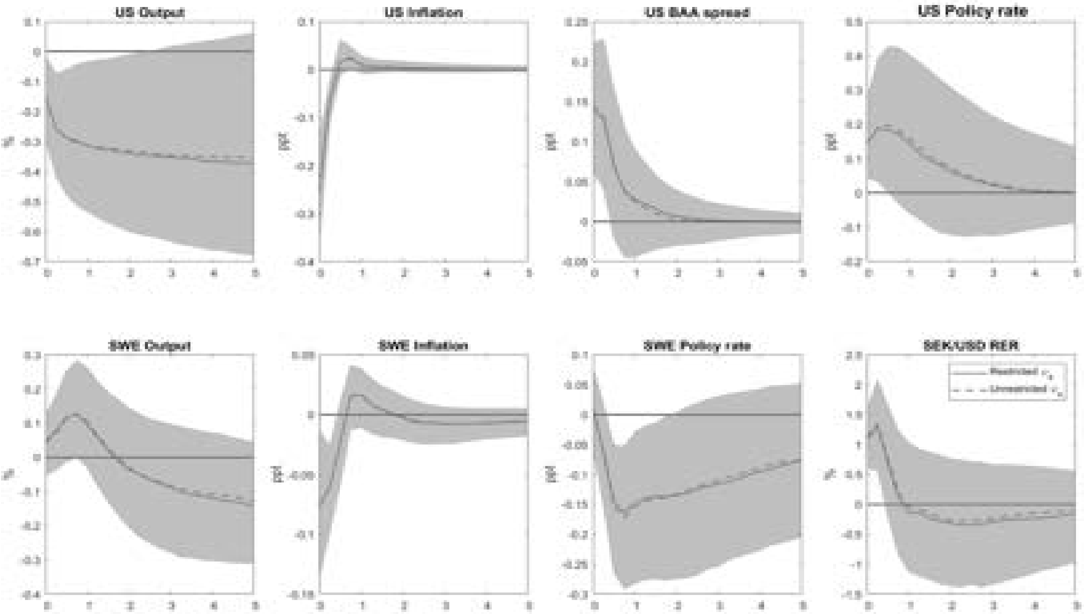

Figure 26: Sweden - IRFs to a one standard deviation contractionary monetary policy shock

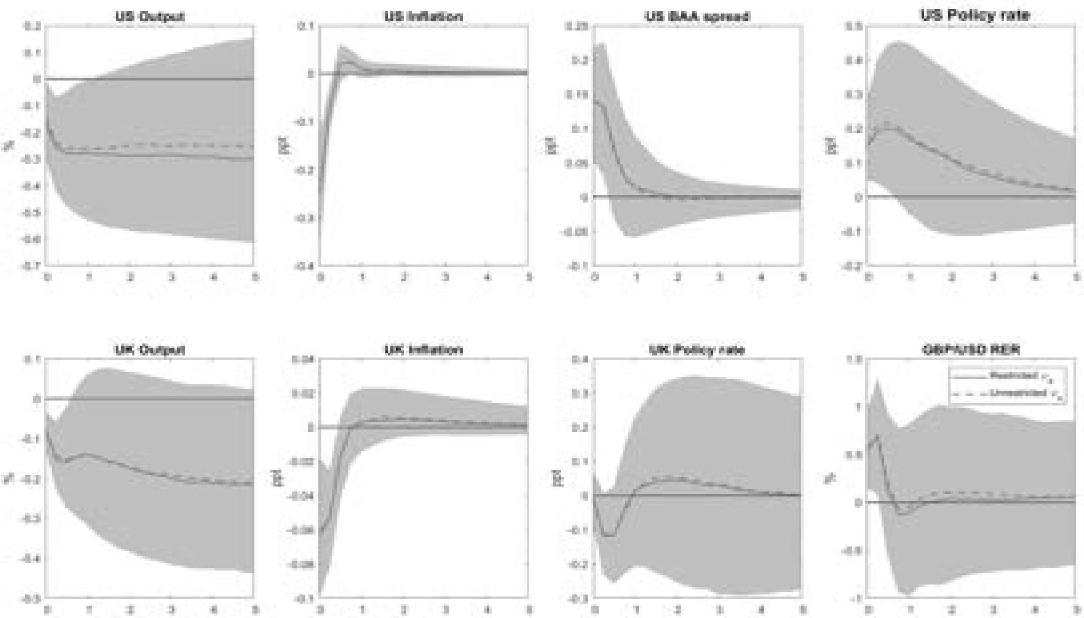

Figure 27: UK - IRFs to a one standard deviation contractionary monetary policy shock

G Baseline

G.1 IRFs to a one standard deviation US monetary policy shocks with/out $\psi_{cs^*} < 0$

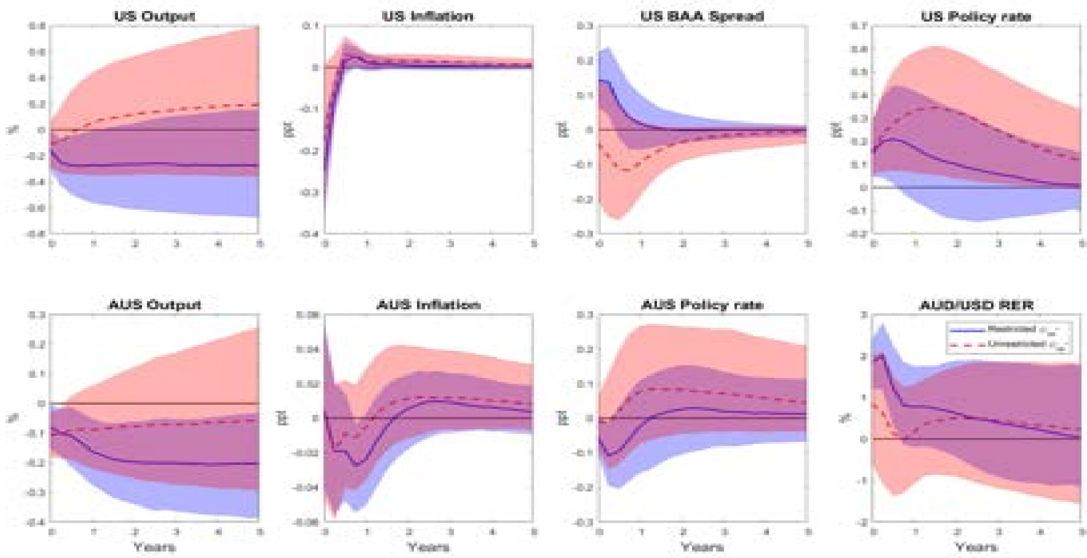

Figure 28: Australia - IRFs to a one standard deviation contractionary monetary policy shock

