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Monetary policy shocks and exchange rate dynamics in small open economies*

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Abstract

This paper investigates the relationship between monetary policy shocks and real exchange rates in several small open economies. To that end, we develop a novel identification strategy for time-varying structural vector autoregressions with stochastic volatility. Our approach combines short-run and long-run restrictions to preserve the contemporaneous interaction between the interest rate and the exchange rate. Using this framework, we find that the volatility of monetary policy shocks has substantially decreased in all countries. This leads to a considerable reduction in the significance of policy shocks in explaining exchange rate and macroeconomic fluctuations since the 1990s. However, we find that the dynamic effects of the policy shocks have remained stable over time. Finally, while we do identify violations of uncovered interest parity (UIP) in some countries, we find no evidence of the 'exchange rate puzzle' or the 'delayed overshooting puzzle' in any country.

JEL Classifications: C32, E52, F31, F41.
Keywords: Monetary policy shocks; Exchange rate; Dornbusch overshooting; UIP; TVP-VARs.

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1 Introduction

The well-known Dornbusch's (1976) Exchange Rate Overshooting Hypothesis (henceforth, EROH) postulates an immediate appreciation of the domestic currency following an increase in the domestic interest rate relative to the foreign interest rate, followed by a slow depreciation of the currency. This is in agreement with the Uncovered Interest Parity (UIP) hypothesis, which postulates that interest rate differentials between two countries should equate to the expected change in the exchange rate. Although there has been extensive research on the effects of monetary policy on exchange rates, empirical investigations on the temporal relationship between exchange rate fluctuations and policy shocks are relatively sparse. This study contributes to this literature by investigating the impact of changes in the size and transmission of monetary policy shocks on exchange rate dynamics over time.

There are several reasons to investigate the presence of time variation in the effects of monetary policy shocks on exchange rates. Firstly, exchange rates have exhibited significant fluctuations over the past few decades (Rossi, 2021). Understanding the extent to which these fluctuations are influenced by monetary policy shocks is crucial for policymakers. Secondly, there are various puzzles regarding the impact of monetary policy shocks on exchange rates. For example, some empirical studies have found that a contractionary monetary policy shock often leads to an immediate depreciation of the exchange rate known as the 'exchange rate puzzle' (see, e.g. Grilli and Roubini, 1995), or when the exchange rate appreciates, there is a substantial delay in its peak response, which is often referred to as the 'delayed overshooting puzzle' (see, e.g. Eichenbaum and Evans, 1995). Both puzzles contradict Dornbusch's (1976) EROH. The delayed overshooting puzzle is often called the 'forward discount puzzle' since it implies the violation of the UIP. Thirdly, there is a puzzling empirical fact about UIP: not only do the estimated coefficients in UIP regressions deviate from the values predicted by theory (as shown by Fama, 1984), but they are also unstable over time. Ismailov and Rossi (2018) show that UIP holds during periods of low uncertainty but breaks down during periods of high uncertainty when examining data from several industrialized countries. However, in the context of monetary policy, it is important to determine whether UIP holds conditional on monetary policy shocks and whether this relationship remains stable over time. Kim et al. (2017) demonstrate using U.S. data that UIP fails during the Volcker era but tends to hold in the post-Volcker era, which they attribute to the lack of policy credibility during the Volcker era. They also show that monetary policy shocks have substantial impacts on exchange rate fluctuations but misleadingly appear to have minimal effects when monetary policy regimes are pooled. Kim and Lim (2018) discuss similar issues related to transitions to inflation-targeting regimes in small open economies and sugest that VAR models should be estimated over stable monetary policy regimes to overcome these puzzling results. In general, if there are time variations in the effects and contributions of policy shocks to exchange rate fluctuations, and these variations are overlooked in the empirical analysis, the reported results on the impact and importance of policy shocks may be misleading as they would reflect an average effect over the estimation sample.

To analyze the temporal effects of monetary policy shocks on exchange rates, we employ a time-varying parameter VAR model with stochastic volatility (TVP-VAR-SV), as in Primiceri (2005). This model offers a flexible framework that accommodates the possibility of time-varying parameters (TVPs) and heteroscedastic shocks through stochastic volatility (SV). Following Bjørnland (2009), we seek to identify the monetary policy shock by applying a combination of zero short-run and long-run restrictions. From an economic standpoint, the short-run restrictions establish a recursive structure between macroeconomic variables and the domestic interest rate. This structure ensures that variables like output and inflation do not contemporaneously react to monetary policy shocks, while allowing for simultaneous feedback from macroeconomic variables to domestic monetary policy. On the other hand, the long-run restrictions guarantee that monetary policy shocks do not exert long-run effects on real exchange rates. In this framework, the exchange rate can react immediately to all shocks, but the pass-through effect to macroeconomic variables is slow. However, a significant challenge in identifying monetary policy shocks within the TVP-VAR-SV framework lies in the traditional methods for imposing a combination of zero short-run and long-run restrictions. These methods are not feasible within a TVP-VAR-SV setting (Bjørnland, 2009) because the classic Blanchard and Quah (1988) long-run restrictions do not account for temporal variations in the size or transmission of structural shocks. To address this issue, we adapt an algorithm based on Rubio-Ramirez et al. (2010) (RWZ hereafter) that enables to incorporate zero short-run and long-run restrictions into TVP-VAR-SV models. As such, our methodology overcomes the difficult problem of combining short-run and long-run restriction, and allows us to analyze the temporal dynamics of the effects of monetary policy shocks on exchange rates effectively.

We estimate the TVP-VAR-SV model for six small open economies: Australia, Canada, Norway, New Zealand, Sweden and the United Kingdom over the post-1980s period. Our findings provide some new insights and extension of early results. Firstly, we find evidence of a substantial decrease in the volatility of monetary policy shocks over time in all countries, which support the presence of SV. Secondly, following a contractionary monetary policy shock, the exchange rate appreciates instantaneously in almost all countries and across time. This finding provides strong evidence against the exchange rate puzzle in all countries, which is in contrast with Grilli and Roubini (1995)'s evidence for G-7 countries excluding the US. Thirdly, the delay in overshooting is relatively short and limited. The maximum impact of the policy shock on the exchange rate is only delayed by one quarter in most countries and the additional appreciation following the impact response is quantitatively small. Following the initial appreciation, the exchange rate depreciates back to its long-run level in all countries, which aligns with the EROH. Fourthly, we find evidence of a forward discount puzzle in three out of six countries. Specifically, with the exception of Australia, Canada and New Zealand, monetary policy shock leads to the violation of UIP in other countries, especially in short horizons. Finally, we find that the monetary policy transmission mechanism has been stable over time in most countries, as evidence of time variation in the dynamic effects of monetary policy shocks is weak with impulse responses to an identified policy shock showing limited variations over time. However, we document a substantial decrease in the contribution of monetary policy shocks in driving exchange rate volatility, particularly since the 1990s. This coincides with the adoption of inflation-targeting and central bank independence in most of the countries studied. In addition to the fact that monetary policy shocks are not important in driving exchange rate fluctuations in these countries since the 2000s, the contribution of policy shocks in explaining inflation and output volatility have also fallen over the same time period.

Closely related to our work are studies by Bjørnland (2009), Inoue and Rossi (2019), Rüth (2020), Castelnuovo et al. (2022) and Rüth and Van der Veken (2022). Similar to

Bjørnland (2009), Castelnuovo et al. (2022) and Rüth and Van der Veken (2022) show that the aforementioned puzzles disappear and the effects of monetary policy shocks on exchange rates fall in line with the EROH when using plausible identification methods. Inoue and Rossi (2019) and Rüth (2020) also find evidence in favour of the EROH using high-frequency data. Our findings depart from Bjørnland (2009) in the sense that sizeable deviations from UIP in some countries are evidenced in our case, despite the exchange rate behavior being consistent with the EROH. The latter funding aligns with Scholl and Uhlig (2008) who show evidence of the forward discount puzzle for the U.S. economy even without delayed overshooting. A growing number of studies have used the TVP-VAR setting to establish empirical evidence on the dynamic structure of the economy.[1] However, few have focused on the relationship between monetary policy and the exchange rate. Yang and Zhang (2021) employ a TVP-VAR model to study the effect of monetary policy on the exchange rate using high frequency surprises in financial markets on central bank announcement days. They find that a contractionary monetary policy lead to a large exchange rate appreciation during unconventional monetary policy periods. Mumtaz and Sunder-Plassmann (2013) investigate the evolving dynamics of the real exchange rate using a TVP-VAR model. The findings of the study indicate that there has been a significant increase in the impact of demand and nominal shocks on the real exchange rate since the mid-1980s. However, it is important to note that the study does not specifically focus on identifying monetary policy shocks or address the open-economy puzzles discussed in this paper.

The paper is organized as follows. Section 2 presents the TVP-VAR model with stochastic volatility, describes the data used in the estimation and outlines the identification strategy. Section 3 documents the main results of the paper. Additional robustness checks are performed in Section 4. Finally, Section 5 concludes. Further empirical results are provided in the Appendix.

[1] See, e.g., Canova and Gambetti (2009); Benati and Lubik (2014a,b); Haque and Magnusson (2021); Haque et al. (2021).

2 Framework

2.1 Time-varying parameter VAR with stochastic volatility

We specify a time-varying parameter vector autoregression (TVP-VAR) with stochastic volatility based on the framework of Primiceri (2005):

$$\boldsymbol{Y}_t = c_t + B_{1,t}\boldsymbol{Y}_{t-1} + \ldots + B_{p,t}\boldsymbol{Y}_{t-p} + u_t, \quad t = p+1, \ldots, T, \tag{2.1}$$

where $\boldsymbol{Y}_t$ is a $n \times 1$ vector of observed endogenous variables, c_t is a $n \times 1$ vector of time-varying intercept coefficients, $B_{i,t}, i = 1, \ldots, p$, are $n \times n$ matrices of time-varying coefficients, u_t are normally distributed, zero mean, heteroscedastic unobservable shocks with variance covariance matrix Ω_t. As Ω_t is positive definite for all t, there exist a lower triangular matrix A_t and a positive definite diagonal matrix Σ_t such that $\Omega_t = A_t^{-1}\Sigma_t\Sigma_t' \, A_t^{-1'}$, where Σ_t and A_t are given by

$$\Sigma_t = \begin{pmatrix} \sigma_{1,t} & 0 & \cdots & 0 \\ 0 & \ddots & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \cdots & 0 & \sigma_{n_t} \end{pmatrix}, \quad A = \begin{pmatrix} 1 & 0 & \cdots & 0 \\ a_{21,t} & \ddots & \ddots & \vdots \\ \vdots & \ddots & \ddots & \\ a_{n1,t} & \cdots & a_{nn-1,t} & 1 \end{pmatrix}.$$

Therefore, we can decompose u_t as $u_t = A_t^{-1}\Sigma_t\varepsilon_t$ with $\varepsilon_t \sim N(0, I_n)$, so that equation (2.1) can be expressed as:

$$\boldsymbol{Y}_t = c_t + B_{1,t}\boldsymbol{Y}_{t-1} + ... + B_{p,t}\boldsymbol{Y}_{t-p} + A_t^{-1}\Sigma_t\varepsilon_t. \tag{2.2}$$

Defining $X_t = I_k \otimes (1, \boldsymbol{Y}_{t-1}', \ldots, \boldsymbol{Y}_{t-p}')$, where $\otimes$ denotes the Kronecker product, and letting $\boldsymbol{\beta}_t = [(\text{vec}(c_t))', (\text{vec}(B_{1,t}))', \ldots, (\text{vec}(B_{p,t}))']'$ be a $(n^2p+n) \times 1$ vector with $\text{vec}(\cdot)$ symbolizing the vectorization operator, we can write equation (2.2) as:

$$\boldsymbol{Y}_t = X_t\boldsymbol{\beta}_t + A_t^{-1}\Sigma_t \, \varepsilon_t, \quad t = p+1, \ldots, T. \tag{2.3}$$

Let $\boldsymbol{a}_t$ be the stacked vector of the lower-triangular elements in A_t: $\boldsymbol{a}_t = (a_{21,t}, a_{31,t}, a_{32,t}, a_{41,t}, \ldots, a_{nn-1,t})'$, and define $\boldsymbol{h}_t$ from the vector of the diagonal elements of the matrix Σ_t: $\boldsymbol{h}_t = (\log(\sigma_{1t}^2), \ldots, \log(\sigma_{nt}^2))'$. Following Primiceri (2005), we assume that these time-varying parameters have the following dynamics:

$$\boldsymbol{\beta}_{t+1} = \boldsymbol{\beta}_t + u_{\beta_t}, \ \boldsymbol{a}_{t+1} = \boldsymbol{a}_t + u_{a_t}, \ \boldsymbol{h}_{t+1} = \boldsymbol{h}_t + u_{h_t}, \tag{2.4}$$

$$\begin{pmatrix} \varepsilon_t \\ u_{\beta_t} \\ u_{a_t} \\ u_{h_t} \end{pmatrix} \sim N\left(0, \begin{pmatrix} I_n & 0 & 0 & 0 \\ 0 & \Sigma_\beta & 0 & 0 \\ 0 & 0 & \Sigma_a & 0 \\ 0 & 0 & 0 & \Sigma_h \end{pmatrix}\right)$$

for $t = p+1, \ldots, n$, where $\boldsymbol{\beta}_{p+1} \sim N(\mu_{\beta_0}, \Sigma_{\beta_0})$, $\boldsymbol{a}_{p+1} \sim N(\mu_{a_0}, \Sigma_{a_0})$ and $\boldsymbol{h}_{p+1} \sim N(\mu_{h_0}, \Sigma_{h_0})$ for some initial vectors and matrices of parameters μ_{β_0}, Σ_{β_0}, μ_{a_0}, Σ_{a_0}, μ_{h_0}, and Σ_{h_0}. The variance and covariance structure for the innovations of the time-varying parameters are controlled by the parameters Σ_β, Σ_a and Σ_h. As in Primiceri (2005), we assume that Σ_a is a diagonal matrix. This implies that the contemporaneous relationships between the variables evolve independently across equations but are correlated within equations. For simplicity we also assume that Σ_β and Σ_h are diagonal matrices. As seen, (2.4) postulates a random walk behavior for all $\boldsymbol{a}_t$, $\boldsymbol{\beta}_t$ and $\boldsymbol{h}_t$. The random walk assumption allows us to focus on gradual and permanent shifts in the coefficients and reduces the number of parameters to estimate in the TVP-VAR model (see Primiceri, 2005).