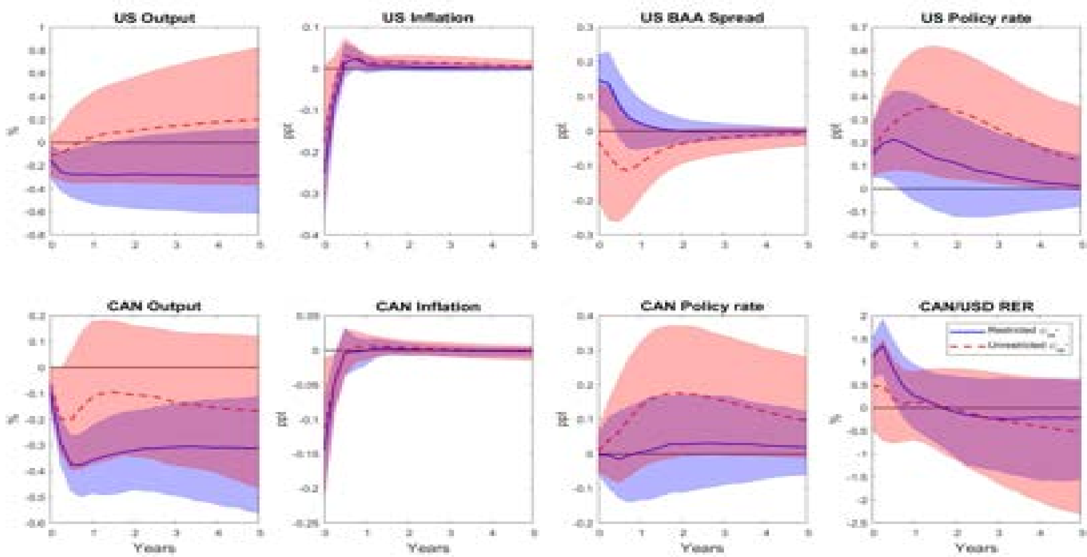

Figure 29: Canada - IRFs to a one standard deviation contractionary monetary policy shock

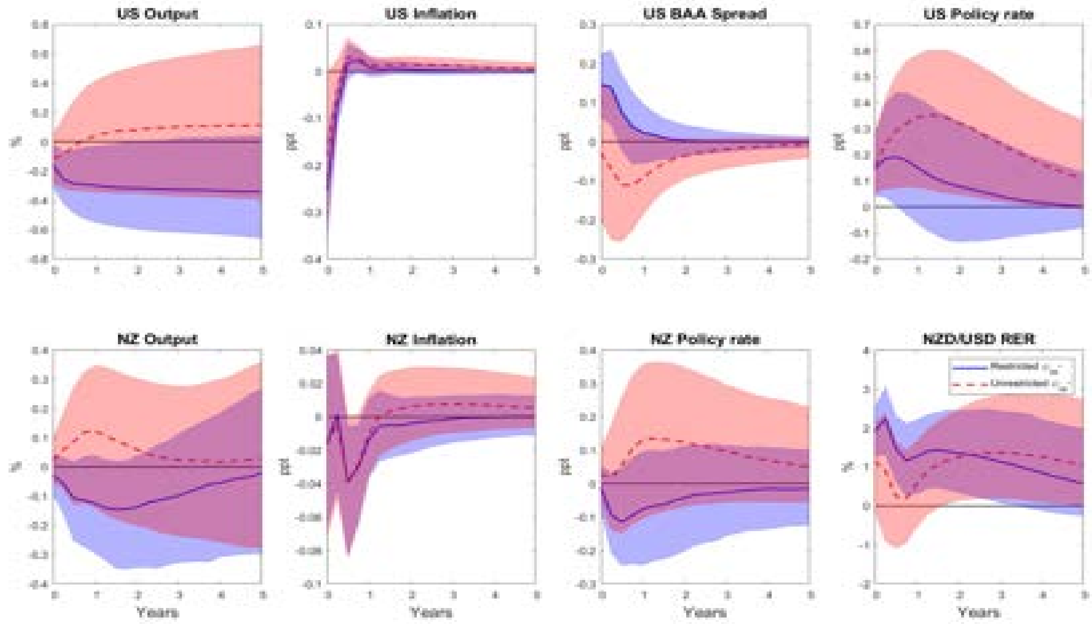

Figure 30: New Zealand - IRFs to a one standard deviation contractionary monetary policy shock

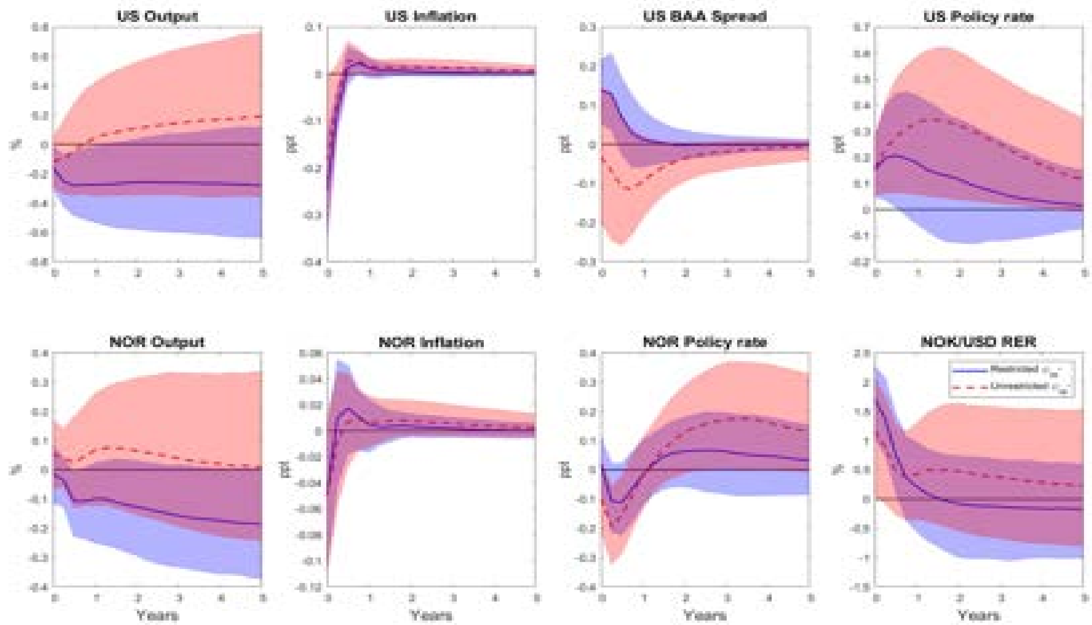

Figure 31: Norway - IRFs to a one standard deviation contractionary monetary policy shock

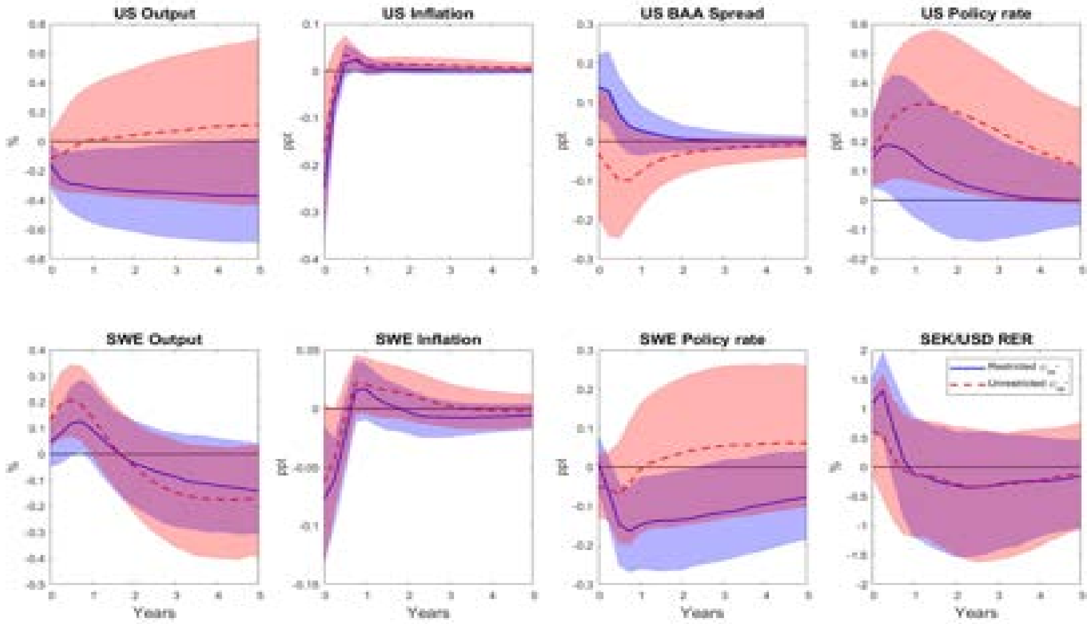

Figure 32: Sweden - IRFs to a one standard deviation contractionary monetary policy shock

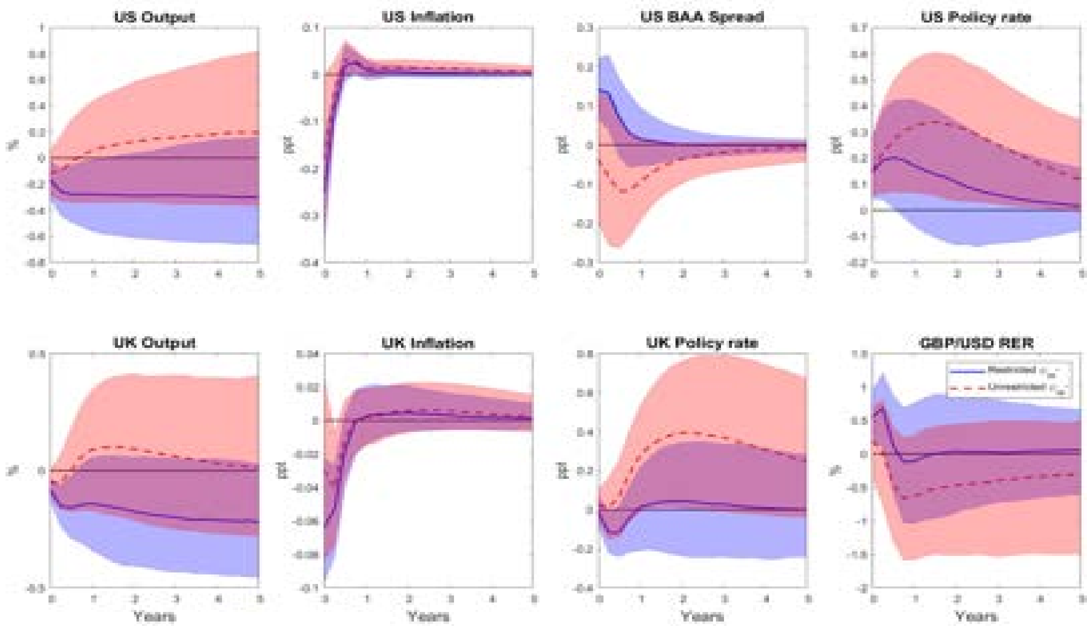

Figure 33: UK - IRFs to a one standard deviation contractionary monetary policy shock