We estimate the model using the standard Bayesian MCMC methods. In particular, we use the MCMC algorithm developed by Nakajima (2011). We use this algorithm as it allows for more efficient joint sampling of the parameters. For the initial states of the parameters, we place relatively uninformative priors. Specifically, we set $\mu_{\beta_0} = \mu_{a_0} = \mu_{h_0} = 0$ and $\Sigma_{\beta_0} = \Sigma_{a_0} = \Sigma_{h_0} = 10 \times I_n$. The corresponding hyper-parameters' priors for each draw i are respectively $(\Sigma_\beta)_i^{-2} \sim Gamma(80, 0.00035)$, $(\Sigma_a)_i^{-2} \sim Gamma(12, 0.01)$, $(\Sigma_h)_i^{-2} \sim Gamma(12, 0.01)$. The relatively tighter hyper-parameter priors for the VAR coefficients, i.e. $(\Sigma_\beta)_i^{-2}$, are chosen such that we can impose a stability constraint on the VAR coefficients. Looser priors for $(\Sigma_\beta)_i^{-2}$ allow for more drifts in the VAR coefficients but can also make the

VAR unstable.[2] To compute the posterior estimates, we collect 5000 *restricted* posterior samples after discarding the initial 1000 as burn-in. Following Cogley and Sargent (2005), our posterior draws are restricted to be comprised of only those that produce stable VAR dynamics at each point in time.[3]

In our empirical application, the vector of dependent variables is $\boldsymbol{Y_t} = [R_t^*, \pi_t, y_t, R_t, \Delta re_t]'$, where R_t^* is the trade-weighted foreign interest rate, π_t is the annual change in the log of consumer prices, i.e., inflation, y_t is the log of real gross domestic product, R_t is the domestic interest rate, and Δre_t is the first difference of the log of the trade-weighted real exchange rate.[4] The model is estimated for six small open economies: Australia, Canada, Norway, New Zealand, Sweden and the United Kingdom. We focus on small open economies as the exchange rate is potentially an important transmission channel for monetary policy. Quarterly data from 1983Q1 to 2019Q4 are used to estimate the model. We quadratically detrend the data before the estimations.[5]

2.2 Identification

Our identification strategy is based on a combination of short-run and long-run restrictions, as in Bjørnland (2009). First, a standard recursive structure is imposed between the macroeconomic variables and the domestic interest rate, ensuring that variables such as output and inflation do not react simultaneously to monetary policy shocks. However, there may be a simultaneous feedback from macroeconomic variables to domestic monetary policy setting. Similar recursive restrictions are also imposed on the relationship between the exchange rate and the macroeconomic variables. The exchange rate can react immediately to all shocks, but there is a slow pass-through of the exchange rate to macroeconomic variables, primarily due to nominal rigidities.

[2] Another reason for choosing tighter priors for the VAR coefficients than those set for the variance and covariance states is because otherwise time variation in the variances of the model may be absorbed by the VAR coefficients, exaggerating the drifts in the systematic relationship between the variables. See Sims's (2001) comment on Cogley and Sargent (2001).

[3] See Appendix B of Cogley and Sargent (2005) for more details on this step.

[4] See the Appendix for a description of the data sources.

[5] Bjørnland (2009) estimates a constant-parameter VAR together with a linear time trend. We choose to quadratically detrend the data instead of linear detrending as we extend the sample. Doing so makes the data comparable to Bjørnland (2009) for the overlapping period (comparison available upon request).

In terms of the ordering of the remaining three variables, the foreign interest rate is placed at the top of the ordering, followed by inflation and output. This assumes that domestic macroeconomic variables will be affected contemporaneously by exogenous foreign monetary policy (which is a plausible assumption of small open economy). However, as discussed in Bjørnland (2009), the responses to a monetary policy shock or an exchange rate shock (which are the last two variables in the ordering) remain invariant regardless of the ordering of the first three variables.

When examining the interaction between monetary policy and the exchange rate, the aforementioned identification assumptions alone are insufficient to separately identify monetary policy and exchange rate shocks. A commonly used identification scheme in the literature is the Cholesky decomposition, which either (i) restricts monetary policy from reacting contemporaneously to an exchange rate shock (Eichenbaum and Evans, 1995), or (ii) restricts the exchange rate from immediately reacting to a monetary policy shock (Marcellino and Favero, 2004; Mojon and Peersman, 2001). However, as discussed in Bjørnland (2009), both restrictions are inconsistent with established theory and central bank practices. Assuming (i) implies that the monetary authority ignores any unexpected changes in exchange rates when making monetary policy decisions, while assumption (ii) is questionable since exchange rates, as asset prices, are known to instantaneously react to monetary policy.

Bjørnland (2009) highlights the importance of allowing for full simultaneity between the interest rate and the exchange rate to avoid bias in identifying the policy shock. Therefore, we do not impose a recursive structure between the domestic interest rate and the real exchange rate. Instead, identification is achieved by assuming that monetary policy shocks do not have long-run effects on real exchange rates.[6] This assumption is a standard neutrality assumption that holds for a large class of monetary models in the literature and is consistent with the EROH. By adopting this approach, the monetary policy shock is uniquely identified, and the identification restrictions permit contemporaneous interaction between monetary policy and the exchange rate.

Implementing these restrictions within a TVP-VAR-SV framework poses a significant challenge as they do not account for temporal variation in the size or transmission of

[6] That is why the exchange rate enters as a first difference so that when long-run restrictions are applied to the first-differenced real exchange rate, the effect on the level of the exchange rate will eventually be zero.

the structural shocks. To address this issue, we extend the Rubio-Ramirez-Waggoner-Zha (RWZ) algorithm to our analysis, thus enabling the combination of short-run and long-run restrictions. To explain the algorithm, we need to introduce some notation. Similar to the RWZ identification scheme, restrictions on the short-run and long-run matrices can be written generally as:

$$f\left(A_t^{-1}\Sigma_t, \mathcal{B}_t\right) = \begin{bmatrix} L_{0,t} \\ --- \\ L_{\infty,t} \end{bmatrix} = \begin{array}{c|ccccc|} & s_1 & s_2 & \dots & s_n \\ v_1 & 0 & \times & \dots & \times \\ v_2 & \times & \times & \dots & \times \\ \vdots & \vdots & \vdots & \vdots & \vdots \\ v_n & \times & \times & \dots & \times \\ --- & --- & --- & --- & --- \\ v_1 & \times & \times & \dots & \times \\ v_2 & \times & \times & \dots & \times \\ \vdots & \vdots & \vdots & \vdots & \vdots \\ v_n & \times & \times & \dots & \times \end{array} \tag{2.5}$$

where $\mathcal{B}_t = \sum_{i=1}^{p} B_{i,t}$, $L_{0,t}$ is the $n \times n$ short-run impact matrix defined by $L_{0,t} = A_t^{-1}\Sigma_t$, and $L_{\infty,t}$ is the $n \times n$ long-run impact matrix given by $L_{\infty,t} = \left(I - \mathcal{B}_t\right)^{-1} L_{0,t}$. The labels on the rows denote variables where the ith variable is denoted by v_i, while the labels on the columns denote shocks such that s_i is the ith shock. Moreover, the symbol "$\times$" means no restriction is imposed while "0" denotes an identification restriction that a shock has no short-run and/or no long-run effect on a variable.

The zero restrictions on each shock can be written in terms of a $n \times 2n$ matrix Q_j as:

$$Q_j f\left(A_t^{-1}\Sigma_t, \mathcal{B}_t\right) e_j = 0, \tag{2.6}$$

where e_j is the jth column of the $n \times n$ identity matrix. Given n shocks, there will be n different Q_j matrices. We order the columns in $f\left(A_t^{-1}\Sigma_t, \mathcal{B}_t\right)$ in descending order of the ranks for the corresponding Q_j matrix, where the rank is denoted by $q_j = rank\left(Q_j\right)$. RWZ show that an SVAR with restrictions as in $f(\cdot,\cdot)$ is exactly identified if and only if $q_j = n - j$ for $1 \leq j \leq n$. When a model is exactly identified, one can use the algorithm for exactly identified models to find the unique solution.

The algorithm is described as follows. For a given short-run and long-run matrix $f(\cdot,\cdot)$, find an orthogonal matrix that rotates the initial candidate matrix until it satisfies the identification restrictions. Let us denote the initial short-run matrix as $L^*_{0,t}$ and the initial long-run matrix as $L^*_{\infty,t}$. One potential convenient candidate for the initial impact matrix is the lower Cholesky decomposition of the covariance matrix Ω_t, i.e.,

$$Z_t = chol\left(\Omega_t\right)', \qquad Z_t Z_t' = \Omega_t,$$

such that the initial short-run matrix is given by $L^*_{0,t} = Z_t$, and the initial long-run matrix consistent with this short-run matrix is given by $L^*_{\infty,t} = \left[I - \mathcal{B}_t\right]^{-1} Z_t$. The initial matrix of short-run and long-run responses consistent with the Cholesky decomposition is then given by

$$F_t = \begin{bmatrix} L^*_{0,t} \\ L^*_{\infty,t} \end{bmatrix}.$$

Next, find an orthogonal rotation matrix P_t such that F_t is consistent with the restrictions imposed by (2.5)

$$F_t P_t = f\left(A_t^{-1}\Sigma_t, \mathcal{B}_t\right), \quad where \quad P_t P_t' = I$$

which implies that

$$L_{0,t} = L^*_{0,t} P_t, \quad L_{\infty,t} = L^*_{\infty,t} P_t, \quad L^*_{0,t} P_t P_t' \left(L^*_{0,t}\right)' = \Omega_t.$$

The combination of short-run and long-run restrictions explained above implies the following form of the matrix $f\left(A_t^{-1}\Sigma_t, \mathcal{B}_t\right)$:

$$
f\left(A_t^{-1}\Sigma_t, \mathcal{B}_t\right) = \begin{bmatrix} L_{0,t} \\ --- \\ L_{\infty,t} \end{bmatrix} = \begin{array}{c} \\ R_t^* \\ \pi_t \\ y_t \\ R_t \\ \Delta re_t \\ --- \\ R_t^* \\ \pi_t \\ y_t \\ R_t \\ \Delta re_t \end{array} \begin{array}{c} \begin{array}{ccccc} s_1 & s_2 & s_3 & MP & s_5 \end{array} \\ \begin{bmatrix} \times & 0 & 0 & 0 & 0 \\ \times & \times & 0 & 0 & 0 \\ \times & \times & \times & 0 & 0 \\ \times & \times & \times & \times & \times \\ \times & \times & \times & \times & \times \\ --- & --- & --- & --- & --- \\ \times & \times & \times & \times & \times \\ \times & \times & \times & \times & \times \\ \times & \times & \times & \times & \times \\ \times & \times & \times & \times & \times \\ \times & \times & \times & 0 & \times \end{bmatrix} \end{array} \tag{2.7}
$$

where MP represents the monetary policy shock, which we are interested in. Given these restrictions, one can verify that the model is exactly identified.

The novelty of our application lies in implementing the RWZ algorithm to impose zero short-run and long-run restrictions within a TVP-VAR-SV setting. Previous studies, such as RWZ, have highlighted that obtaining accurate small-sample inferences in time-varying structural VARs using conventional methods is prohibitively expensive and practically infeasible. For instance, methods like Gali (1992) involve solving a system of nonlinear equations for each draw of the parameters at every time point, leading to computational challenges that quickly become infeasible. In contrast, RWZ demonstrate that the specific structure of restrictions in exactly identified systems can be leveraged to develop an efficient method for obtaining the orthogonal matrix P_t.[7]

[7] RWZ also derive computationally efficient algorithm for sign restrictions. See Binning (2013) for an algorithm combining zero short-run and long-run restrictions with sign restrictions.

3 Empirical results

The model is estimated using a lag length of three for Australia, Canada, New Zealand, Norway, Sweden, and the United Kingdom, following the approach employed by Bjørnland (2009).[8] Figure 1 displays the median estimates (solid lines) and the 68% intervals (gray shaded areas) of the posterior distributions of the standard deviations of the estimated monetary policy shocks for each of the six countries. The figure reveals a significant decrease in the volatility of monetary policy shocks across all countries over time, with policy shocks becoming almost negligible since the 2000s. This decline in volatility, coupled with the adoption of inflation-targeting, provides evidence of improved monetary policy, indicating that policy has become less idiosyncratic over time. This finding underscores the importance of incorporating stochastic volatility in the estimated VAR. As discussed later in Section 3.3, a direct implication of this finding is that the impact of monetary policy shocks on explaining exchange rate and macroeconomic volatility has also substantially diminished over time.

3.1 The Dornbusch overshooting hypothesis

In this section, we analyze the dynamics of policy shocks and examine their relationship with the Dornbusch EROH. Figures 2-7 present the posterior median impulse responses of inflation, output, domestic interest rate, and the level of the real exchange rate to a contractionary monetary policy shock for each of the six countries, respectively. To ensure comparability across different time periods, we adopt the approach suggested by Nakajima (2011) and calculate the impulse responses by setting the initial shock size equal to the average stochastic volatility observed over the sample period.[9]

[8] Our results remain essentially unchanged when using two lags instead as in Cogley and Sargent (2005); Primiceri (2005).

[9] The impact response of the domestic interest rate nonetheless shows time variation due to the long-run restriction imposed, which, in turn, relies on accumulated responses over time.