H Split sample in 2008:Q3

H.1 IRFs to a one standard deviation SOE monetary policy shocks (1992:Q1-2008:Q3)

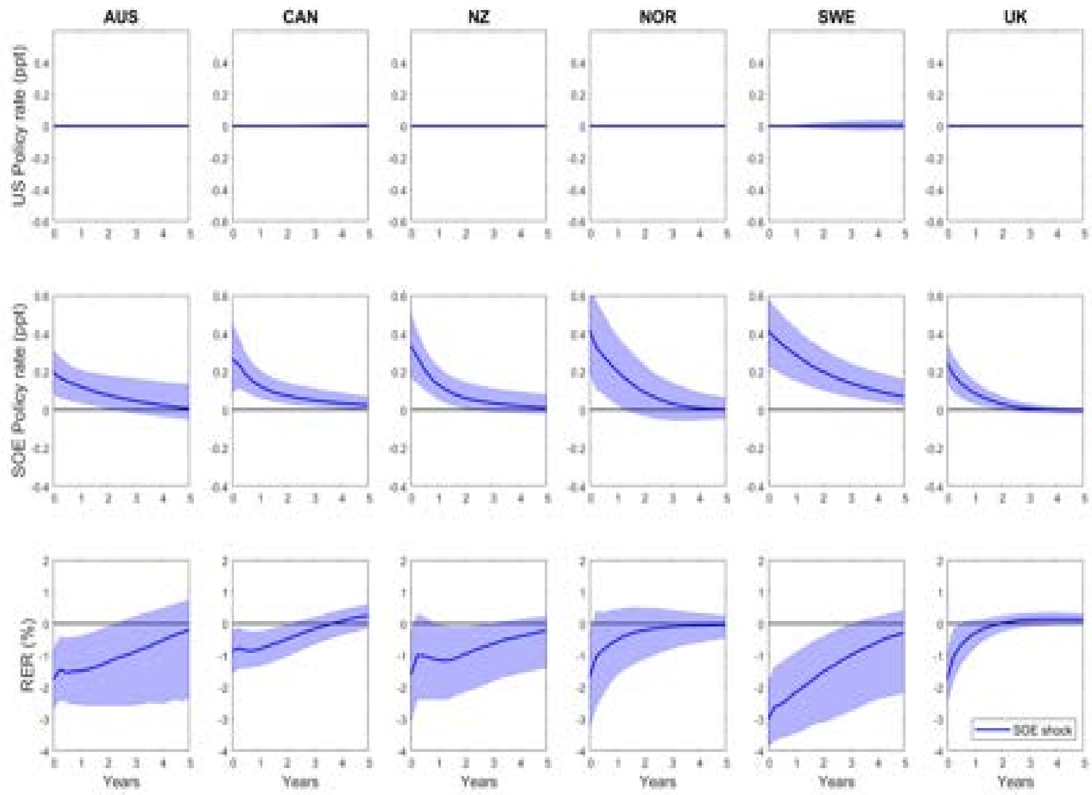

Figure 34: IRFs to a one standard deviation contractionary SOE monetary policy shock for the period 1992:Q1-2008:Q3

H.2 IRFs to a one standard deviation SOE monetary policy shocks (2008:Q4-2019:Q4)

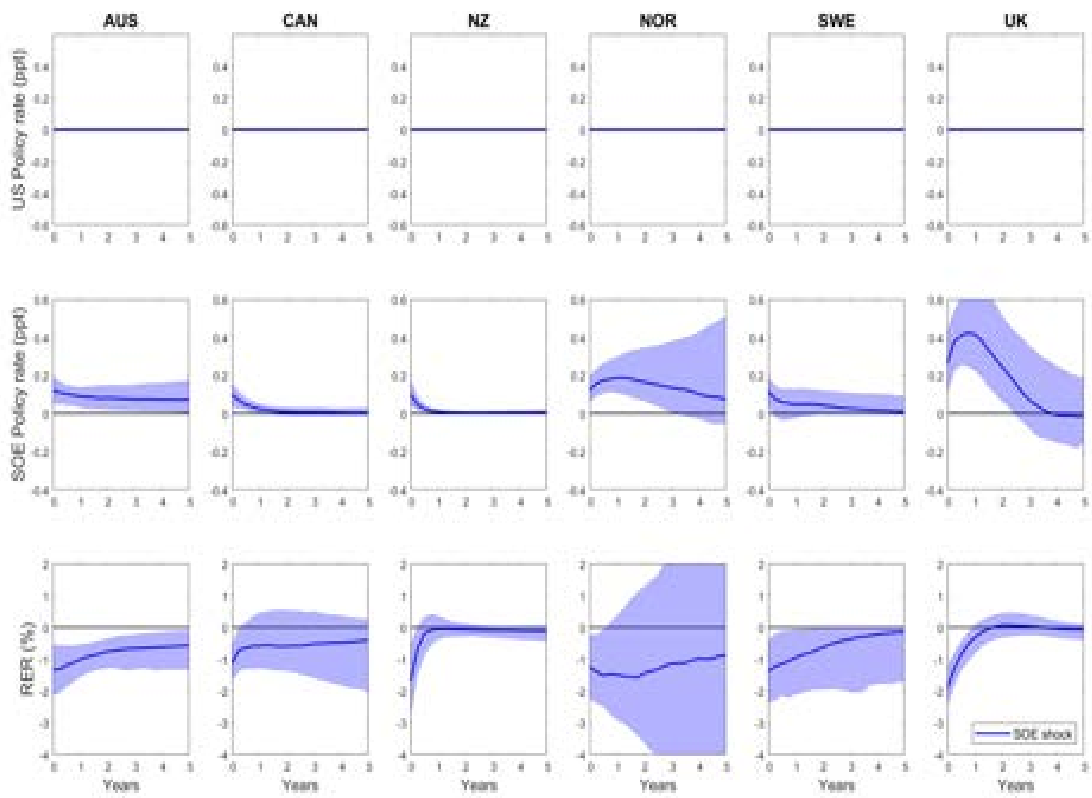

Figure 35: IRFs to a one standard deviation contractionary SOE monetary policy shock for the period 2008:Q4-2019:Q4