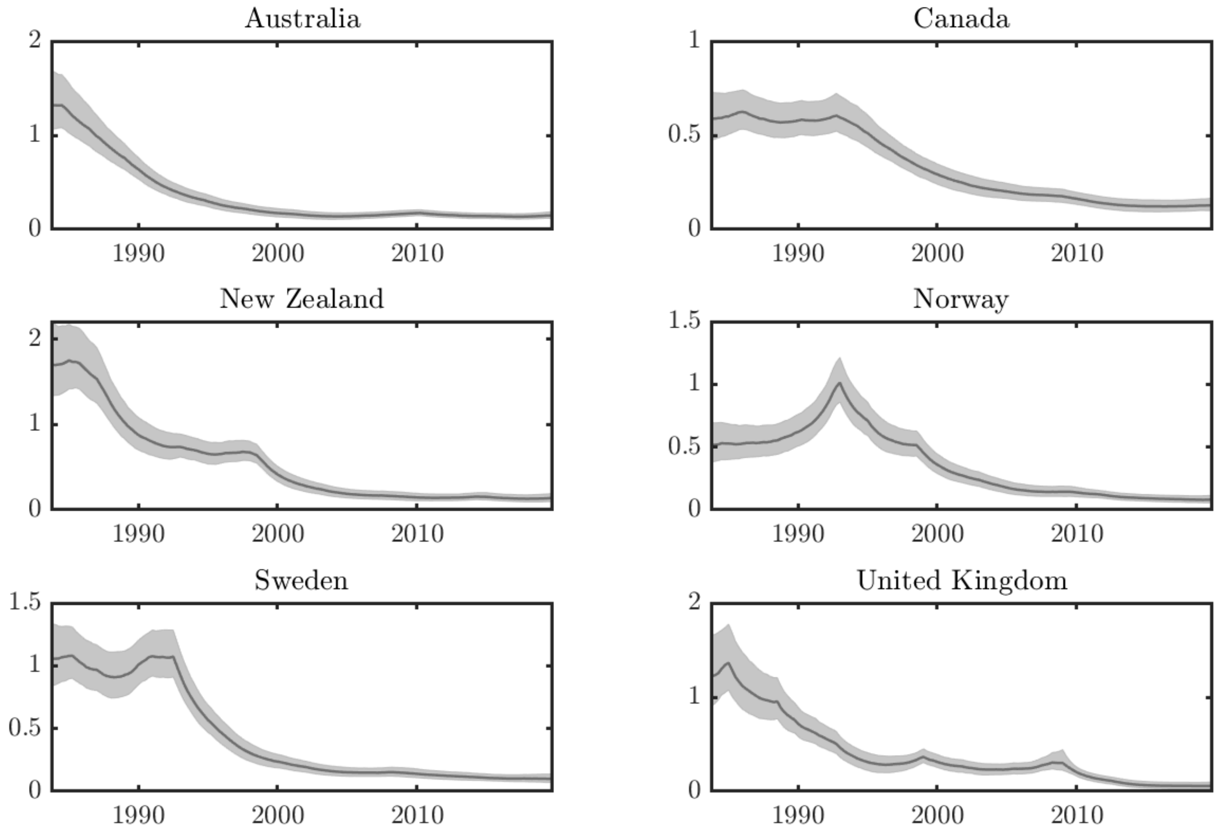

Figure 1: Estimated standard deviations of monetary policy shocks. Solid lines depict the posterior median estimates from the TVP-VAR-SV model while the gray shaded area represents 68% posterior credible intervals around the posterior median.

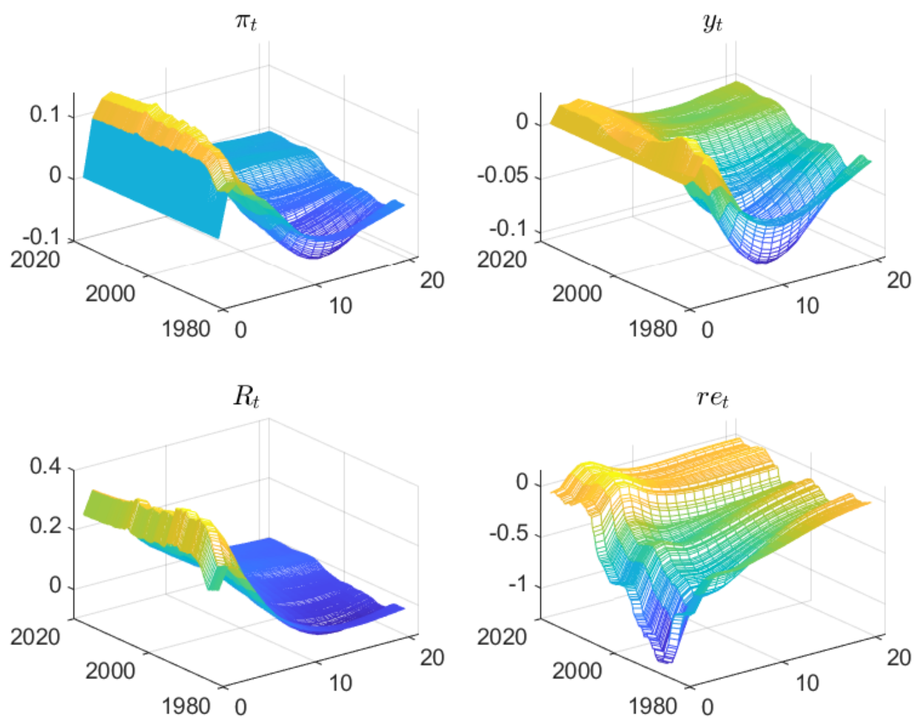

Figure 2: Australia – median impulse responses to a monetary policy shock from the TVP-VAR-SV model.

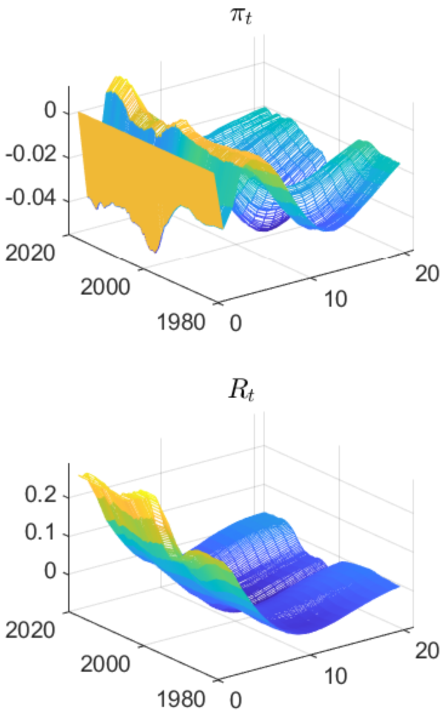

Figure 3: Canada – median impulse responses to a monetary policy shock from the TVP-VAR-SV model.

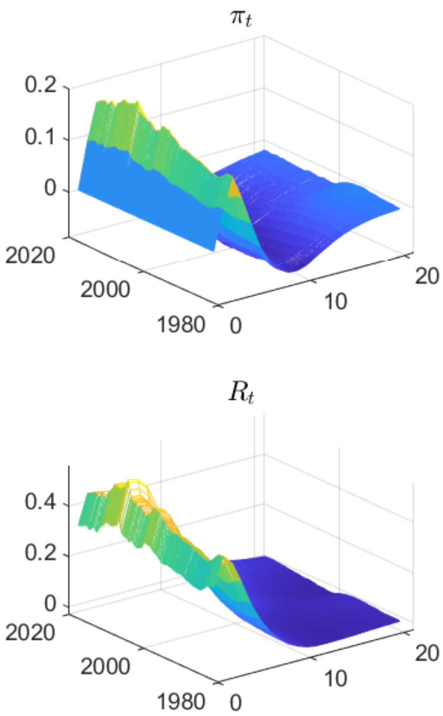

Figure 4: New Zealand – median impulse responses to a monetary policy shock from the TVP-VAR-SV model.

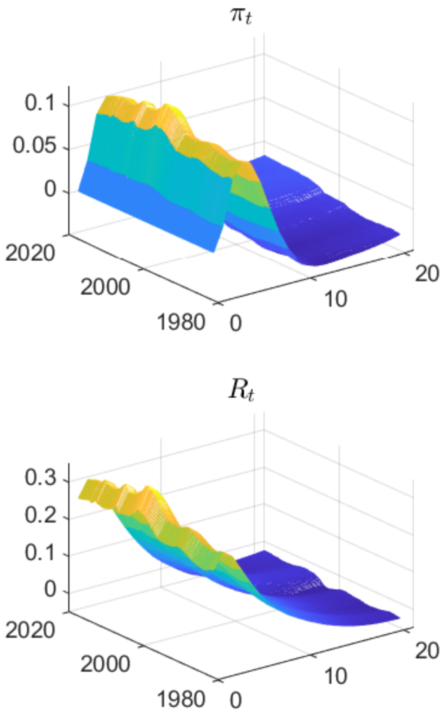

Figure 5: Norway – median impulse responses to a monetary policy shock from the TVP-VAR-SV model.

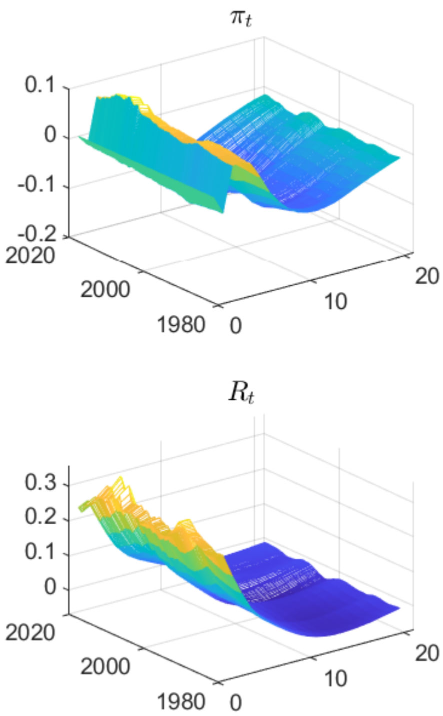

Figure 6: Sweden – median impulse responses to a monetary policy shock from the TVP-VAR-SV model.

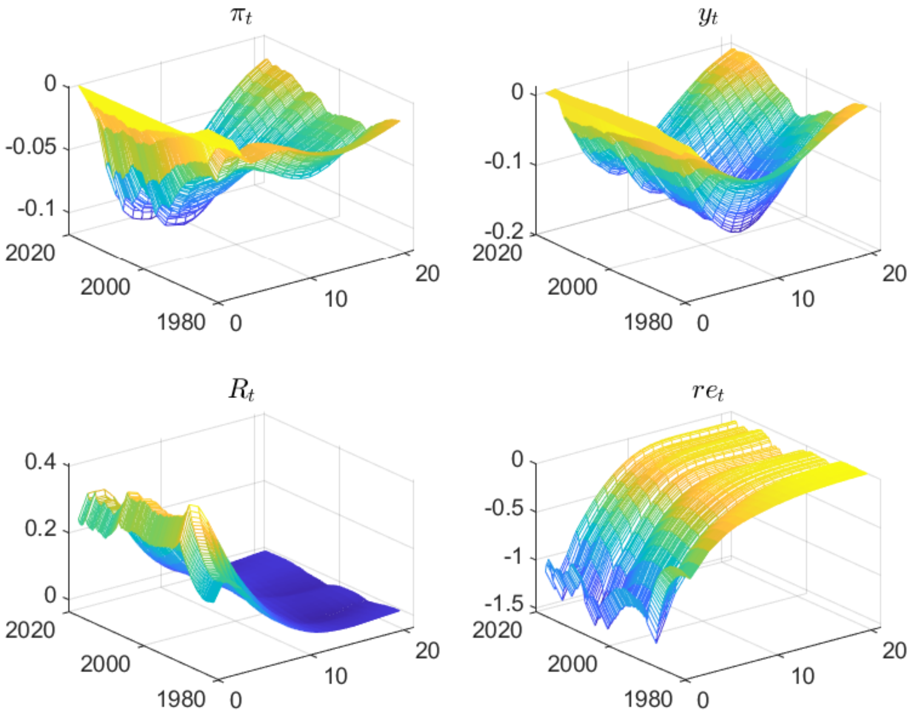

Figure 7: United Kingdom – median impulse responses to a monetary policy shock from the TVP-VAR-SV model.

As observed in the figures, the contractionary policy shock leads to a significant and persistent increase in the interest rate across all countries. The shock is gradually offset and the interest rate returns to its steady-state level after approximately three years in most countries. In some cases, the interest rate falls below the steady-state level before eventually returning to its pre-shock level (see Bjørnland, 2009).

Regarding the effect on output, we find that, in line with economic theory, output experiences a decline in most countries following the policy shock. Furthermore, we observe some time variation in the response of output in Australia, where the negative impact becomes less pronounced over time. Conversely, the response of output becomes more negative in Canada during the 1990s. For the remaining countries, the responses of output remain relatively stable over time.

In Australia, New Zealand, and Norway, we uncover evidence of an initial "price puzzle" phenomenon, whereby inflation initially rises following a contractionary monetary policy shock. However, after this initial phase, the effect on inflation eventually becomes negative. This finding aligns with Bjørnland (2009), who also identifies initial price puzzles for Australia and New Zealand. These price puzzles can be explained by the cost channel of the interest rate, where higher interest rates increase borrowing costs and marginal costs for firms. Consequently, firms may pass on these increased costs to consumers in the form of higher prices. This mechanism is supported by studies such as Ravenna and Walsh (2006); Chowdhury et al. (2006), which highlight the link between interest rates, borrowing costs, and consumer prices.

When examining the response of the real exchange rate, we find that it appreciates instantaneously in almost all countries following the policy shock. This suggests that there is no evidence of the exchange rate puzzle in any of the countries studied, which contrasts with the findings of Grilli and Roubini (1995). However, there is some evidence of time variation in the response of the real exchange rate, particularly in Australia, where the extent of appreciation diminishes over time. On the other hand, in Canada, New Zealand, and Sweden, the exchange rate exhibits somewhat larger appreciation in the latter part of the sample.

While the median responses indicate time variation in the exchange rate, they only provide an average depiction of the quantitative effects and do not inform us about the statistical significance of these effects. Therefore, we also consider the 68% posterior credible interval for the exchange rate response at selected time periods, along with the median response. Figure 8 illustrates the results. As shown, the exchange rate responses are not statistically significantly different from zero in the post-2000s for Australia and Norway. Moreover, the delay in overshooting is relatively short, with the maximum impact of the policy shock being delayed by only one quarter in most countries over time, except for Norway, where delayed overshooting is evident, especially in the first half of the sample. However, the additional appreciation following the impact response is quantitatively small for all countries, and the exchange rate subsequently depreciates back to its long-run level.[10]

These findings align with economic theory, specifically Dornbusch's (1976) EROH, which suggests that an increase in the interest rate should cause the exchange rate to appreciate instantaneously and then gradually depreciate. Our results are consistent with those of Bjørnland (2009), who also confirms Dornbusch's theory. However, it is worth noting that Bjørnland (2009) estimates a time-invariant VAR over the post-1980s, representing the average effect over that time period, which encompasses both pre- and post-inflation targeting

[10] In the long-run, the policy shock has no effect on the level of the real exchange rate by construction (i.e., the long-run restriction imposed on the VAR).

regimes. In contrast, our results indicate that Dornbusch's EROH holds not only on average over time but also across time for most countries in our sample, encompassing both pre- and post-inflation targeting periods. In the context of small open economies, Kim and Lim (2018) have presented evidence of puzzling results when estimating their VAR with pooled data from different monetary policy regimes. They suggest estimating the model over a homogeneous (inflation targeting) monetary policy regime to avoid strong puzzles. However, our results suggest that Dornbusch's EROH holds consistently across time for most countries in our sample, including both pre- and post-inflation targeting periods.[11] This underscores the importance of our TP-VAR-SV setting along with the combination of short-run and long-run restrictions.