H.3 UIP deviations conditional on US and SOE monetary policy shocks for two sub-samples

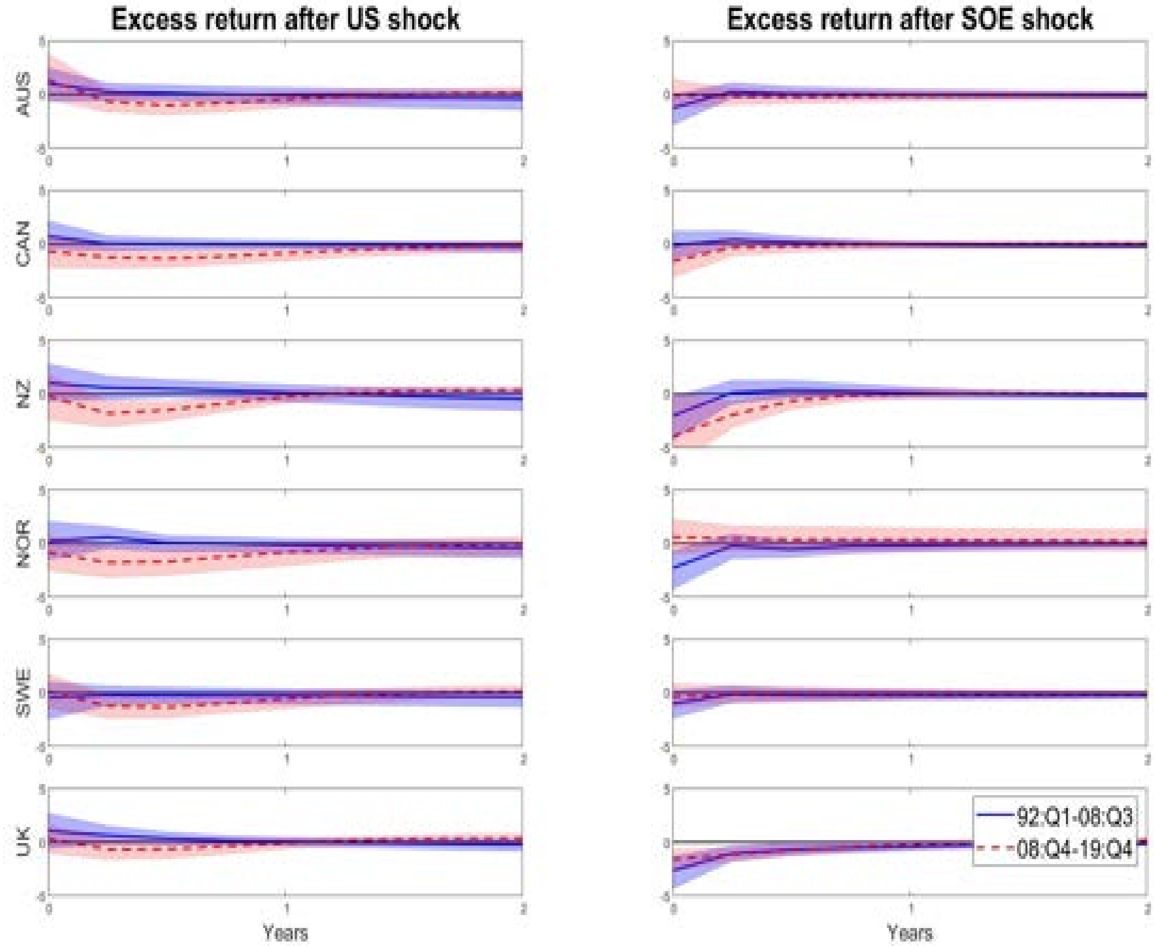

Figure 36: UIP deviations conditional on US and SOE monetary policy shocks for two sub-samples

I Additional results

I.1 IRFs to a one standard deviation SOE monetary policy shocks with/out $\psi_e > 0$ (1992:Q1-2008:Q3)

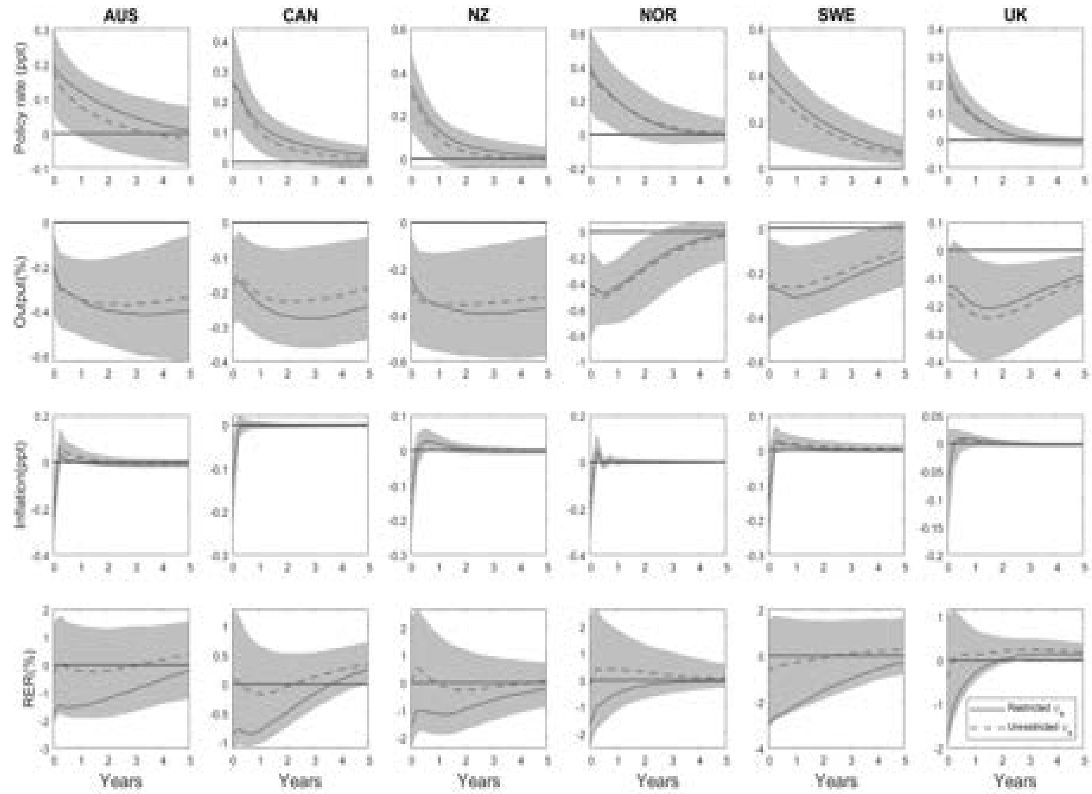

Figure 37: IRFs to a one standard deviation SOE contractionary monetary policy shock with/without $\psi_e > 0$ for sample 1992:Q1-2008:Q3. *Note:* Solid lines are point-wise posterior median responses with Restriction 4 ($\psi_e > 0$). Dashed lines are point-wise posterior median responses after relaxing Restriction 4 with corresponding 68% equal-tailed posterior probability bands.

I.2 Forecast error variance decomposition (1992:Q1-2019:Q4)

Figure 38 shows the forecast-error variance decomposition of the six SOE/US real exchange rates. With the exception of Canada and New Zealand, domestic monetary policy shocks account for a greater share of the RER volatility than US shocks. Depending on the country, domestic monetary policy shocks roughly explain 10 to 25 percents of the volatility of the exchange rate in the short run, while the share attributed to US shocks varies from 3 to 18 percents. Depending on the country, the joint contribution of US and SOE monetary disturbances to the short-run volatility of the SOE/US RER ranges from 25 to 35 percents.

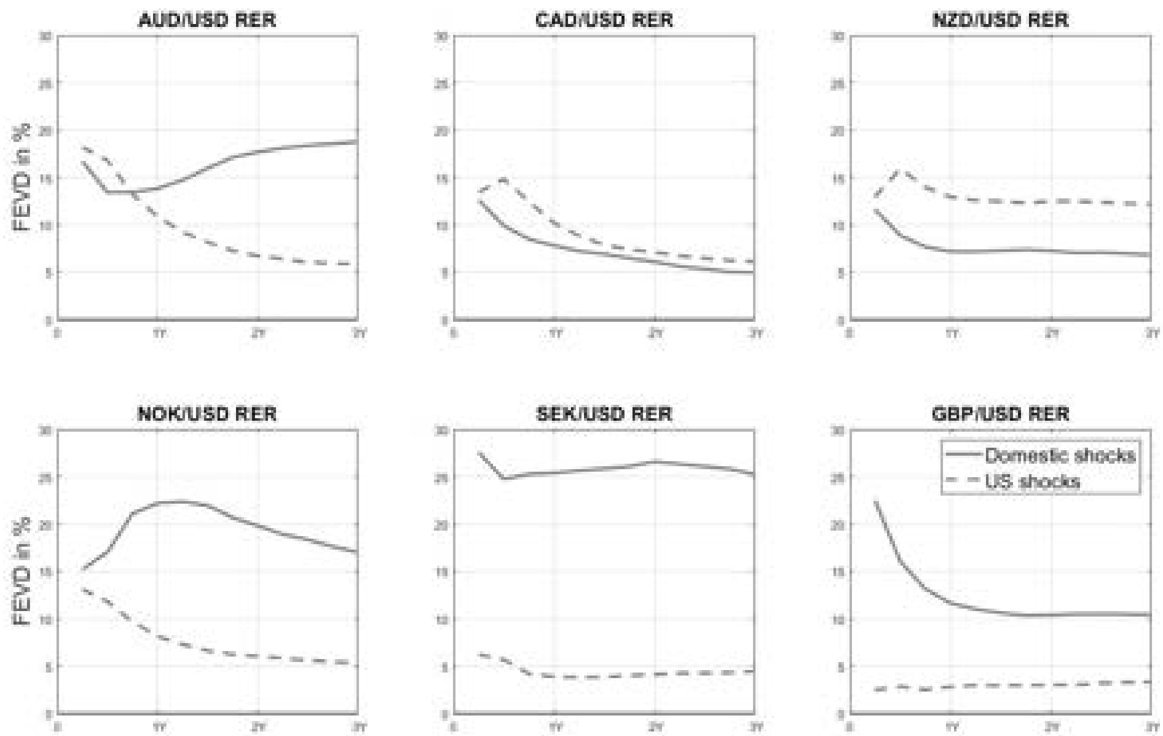

Figure 38: Forecast error variance decomposition of the SOE/US RER. *Note:* In each panel, the solid line represents the posterior median estimate of the contribution of SOE monetary shocks to the forecast-error variance of the RER, while the dashed line shows the contribution of US monetary shocks.