[11] The only exceptions are Australia and Norway where monetary policy shocks do not have any significant effects on exchange rates in the post-2000s period.

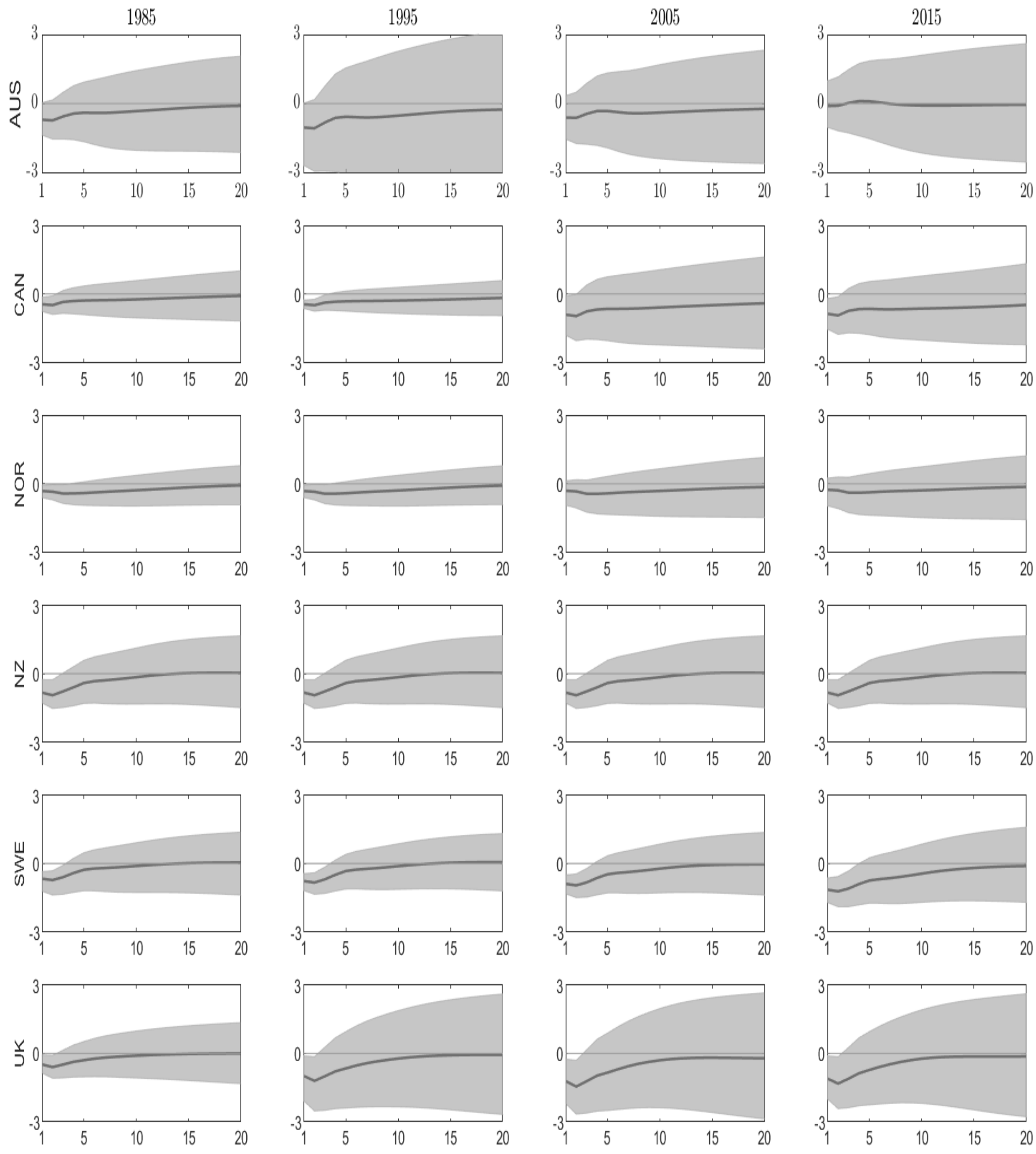

Figure 8: Impulse responses of the real exchange to a monetary policy shock in selected periods.

3.2 Model selection

In the previous section, we observed that there is limited time variation in the estimated impulse responses for most countries in our dataset, indicating that time-varying VAR coefficients may not play a significant role in explaining the dynamics of exchange rate fluctuations. However, we also found substantial variation in the estimated standard deviations of the policy shocks for all six countries, indicating a significant decline in the volatility of monetary policy shocks over time. In this section, we aim to formally assess the importance of incorporating time variation in the estimated coefficients and stochastic volatility in the shocks. To achieve this, we estimate three alternative models:

1. A conventional VAR with constant coefficients and homoscedastic shocks (henceforth, VAR); see Sims (1980).

2. A VAR with time-varying coefficients and homoscedastic shocks (henceforth, TVP-VAR); see Cogley and Sargent (2001).

3. A VAR with constant coefficients and stochastic volatility (henceforth, VAR-SV).

Equation (2.1) represents the general framework within which all three alternative models (VAR, TVP-VAR, and VAR-SV) can be nested. For example, the TVP-VAR model is obtained from the TVP-VAR-SV model (2.1) by imposing the restriction $\Omega_t = \Omega$ for all $t = 1, \ldots, T$. The standard VAR model is obtained from model (2.1) by imposing the restrictions $\Omega_t = \Omega$, $B_{i,t} = B_i$ and $A_t = A$ for all $t = 1, \ldots, T$ and $i = 1, \ldots, p$. Finally, the VAR-SV model is a nested version of the TVP-VAR-SV model (2.1) in which the coefficients are not time-varying but the shocks feature stochastic volatility. Therefore, all these models can be estimated using the same Bayesian MCMC technique as above, by specifying dogmatic priors of zero time variation for the respective matrices.

A formal way to choose between alternate models in a Bayesian framework is to compare their Bayes Factors. To this end, suppose we want to compare the in-sample fit of an arbitrary model M_i with a distinct model M_j. Each model $M_k, k = i, j$, is defined by a likelihood function of observing the data given the parameters of the model, denoted by $p\left(\boldsymbol{Y}|\boldsymbol{\theta}_k, M_k\right)$ where $\boldsymbol{Y}$ is a vector or matrix of data, and a prior specification for the parameters vector or matrix $\boldsymbol{\theta}_k$ of the model M_k is denoted by $p\left(\boldsymbol{\theta}_k|M_k\right)$. Then, the Bayes Factor of M_i against M_j is defined as

$$BF_{ij} = \frac{p\left(\boldsymbol{Y}|M_i\right)}{p\left(\boldsymbol{Y}|M_j\right)}, \tag{3.1}$$

where the marginal likelihood of model M_k, $p\left(\boldsymbol{Y}|M_k\right)$, is given by

$$p\left(\boldsymbol{Y}|M_k\right) = \int p\left(\boldsymbol{Y}|\boldsymbol{\theta}_k, M_k\right) p\left(\boldsymbol{\theta}_k|M_k\right) d\boldsymbol{\theta}_k. \tag{3.2}$$

The marginal likelihood can serve as a measure to compare the fit of different models, as it can be interpreted as a density forecast of the observed data $\boldsymbol{Y}$ under a specific model M_k (Geweke and Amisano, 2011). When comparing two models, M_i and M_j, a Bayes Factor (BF_{ij}) greater than 1 indicates that the observed data are more likely under model M_i than model M_j. Kass and Raftery (1995) propose a guideline for interpreting Bayes Factors. According to their suggestion, a Bayes Factor below 2 is considered "not worth more than a bare mention," indicating weak evidence in favor of either model. A Bayes Factor between 2 and 6 suggests "positive" evidence in favor of one of the models, between 6 and 10 suggests "strong" evidence, and a Bayes Factor larger than 10 suggests "very strong" evidence. Therefore, when comparing models using the Bayes Factor, the magnitude of the Bayes Factor can provide an indication of the strength of evidence in favor of one model over the other.

While the Bayes factor has simple theoretical foundations, computation of marginal likelihoods of the highly dimensional time-varying models presents a non-trivial computational concern. Early attempts to compute the likelihood of such highly dimensional models relied on estimation techniques which utilize the conditional likelihood (e.g., Koop et al., 2009). More recently, authors have shown that this approach can be extremely inaccurate. For instance, Chan and Grant (2015) show that the marginal likelihood estimates computed using the (modified) harmonic mean as in Gelfand and Dey (1994) can have a substantial finite sample bias and can thus lead to inaccurate model selection. Frühwirth-Schnatter and Wagner (2010) provide a similar inference for Chib's marginal likelihood method (see, Chib, 1995). To overcome these concerns we follow a series of work on the efficient estimation of marginal likelihood functions for TVP-VARs which are not based on the conditional likelihood as those previously mentioned, but instead, the integrated likelihood (i.e., the marginal density of the data unconditional on the time-varying coefficients and log-volatilities).

We estimate the marginal likelihood for the VAR and TVP-VAR following Chan and Grant (2016), and that of the VAR-SV and TVP-VAR-SV through the efficient sampling algorithm of Chan and Eisenstat (2018). The model comparison results are reported in Table 1. We report the log-marginal likelihoods (ML), where the model with the highest log-ML is preferred, along with the Bayes factor. The Bayes Factor is calculated as 2(log-data density H1 - log-data density H0), where the null hypothesis (H0) is taken to be the constant parameter VAR model. The results show that the VAR-SV model provides the best fit for all countries in our sample, fitting substantially better than the other specifications. This suggests that stochastic volatility is a significant modelling feature and ignoring it would lead to biased estimates of the underlying shocks. Table 1 also shows that the fit of the TVP-VAR model is worse than a conventional constant parameter VAR for all countries. This shows clear evidence that the dynamic relationships between the observed variables have remained stable across time in these countries.

Table 1: Log-marginal likelihoods

Country	VAR	VAR-SV	TVP-VAR	TVP-VAR-SV
Australia	-1131.5 (0)	**-988** (287)	-1244 (-225)	-1103.3 (62.4)
Canada	-1079.1 (0)	**-1012** (134)	-1180.4 (-203)	-1107.5 (-57)
New Zealand	-1357.2 (0)	**-1256.2** (202)	-1457.2 (-200)	-1364.8 (-15)
Norway	-1169.7 (0)	**-1086.5** (166)	-1267.5 (-196)	-1176.2 (-13)
Sweden	-1130.4 (0)	**-1031.2** (198)	-1169.6 (-78)	-1123.4 (14)
United Kingdom	-1052.8 (0)	**-956.2** (193.2)	-1154.4 (-203)	-1057.4 (-9.2)

Note: Bayes Factor in parentheses calculated as 2(log-data density H1 - log-data density H0), where the null hypothesis (H0) is taken to be the constant parameter VAR model.

After establishing that the VAR-SV model offers the most accurate fit within the sample, we now proceed to analyze the estimation results derived from this model. For the VAR-SV model, equation (2.1) becomes

$$\boldsymbol{Y}_t = c + B_1\boldsymbol{Y}_{t-1} + \ldots + B_p\boldsymbol{Y}_{t-p} + u_t, \quad t = p+1, \ldots, T, \tag{3.3}$$

where $\text{var}(u_t) = \Omega_t = A^{-1}\Sigma_t\Sigma_t' A^{-1'}$, B_j $(j = 1, \ldots, p)$ and A are time-invariant but Σ_t, which captures stochastic volatility, is time-varying.

Figure 16 in Appendix B displays the posterior estimates, revealing a substantial decline in the volatility of monetary policy shocks over time. We note that the volatility estimates remain remarkably similar throughout the study period and closely resembles Figure 1. Moving forward, we examine the estimated impulse responses to an 'average-sized' monetary policy shock using the VAR-SV model. The corresponding results for the six countries are presented in Figures 17-22 in the Appendix.

Upon analyzing these figures, we observe that the constant-parameter VAR-SV model effectively captures the dynamics of monetary policy shocks. Specifically, the real exchange rate response shows an immediate appreciation in nearly all countries, indicating no evidence of the exchange rate puzzle. Additionally, any delay in overshooting is relatively short, typically lasting no longer than one quarter. Subsequently, the exchange rate follows a depreciation pattern consistent with Dornbusch (1976). However, it is worth noting that we also uncover some disparities in the estimated response of the exchange rate. Particularly, the estimated response turns out to be statistically insignificant in Australia and Norway, even during the earlier part of the sample. Interestingly, this finding persists in several robustness checks, as discussed later in the analysis.

3.3 Monetary policy and exchange rate volatility

Given the significance of stochastic volatility and the decline in the volatility of policy shocks, the next step is to examine the implications for exchange rate and macroeconomic fluctuations. To do this, we analyze the forecast error variance decompositions (FEVDs) for monetary policy shocks using the VAR-SV model. Figures 9-14 present these decompositions for the six small open economies at various horizons. The solid lines represent median responses, while the shaded areas show 68% credible intervals. By studying these decompositions, we gain insights into the fraction of forecast error variances explained by monetary policy shocks and how it has evolved over time.

The figures illustrate that monetary policy shocks accounted for about 20%-60% of the exchange rate fluctuations in the 1980s. However, there has been a significant decrease in the importance of policy shocks in explaining exchange rate movements since the 1990s. This decline aligns with the adoption of inflation-targeting policies in many countries within our sample, accompanied by a reduction in the volatility of monetary policy shocks. In fact, the impact of policy shocks on exchange rate volatility has become negligible in Australia, Norway, Sweden (since the 2000s), and in Canada and the United Kingdom (since the 2010s). Our findings diverge from Bjørnland (2009), who observe that policy shocks played a substantial role in explaining exchange rate volatility in several small open economies. Instead, our results suggest that once stochastic volatility is considered, monetary policy shocks are no longer a significant driver of exchange rate fluctuations in small open economies, particularly since the 2000s. This result is primarily driven by the substantial decline in the estimated volatility of monetary policy shocks over the analyzed period. Failing to account for this declining trend over time will lead to overestimating the contribution of monetary policy shocks in explaining exchange rate fluctuations.

Regarding the other variables, our findings indicate that monetary policy shocks only account for a small fraction of output and inflation volatility, which aligns with existing literature on monetary policy. Moreover, the contributions of policy shocks in explaining volatility in output and inflation have diminished over time. As a result, policy shocks have become almost insignificant in explaining fluctuations in inflation and output for most countries, particularly in the post-2000s period.

Overall, our results suggest that monetary policy in small open economies has become less idiosyncratic over time, especially after the implementation of inflation-targeting frameworks. This has led to a decline in the volatility of monetary policy shocks and a subsequent decrease in the significance of policy shocks in driving both macroeconomic fluctuations and exchange rate volatility.

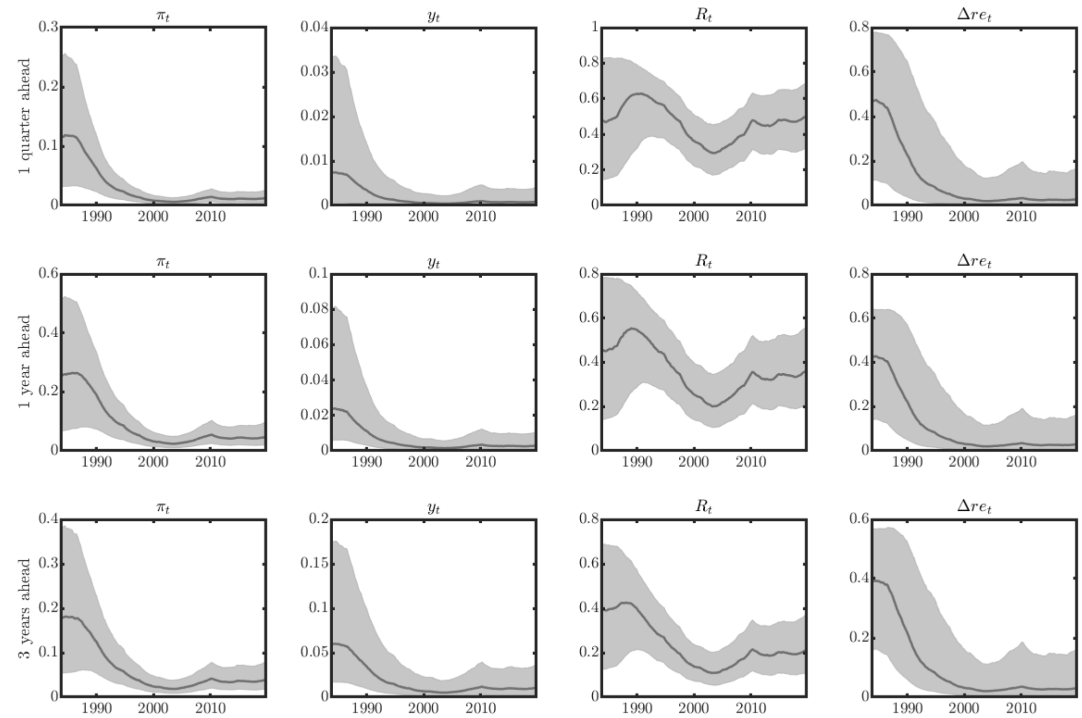

Figure 9: Australia – FEVD for monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

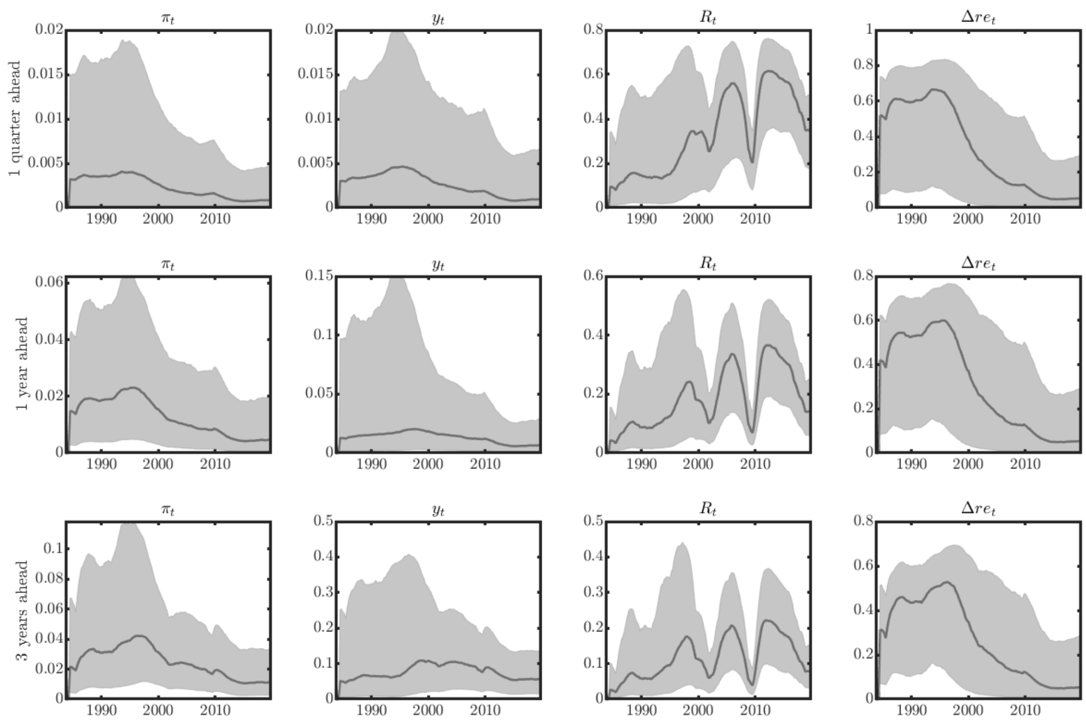

Figure 10: Canada – FEVD for monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

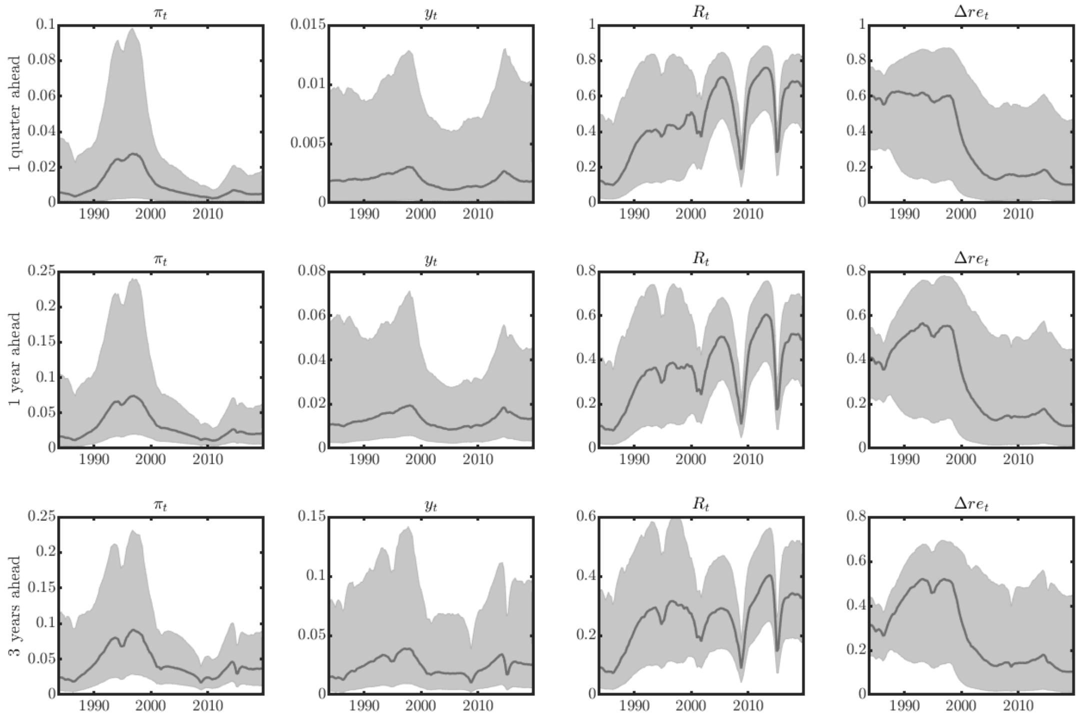

Figure 11: New Zealand – FEVD for monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

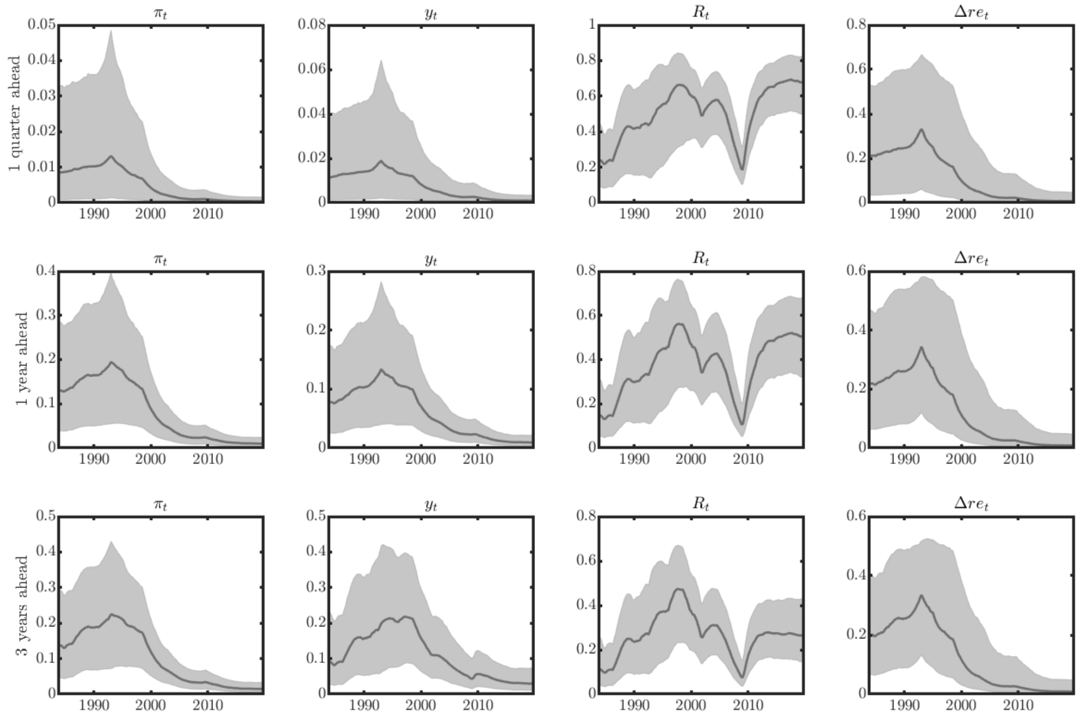

Figure 12: Norway – FEVD for monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

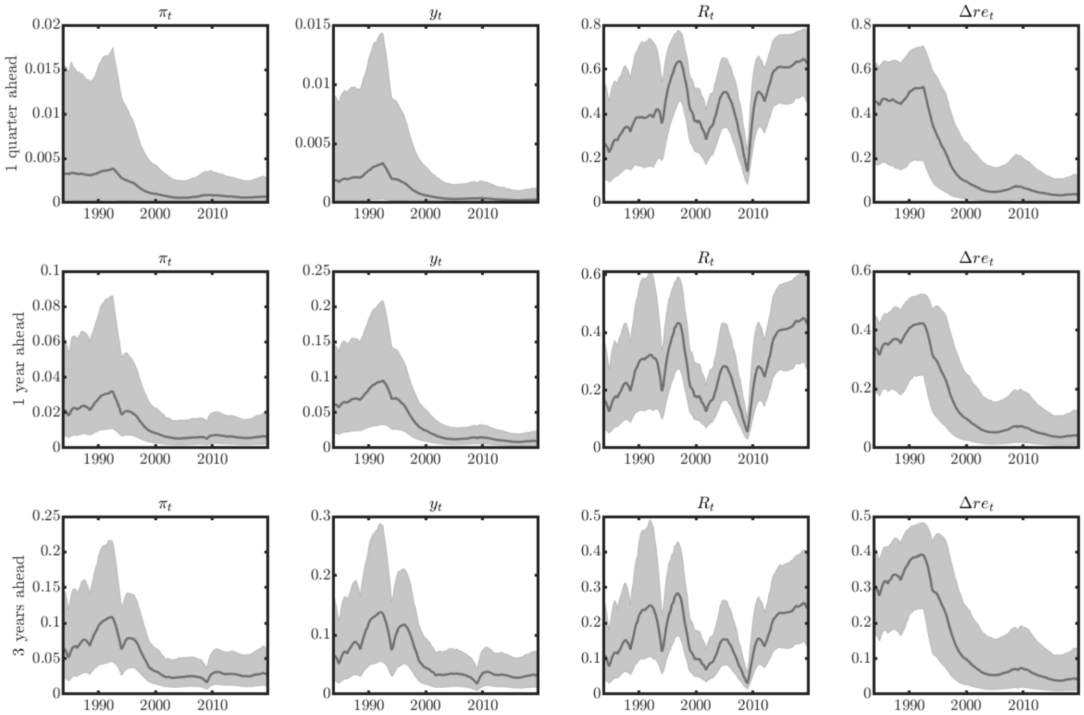

Figure 13: Sweden – FEVD for monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