I.3 Robustness check 1: Forecast Error Variance Decomposition (1992:Q1-2008:Q3)

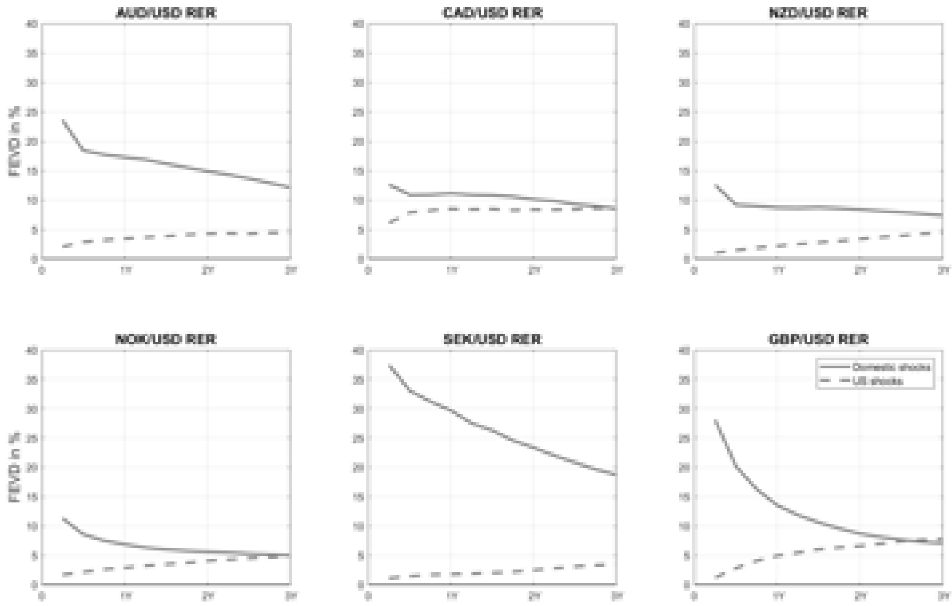

Figure 39: Contribution of one-standard deviation US and domestic monetary policy shocks to time-series fluctuations for Robustness check 1 (Excluding ZLB episodes). *Note:* The solid lines are the contribution of domestic monetary policy shock and dashed lines are the contribution of US monetary policy shock.

I.4 Robustness check 2: Forecast Error Variance Decomposition (1992:Q1-2019:Q4)

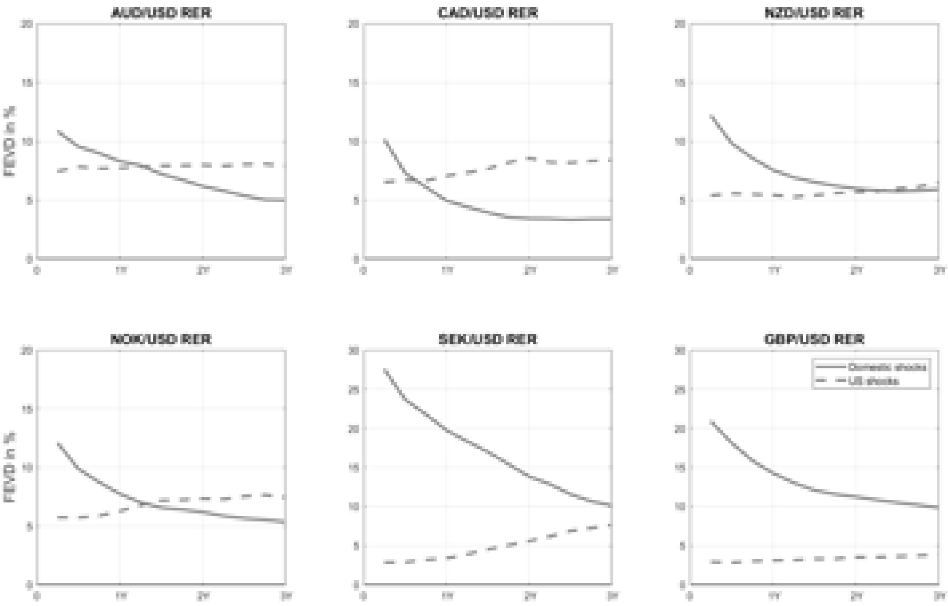

Figure 40: Contribution of one-standard deviation US and domestic monetary policy shocks to time-series fluctuations for Robustness check 2 (SOE Commodity Terms of Trade). *Note:* The solid lines are the contribution of domestic monetary policy shock and dashed lines are the contribution of US monetary policy shock.

J Baseline

J.1 Contemporaneous coefficients in SOE monetary policy equations.

	Coefficients		
SOEs	ψ_y	ψ_π	ψ_e
AUS	0.72	0.65	0.09
	[0.24;2.20]	[0.17;2.75]	[0.02;0.35]
CAN	0.99	1.32	0.16
	[0.24;3.10]	[0.37;3.63]	[0.05;0.49]
NZ	0.87	1.44	0.11
	[0.24;3.99]	[0.36;5.49]	[0.03;0.46]
NOR	0.62	1.31	0.16
	[0.16;2.73]	[0.34;5.80]	[0.04;0.69]
SWE	0.94	1.92	0.14
	[0.21;4.01]	[0.58;8.11]	[0.04;0.64]
UK	1.85	3.64	0.18
	[0.54;7.48]	[0.79;14.28]	[0.05;0.72]

Table 1: Contemporaneous coefficients in SOE monetary policy equations using full sample period from 1992:Q1 to 2019:Q4. *Note:* The entries in the table are the posterior median estimates and the entries in the brackets are the respective 68% probability intervals.

J.2 Contemporaneous coefficients in US monetary policy equations.