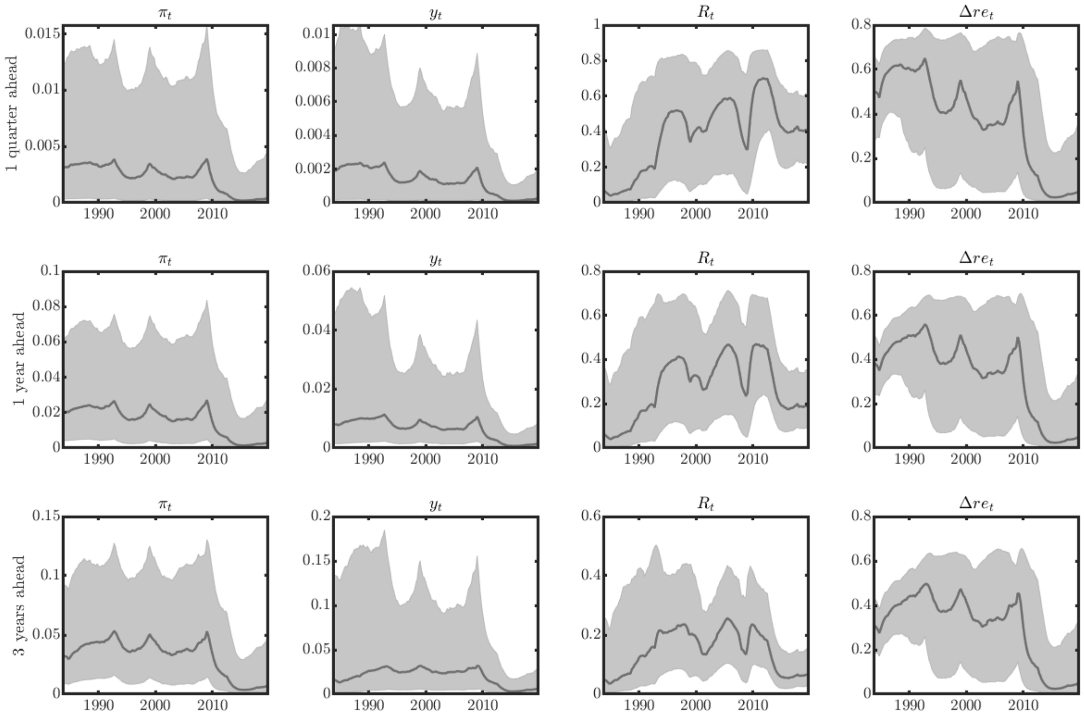

Figure 14: United Kingdom – FEVD for monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

3.4 Uncovered interest parity (UIP)

The literature has provided substantial evidence that the uncovered interest rate parity (UIP) does not hold; see e.g Fama (1984); Engel (1996); Burnside et al. (2006); Burnside (2019). However, a key concern for monetary policymakers is whether UIP holds following a monetary policy shock. Unfortunately, the answer is mixed. While some studies such as Eichenbaum and Evans (1995) and Faust and Rogers (2003) find that UIP does not hold, others such as Bjørnland (2009) suggest that monetary policy shocks generate exchange rate movements that are broadly consistent with UIP. In support of the former strand of literature, Scholl and Uhlig (2008) find that UIP does not hold in the U.S. data following a monetary policy even in the absence of the delayed overshooting puzzle. They argue that the forward discount puzzle, which implies a failure of UIP, is a distinct phenomenon rather than a mere reflection of the delayed overshooting puzzle.

To investigate this further, we follow Eichenbaum and Evans (1995) and Bjørnland (2009) and define ψ_t as the ex-post difference in return between holding one period of foreign bonds or one period of domestic bonds. Measured in domestic currency, ψ_t is then given by

$$\psi_t = R_t^* - R_t + 4 * (s_{t+1} - s_t) \tag{3.4}$$

where s_t is the nominal exchange rate and s_{t+1} is the one-quarter ahead forecast of the exchange rate.[12] One implication of UIP is that

$$E_t(\psi_{t+j}) = 0 \tag{3.5}$$

for all $j \geq 0$, where $E_t(\cdot)$ denotes the expectations operator given the information available at time t.

Figure 15 presents the median and 68% credible intervals of the dynamic response function (3.5) based on the estimated impulse responses from the VAR-SV model.[13] Under the uncovered interest parity (UIP), these responses should ideally be zero. However, the

[12] The exchange rate is multiplied by 4 to be annualized since the interest rate is measured in annual terms.
[13] Note that the effect of policy shocks on prices has been adjusted to get the effect on the nominal exchange rate, which is given by $s_t = re_t - p_t^* + p_t$. However, we can only correct for domestic prices as foreign prices are not included in the VAR (see Bjørnland, 2009). This restriction is equivalent to assuming that domestic monetary policy has a negligible effect on foreign prices, which is a plausible small open economy assumption.

figure suggests that we can reject the hypothesis that the response functions are equal to zero for certain countries.

For example, in Norway, the expected return from investing in foreign short-term bonds decreases relative to the returns from investing in domestic bonds following a contractionary monetary policy shock. In Sweden and the United Kingdom, there is an immediate negative deviation from zero. Thereafter, the responses fluctuate around zero for the UK while it turns positive for Sweden before returning back to zero. In contrast, in Australia, Canada, and New Zealand, the responses predominantly fluctuate around zero, which is consistent with UIP.

Overall, our findings indicate deviations from UIP in three out of the six countries, particularly in the short term. These results partially align with Scholl and Uhlig (2008), who also provide evidence of the forward discount puzzle even in the absence of delayed overshooting.

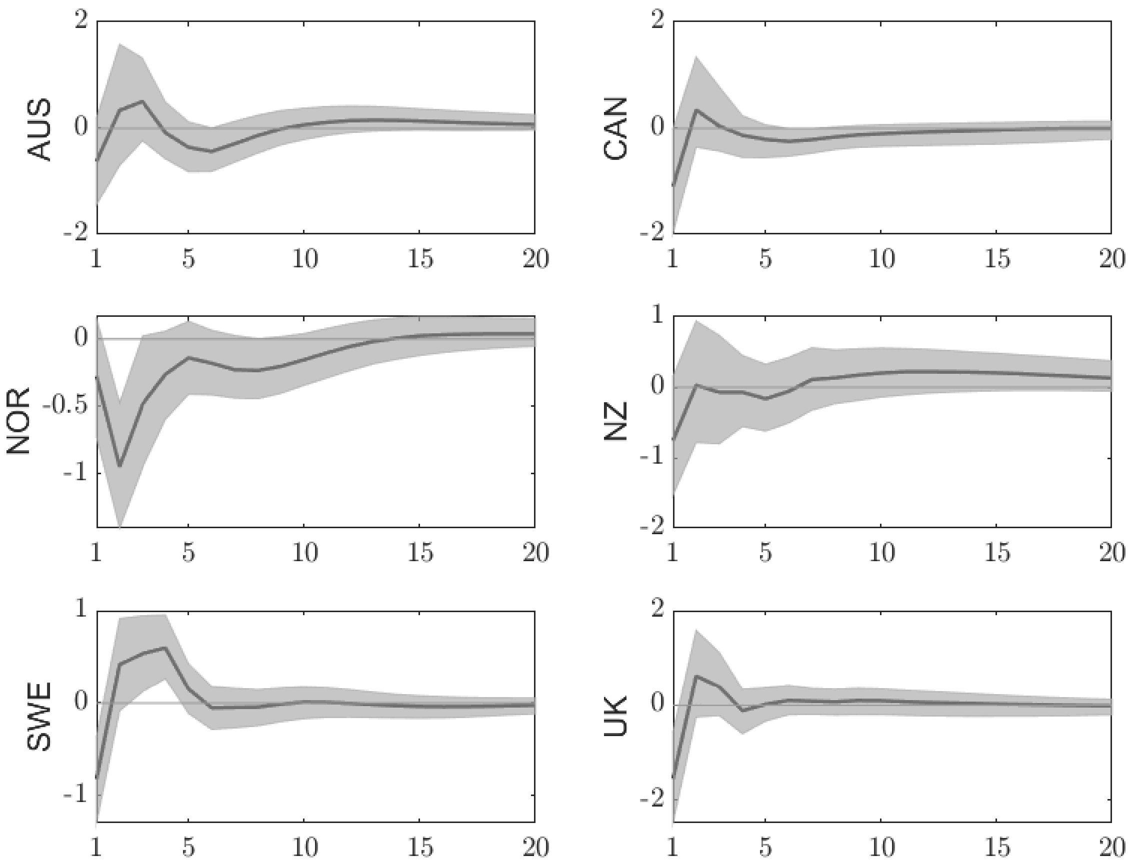

Figure 15: Conditional excess returns from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

4 Robustness checks

We conduct robustness checks on our results in several dimensions: model specification, alternative VAR lag length, alternative priors, and using additional variables in the specification. These checks were performed for the VAR-SV model. Regarding the model specification, we implement the estimations without detrending the data, except for the output variable. Next, we examine the robustness of our results to the lag length by estimating the VARs with two lags instead of three, as done in Cogley and Sargent (2005) and Primiceri (2005). We also assess the robustness to the priors by using tighter priors for the variance and covariance states. Our baseline results indicate significant time variation in the volatility of monetary policy shocks. By applying a tighter prior, we aim to determine whether substantial volatility drifts in policy shocks over time still persist. Finally, we examine the robustness of our results by including oil price in the VAR. Oil price is an important leading indicator that monetary policymakers may react to. Additionally, as discussed in Bjørnland and Halvorsen (2014), some of the countries in our study, namely Canada, Norway, and the United Kingdom, are net oil exporters. Therefore, fluctuations in oil prices in these countries may affect exchange rates. Specifically, we include the growth rate of oil price in the estimated VAR and treat it as the most exogenous variable, given the nature of these small open economies.

Figures 23-35 in Appendix C present the results of the robustness exercises conducted for all six countries. Similar to the previous analysis, we illustrate the impulse responses (Figures 23-28), the conditional excess returns (Figure 29), and the forecast error variance decompositions (FEVDs) (Figures 30-35) for each country. In each figure, we display the baseline results for the VAR-SV model and overlay the estimated posterior medians obtained from the robustness exercises. Importantly, these changes do not significantly alter the main findings, indicating the robustness of our results. Specifically, we find no evidence of the exchange rate puzzle or delayed overshooting. However, we do observe evidence of the forward discount puzzle in certain countries, suggesting a violation of the uncovered interest parity (UIP) in those specific contexts. Additionally, we note a considerable decline in the relevance of monetary policy shocks in explaining both exchange rate fluctuations and macroeconomic volatility since the 1990s.

5 Conclusion

This paper examines the dynamic and changing impacts of monetary policy shocks on six small open economies. To achieve this, we employ a Bayesian framework to estimate a time-varying parameter VAR model with stochastic volatility. In order to establish identification, we show how to implement a combination of short-run and long-run restrictions that maintain the contemporaneous relationship between the interest rate and the exchange rate. While the motivation for our algorithm stemmed from the desire to address this pertinent empirical question, we emphasize that it is general in the sense that it can be applied to other applications that require the use of sign and zero restrictions for identification purposes.

Our findings reveal several key observations. Firstly, in most countries, a contractionary monetary policy shock leads to an immediate appreciation of the exchange rate. This finding contradicts the exchange rate puzzle. Additionally, the delay in overshooting is relatively short, with the maximum impact of the policy shock on the exchange rate being delayed by only one quarter. Any additional appreciation that follows the initial response is quantitatively small, and the exchange rate subsequently depreciates back to its long-run level. This pattern aligns with the Dornbusch overshooting hypothesis. Furthermore, the estimation results obtained from the time-varying VAR model indicate that the overshooting hypothesis holds not just on average, as previous studies have shown, but consistently over time for all countries studied.

However, we do find evidence of the forward discount puzzle, indicating a violation of the uncovered interest parity (UIP) after a monetary policy shock in three out of the six countries. The exceptions to this puzzle are Australia, Canada, and New Zealand, where the responses align with UIP. Hence, our findings demonstrate the presence of the forward discount puzzle even in the absence of delayed overshooting. This confirms that the forward discount puzzle is not merely a byproduct or reflection of delayed overshooting but rather an independent characteristic observed in many countries.

In terms of temporal changes, our analysis provides no evidence of time variations in the dynamic effects of monetary policy shocks. Indeed, the impulse responses to an identified policy shock exhibit limited variations over time. However, we do observe a significant decline in the estimated volatility of policy shocks across all countries as time progresses. To validate this finding, we conduct formal model selection exercises by calculating and comparing marginal likelihoods and Bayes factors of different models. These exercises confirm that a VAR model without time-varying dynamics but with heteroscedastic shocks is favored by the data in all countries.

The decrease in the estimated volatility carries important implications. Specifically, it leads to a notable reduction in the influence of monetary policy shocks on both exchange rates and macroeconomic fluctuations since the 1990s. Remarkably, this time-frame aligns with the adoption of inflation-targeting and central bank independence in many countries. Hence, our findings suggest a link between the implementation of these policy frameworks and the diminishing role of monetary policy shocks in driving economic and exchange rate dynamics.

References

Benati, L. and T. A. Lubik (2014a). Sales, inventories and real interest rates: A century of stylized facts. *Journal of Applied Econometrics 29*(7), 1210–1222.

Benati, L. and T. A. Lubik (2014b). The time-varying Beveridge curve. In *Advances in Non-Linear Economic Modeling*, pp. 167–204. Springer.

Binning, A. (2013). Underidentified SVAR models: A framework for combining short and long-run restrictions with sign-restrictions. Technical report, Norges Bank.

Bjørnland, H. C. (2009). Monetary policy and exchange rate overshooting: Dornbusch was right after all. *Journal of International Economics 79*(1), 64–77.

Bjørnland, H. C. and J. I. Halvorsen (2014). How does monetary policy respond to exchange rate movements? New international evidence. *Oxford Bulletin of Economics and Statistics 76*(2), 208–232.

Blanchard, O. J. and D. Quah (1988). The dynamic effects of aggregate demand and supply disturbances.

Burnside, C. (2019). Exchange rates, interest parity, and the carry trade. In *Oxford Research Encyclopedia of Economics and Finance*.

Burnside, C., M. S. Eichenbaum, I. Kleshchelski, and S. Rebelo (2006). The returns to currency speculation.

Canova, F. and L. Gambetti (2009). Structural changes in the US economy: Is there a role for monetary policy? *Journal of Economic Dynamics and Control 33*(2), 477–490.

Castelnuovo, E., G. Pellegrino, G. Ranzato, et al. (2022). *Delayed Overshooting Puzzle: Does Systematic Monetary Policy Matter?* Department of Economics and Management" Marco Fanno", University of Padova.

Chan, J. C. and E. Eisenstat (2018). Bayesian model comparison for time-varying parameter VARs with stochastic volatility. *Journal of Applied Econometrics 33*(4), 509–532.

Chan, J. C. and A. L. Grant (2015). Pitfalls of estimating the marginal likelihood using the modified harmonic mean. *Economics Letters 131*, 29–33.

Chan, J. C. and A. L. Grant (2016). Fast computation of the deviance information criterion for latent variable models. *Computational Statistics & Data Analysis 100*, 847–859.

Chib, S. (1995). Marginal likelihood from the gibbs output. *Journal of the american statistical association 90*(432), 1313–1321.

Chowdhury, I., M. Hoffmann, and A. Schabert (2006). Inflation dynamics and the cost channel of monetary transmission. *European Economic Review 50*(4), 995–1016.

Cogley, T. and T. J. Sargent (2001). Evolving post-World War II US inflation dynamics. *NBER Macroeconomics Annual 16*, 331–373.

Cogley, T. and T. J. Sargent (2005). Drifts and volatilities: monetary policies and outcomes in the post WWII US. *Review of Economic Dynamics 8*(2), 262–302.

Dornbusch, R. (1976). Expectations and exchange rate dynamics. *Journal of Political Economy 84*(6), 1161–1176.

Eichenbaum, M. and C. L. Evans (1995). Some empirical evidence on the effects of shocks to monetary policy on exchange rates. *The Quarterly Journal of Economics 110*(4), 975–1009.

Engel, C. (1996). The forward discount anomaly and the risk premium: A survey of recent evidence. *Journal of empirical finance 3*(2), 123–192.

Fama, E. F. (1984). Forward and spot exchange rates. *Journal of Monetary Economics 14*(3), 319–338.

Faust, J. and J. H. Rogers (2003). Monetary policy's role in exchange rate behavior. *Journal of Monetary Economics 50*(7), 1403–1424.

Frühwirth-Schnatter, S. and H. Wagner (2010). Stochastic model specification search for gaussian and partial non-gaussian state space models. *Journal of Econometrics 154*(1), 85–100.

Gali, J. (1992). How well does the is-lm model fit postwar us data? *The Quarterly Journal of Economics 107*(2), 709–738.

Gelfand, A. E. and D. K. Dey (1994). Bayesian model choice: asymptotics and exact calculations. *Journal of the Royal Statistical Society: Series B (Methodological) 56*(3), 501–514.

Geweke, J. and G. Amisano (2011). Hierarchical markov normal mixture models with applications to financial asset returns. *Journal of Applied Econometrics 26*(1), 1–29.

Grilli, V. and N. Roubini (1995). Liquidity and exchange rates: puzzling evidence from the G-7 countries. Technical report.

Haque, Q. and L. M. Magnusson (2021). Uncertainty shocks and inflation dynamics in the us. *Economics Letters 202*, 109825.

Haque, Q., L. M. Magnusson, and K. Tomioka (2021). Empirical evidence on the dynamics of investment under uncertainty in the US. *Oxford Bulletin of Economics and Statistics 83*(5), 1193–1217.

Inoue, A. and B. Rossi (2019). The effects of conventional and unconventional monetary policy on exchange rates. *Journal of International Economics 118*, 419–447.

Ismailov, A. and B. Rossi (2018). Uncertainty and deviations from uncovered interest rate parity. *Journal of International Money and Finance 88*, 242–259.

Kass, R. E. and A. E. Raftery (1995). Bayes factors. *Journal of the American Statistical Association 90*(430), 773–795.

Kim, S. and K. Lim (2018). Effects of monetary policy shocks on exchange rate in small open economies. *Journal of Macroeconomics 56*, 324–339.

Kim, S.-H., S. Moon, and C. Velasco (2017). Delayed overshooting: Is it an '80s puzzle? *Journal of Political Economy 125*(5), 1570–1598.

Koop, G., R. Leon-Gonzalez, and R. W. Strachan (2009). On the evolution of the monetary policy transmission mechanism. *Journal of Economic Dynamics and Control 33*(4), 997–1017.

Marcellino, M. and C. A. Favero (2004). Large datasets, small models and monetary policy in Europe.

Mojon, B. and G. Peersman (2001). A VAR description of the effects of monetary policy in the individual countries of the euro area. In *European Central Bank Working Paper Series*, Volume 92, pp. 1–49. European Central Bank.

Mumtaz, H. and L. Sunder-Plassmann (2013). Time-varying dynamics of the real exchange rate: An empirical analysis. *Journal of Applied Econometrics 28*(3), 498–525.

Nakajima, J. (2011). Time-varying parameter VAR model with stochastic volatility: An overview of methodology and empirical applications. *Monetary and Economic Studies 29*(1), 107–142.

Primiceri, G. E. (2005). Time varying structural vector autoregressions and monetary policy. *The Review of Economic Studies 72*(3), 821–852.

Ravenna, F. and C. E. Walsh (2006). Optimal monetary policy with the cost channel. *Journal of Monetary Economics 53*(2), 199–216.

Rossi, B. (2021). Forecasting in the presence of instabilities: How we know whether models predict well and how to improve them. *Journal of Economic Literature 59*(4), 1135–90.

Rubio-Ramirez, J. F., D. F. Waggoner, and T. Zha (2010). Structural vector autoregressions: Theory of identification and algorithms for inference. *The Review of Economic Studies 77*(2), 665–696.

Rüth, S. K. (2020). Shifts in monetary policy and exchange rate dynamics: Is dornbusch's overshooting hypothesis intact, after all? *Journal of International Economics 126*, 103344.

Rüth, S. K. and W. Van der Veken (2022). Monetary policy and exchange rate anomalies in set-identified svars: Revisited. Technical report, University of Erfurt and University of Ghent, mimeo.

Scholl, A. and H. Uhlig (2008). New evidence on the puzzles: Results from agnostic identification on monetary policy and exchange rates. *Journal of International Economics 76*(1), 1–13.

Sims, C. A. (1980). Macroeconomics and reality. *Econometrica: Journal of the Econometric Society*, 1–48.

Sims, C. A. (2001). Evolving post-World War II US inflation dynamics: comment. *NBER Macroeconomics Annual 16*, 373–379.

Yang, Y. and J. Zhang (2021). Effects of monetary policy on the exchange rates: A time-varying analysis. *Finance Research Letters*, 102–114.

Online Appendix

A Data Sources

All data are retrieved from the Federal Reserve Economic Data (FRED).

A.1 Australia

- **Consumer Price Index of All Items in Australia**

  https://fred.stlouisfed.org/series/AUSCPIALLQINMEI

- **Constant Price Gross Domestic Product in Australia**

  https://fred.stlouisfed.org/series/AUSGDPRQDSMEI

- **3-Month or 90-day Rates and Yields: Interbank Rates for Australia**

  https://fred.stlouisfed.org/series/IR3TIB01AUM156N

- **Real Effective Exchange Rates Based on Manufacturing Consumer Price Index for Australia**

  https://fred.stlouisfed.org/series/CCRETT01AUM661N

A.2 Canada

- **Consumer Price Index of All Items in Canada**

  https://fred.stlouisfed.org/series/CANCPIALLMINMEI

- **Gross Domestic Product by Expenditure in Constant Prices: Total Gross Domestic Product for Canada**

  https://fred.stlouisfed.org/series/NAEXKP01CAQ189S

- **3-Month or 90-day Rates and Yields: Interbank Rates for Canada**

  https://fred.stlouisfed.org/series/IR3TIB01CAM156N

- **Real Effective Exchange Rates Based on Manufacturing Consumer Price Index for Canada**

  https://fred.stlouisfed.org/series/CCRETT01CAM661N

A.3 New Zealand

- **Consumer Price Index: All Items for New Zealand**
  https://fred.stlouisfed.org/series/NZLCPIALLQINMEI

- **Leading Indicators OECD: Reference Series: Gross Domestic Product: Original Series for New Zealand**
  https://fred.stlouisfed.org/series/LORSGPORNZQ661S;

- **3-Month or 90-day Rates and Yields: Interbank Rates for New Zealand**
  https://fred.stlouisfed.org/series/IR3TIB01NZM156N

- **Real Effective Exchange Rates Based on Manufacturing Consumer Price Index for New Zealand**
  https://fred.stlouisfed.org/series/CCRETT01NZM661N

A.4 Norway

- **Consumer Price Index: All Items in Norway**
  https://fred.stlouisfed.org/series/NORCPIALLMINMEI

- **Real Gross Domestic Product for Norway**
  https://fred.stlouisfed.org/series/CLVMNACSCAB1GQNO

- **3-Month or 90-day Rates and Yields: Interbank Rates for Norway**
  https://fred.stlouisfed.org/series/IR3TIB01NOM156N

- **Real Effective Exchange Rates Based on Manufacturing Consumer Price Index for Norway**
  https://fred.stlouisfed.org/series/CCRETT01NOM661N

A.5 Sweden

- **Consumer Price Index: All Items in Sweden**
  https://fred.stlouisfed.org/series/SWECPIALLMINMEI

- **Leading Indicators OECD: Reference Series: Gross Domestic Product: Original Series for Sweden**

  https://fred.stlouisfed.org/series/LORSGPORSEQ661S

- **3-Month or 90-day Rates and Yields: Interbank Rates for Sweden**

  https://fred.stlouisfed.org/series/IR3TIB01SEM156N

- **Real Effective Exchange Rates Based on Manufacturing Consumer Price Index for Sweden**

  https://fred.stlouisfed.org/series/CCRETT01SWM661N

A.6 United Kingdom

- **Consumer Price Index of All Items in the United Kingdom**

  https://fred.stlouisfed.org/series/GBRCPIALLMINMEI

- **Gross Domestic Product by Expenditure in Constant Prices: Total Gross Domestic Product for the United Kingdom**

  https://fred.stlouisfed.org/series/NAEXKP01GBQ652S

- **3-Month or 90-day Rates and Yields: Interbank Rates for the United Kingdom**

  https://fred.stlouisfed.org/series/IR3TIB01GBM156N

- **Real Effective Exchange Rates Based on Manufacturing Consumer Price Index for the United Kingdom**

  https://fred.stlouisfed.org/series/CCRETT01GBM661N

A.7 Additional Data

- **Trade-weighted foreign interest rate.**

  We follow Bjørnland and Halvorsen (2014) in constructing data for the foreign interest rate. For Australia, New Zealand, Norway and Sweden, the foreign interest rate is a weighted average of the interest rates in the major trading partners, based on data from the respective central banks. For Canada and the UK, the foreign interest rate is assumed to be the Federal Funds rate since the U.S. comprises the bulk of the foreign trade weight in these countries.

- **Spot Crude Oil Price: West Texas Intermediate (WTI)**
  https://fred.stlouisfed.org/series/WTISPLC#0

B Estimation Results from the VAR-SV model

B.1 Standard Deviations

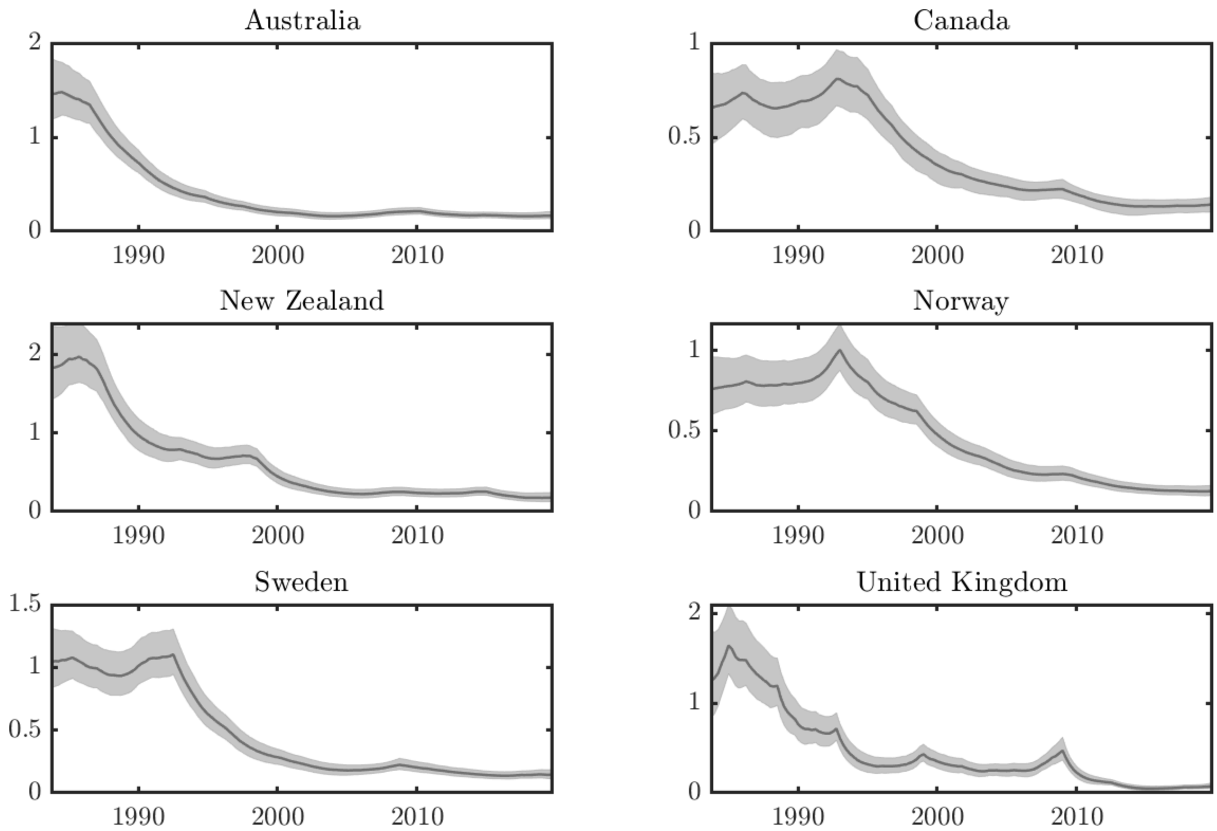

Figure 16: Estimated standard deviations from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