	Coefficients		
US/SOEs	ψ_{y^*}	ψ_{π^*}	ψ_{cs^*}
US/AUS	0.41 [0.11;0.87]	0.38 [0.10;1.00]	-0.57 [-1.36;-0.19]
US/CAN	0.40 [0.14;0.90]	0.39 [0.11;1.00]	-0.59 [-1.38;-0.17]
US/NZ	0.41 [0.12;0.99]	0.40 [0.12;0.99]	-0.57 [-1.42;-0.15]
US/NOR	0.43 [0.13;0.91]	0.38 [0.10;0.97]	-0.55 [-1.33;-0.14]
US/SWE	0.42 [0.12;0.98]	0.39 [0.10;1.09]	-0.54 [-1.34;-0.16]
US/UK	0.43 [0.13;0.93]	0.39 [0.11;1.00]	-0.56 [-1.33;-0.15]

Table 2: Contemporaneous coefficients in US monetary policy equations using full sample period from 1992:Q1 to 2019:Q4. *Note:* The entries in the table are the posterior median estimates and the entries in the brackets are the respective 68% probability intervals.

K Robustness check 1 - Sample: 1992:Q1-2008:Q3

K.1 Contemporaneous coefficients in SOE monetary policy equations.

	Coefficients		
SOEs	ψ_y	ψ_π	ψ_e
AUS	0.72 [0.19;2.77]	0.73 [0.16;2.56]	0.11 [0.02;0.45]
CAN	1.16 [0.30;4.00]	1.42 [0.46;4.36]	0.24 [0.06;0.70]
NZ	0.99 [0.29;3.80]	1.67 [0.45;6.31]	0.14 [0.04;0.56]
NOR	0.79 [0.21;3.06]	1.80 [0.46;6.93]	0.20 [0.05;0.79]
SWE	1.19 [0.31;4.65]	2.03 [0.55;7.90]	0.15 [0.03;0.72]
UK	1.21 [0.30;5.12]	2.62 [0.87;9.29]	0.14 [0.03;0.65]

Table 3: Contemporaneous coefficients in SOE monetary policy equations using shorter sample period from 1992:Q1 to 2008:Q3. *Note:* The entries in the table are the posterior median estimates and the entries in the brackets are the respective 68% probability intervals.

K.2 Contemporaneous coefficients in US monetary policy equations.

	Coefficients		
US/SOEs	ψ_{y^*}	ψ_{π^*}	ψ_{cs^*}
US/AUS	0.50 [0.14;1.23]	0.67 [0.18;1.74]	-0.99 [-2.02;-0.29]
US/CAN	0.51 [0.16;1.24]	0.73 [0.20;2.02]	-1.01 [-2.09;-0.31]
US/NZ	0.53 [0.16;1.26]	0.76 [0.20;2.05]	-0.98 [-2.11;-0.31]
US/NOR	0.51 [0.15;1.19]	0.75 [0.21;1.89]	-1.01 [-2.17;-0.30]
US/SWE	0.49 [0.14;1.20]	0.73 [0.20;1.84]	-1.01 [-2.12;-0.34]
US/UK	0.50 [0.15;1.21]	0.72 [0.21;1.88]	-0.96 [-2.12;-0.28]

Table 4: Contemporaneous coefficients in US monetary policy equations using shorter sample period from 1992:Q1 to 2008:Q3. *Note:* The entries in the table are the posterior median estimates and the entries in the brackets are the respective 68% probability intervals.

L Robustness check 2

L.1 Contemporaneous coefficients in SOE monetary policy equations.

	Coefficients		
US/SOEs	ψ_y	ψ_π	ψ_e
US/AUS	0.77 [0.25;2.50]	0.75 [0.20;2.50]	0.11 [0.03;0.43]
US/CAN	0.83 [0.23;2.28]	1.14 [0.31;2.95]	0.16 [0.04;0.44]
US/NZ	0.96 [0.25;3.90]	1.51 [0.38;5.69]	0.14 [0.04;0.62]
US/NOR	0.61 [0.16;2.11]	1.38 [0.38;4.91]	0.17 [0.04;0.69]
US/SWE	0.98 [0.26;3.77]	1.83 [0.47;6.52]	0.12 [0.02;0.58]
US/UK	2.08 [0.61;6.76]	3.32 [0.89;12.47]	0.19 [0.04;0.78]

Table 5: Contemporaneous coefficients in SOE monetary policy equations using full sample from 1992:Q1 to 2019:Q4 for 9-variable SVAR models that include commodity terms of trade. *Note:* The entries in the table are the posterior median estimates and the entries in the brackets are the respective 68% probability intervals.

L.2 Contemporaneous coefficients in US monetary policy equations.

	Coefficients			
SOEs	ψ_{y^*}	ψ_{π^*}	ψ_{cs^*}	ψ_{cp^*}
AUS	0.50 [0.14;1.20]	0.66 [0.17;1.67]	-0.80 [-1.88;-0.24]	-0.71 [-1.82;-0.07]
CAN	0.60 [0.19;1.45]	0.95 [0.25;2.45]	-0.87 [-2.06;-0.28]	-0.94 [-2.29;-0.20]
NZ	0.58 [0.18;1.43]	0.62 [0.16;1.82]	-0.92 [-2.31;-0.24]	-1.32 [-3.92;0.12]
NOR	0.62 [0.21;1.67]	1.13 [0.30;3.30]	-1.04 [-2.59;-0.30]	-0.35 [-0.95;-0.08]
SWE	0.56 [0.15;1.48]	0.77 [0.24;2.15]	-0.85 [-2.20;-0.23]	-1.95 [-5.22;-0.26]
UK	0.60 [0.19;1.47]	0.71 [0.20;2.06]	-0.86 [-2.12;-0.24]	-2.21 [-5.97;-0.20]

Table 6: Contemporaneous coefficients in US monetary policy equations using full sample period from 1992:Q1 to 2019:Q4 for 9-variable SVAR models that include commodity terms of trade. *Note:* The entries in the table are the posterior median estimates and the entries in the brackets are the respective 68% probability intervals.