B.2 Impulse Responses

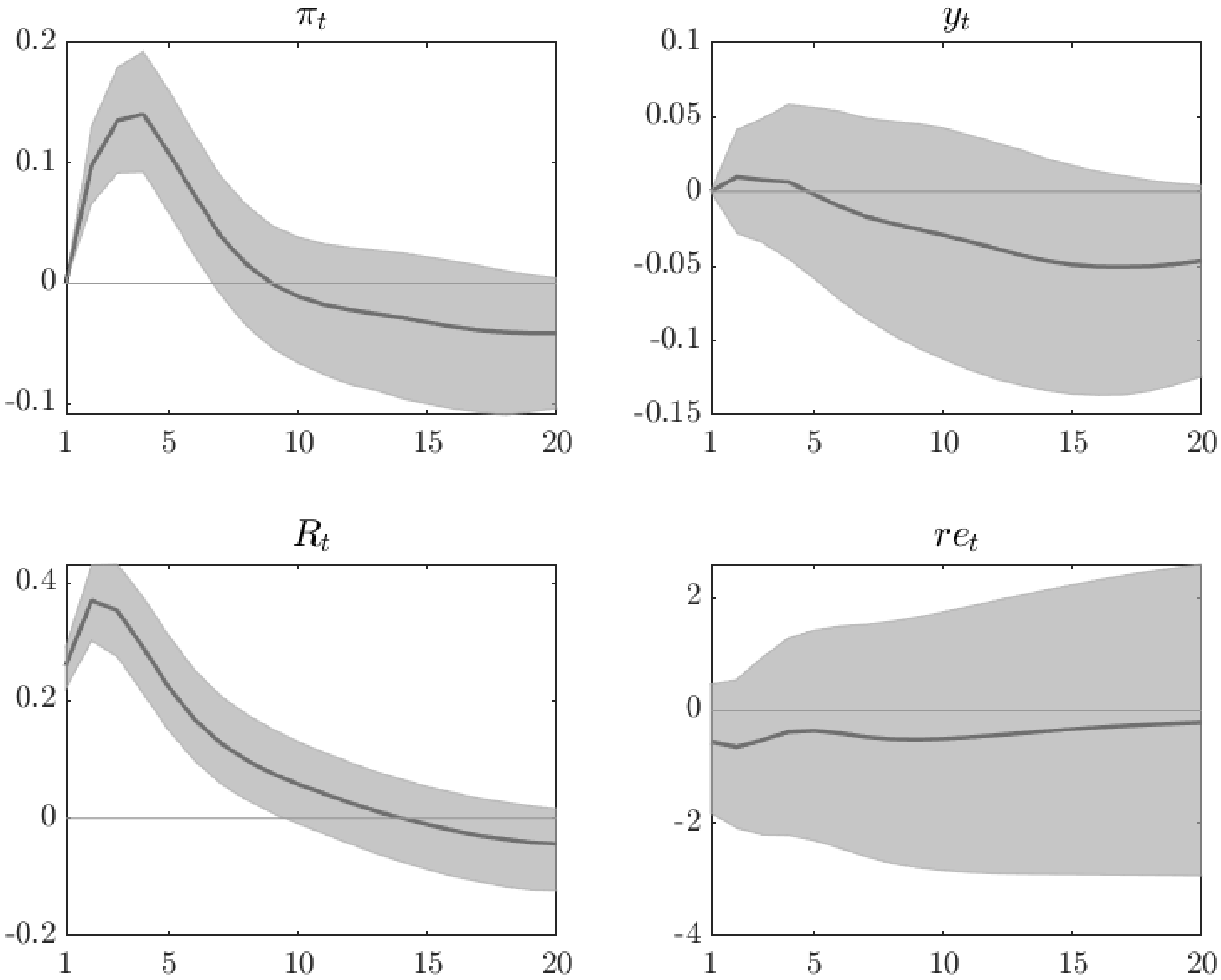

Figure 17: Australia – impulse responses to a monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

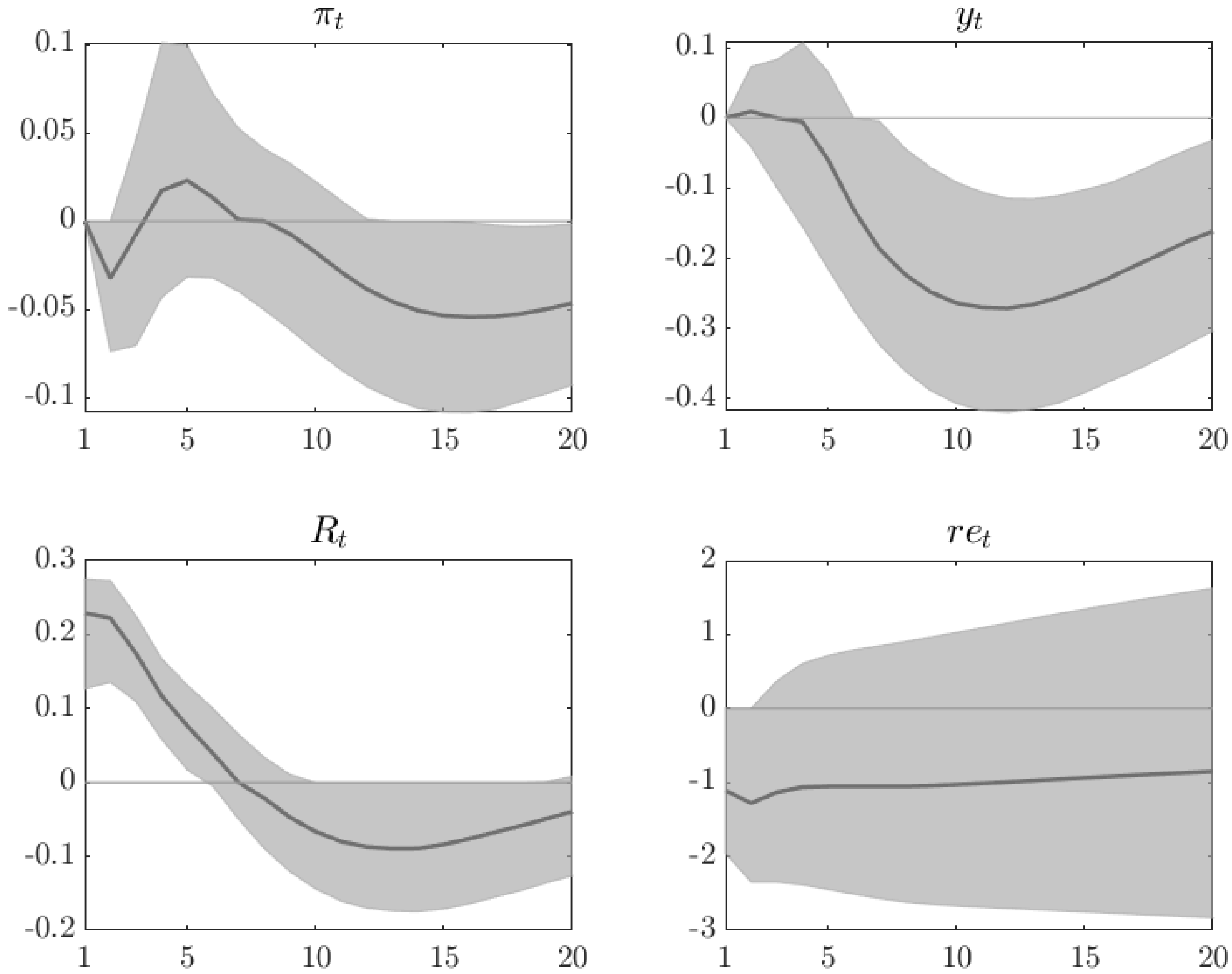

Figure 18: Canada – impulse responses to a monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

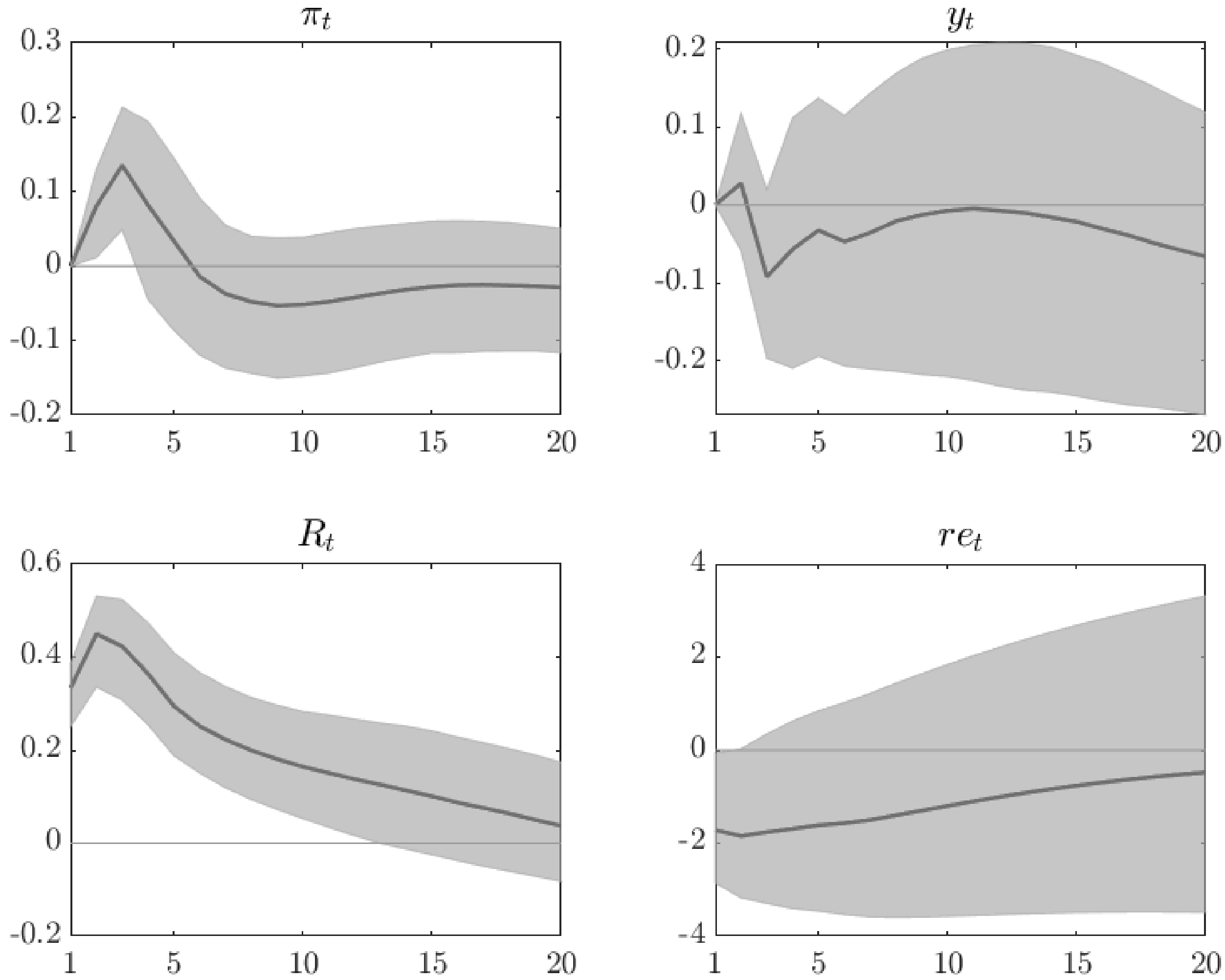

Figure 19: New Zealand – impulse responses to a monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

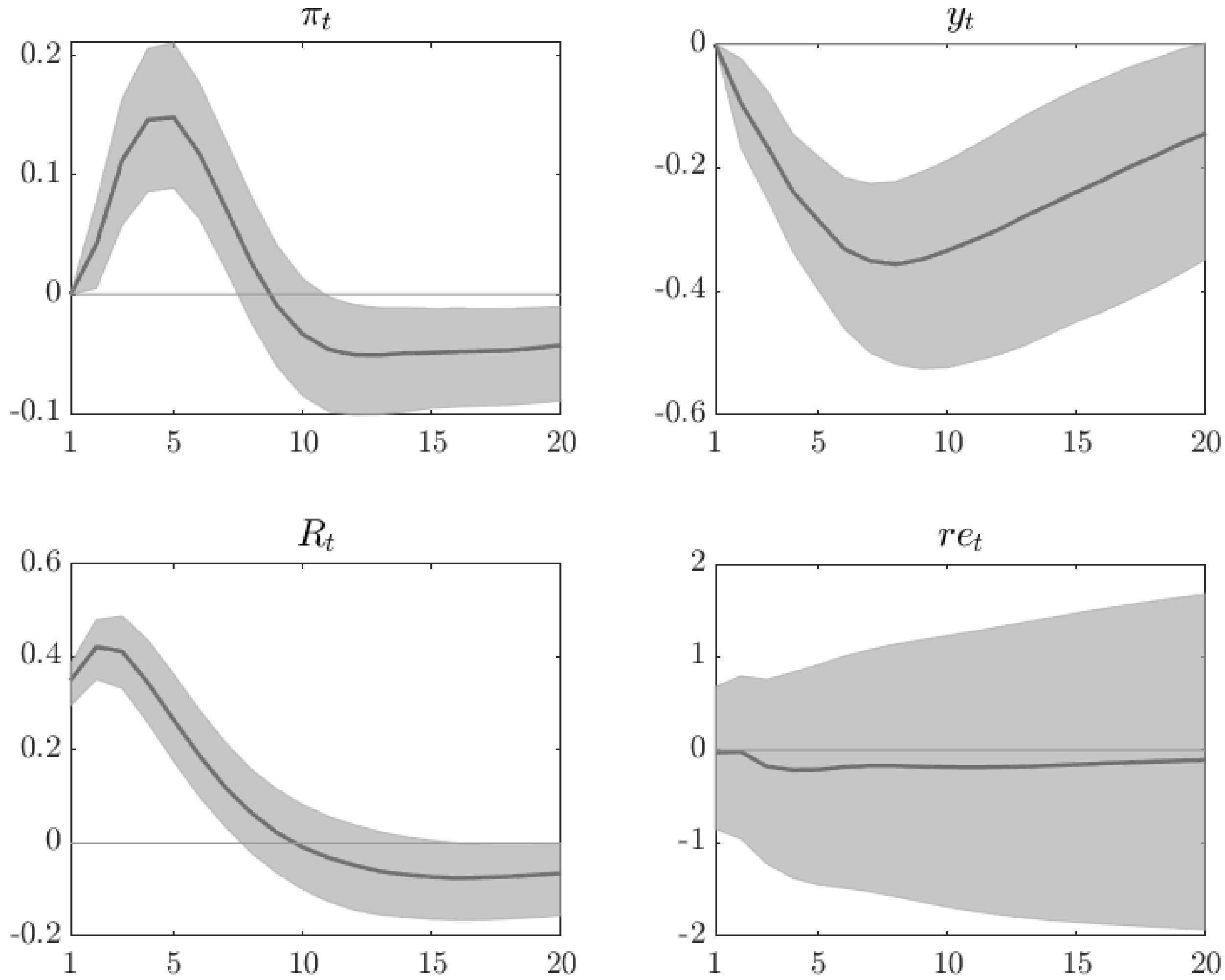

Figure 20: Norway – impulse responses to a monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

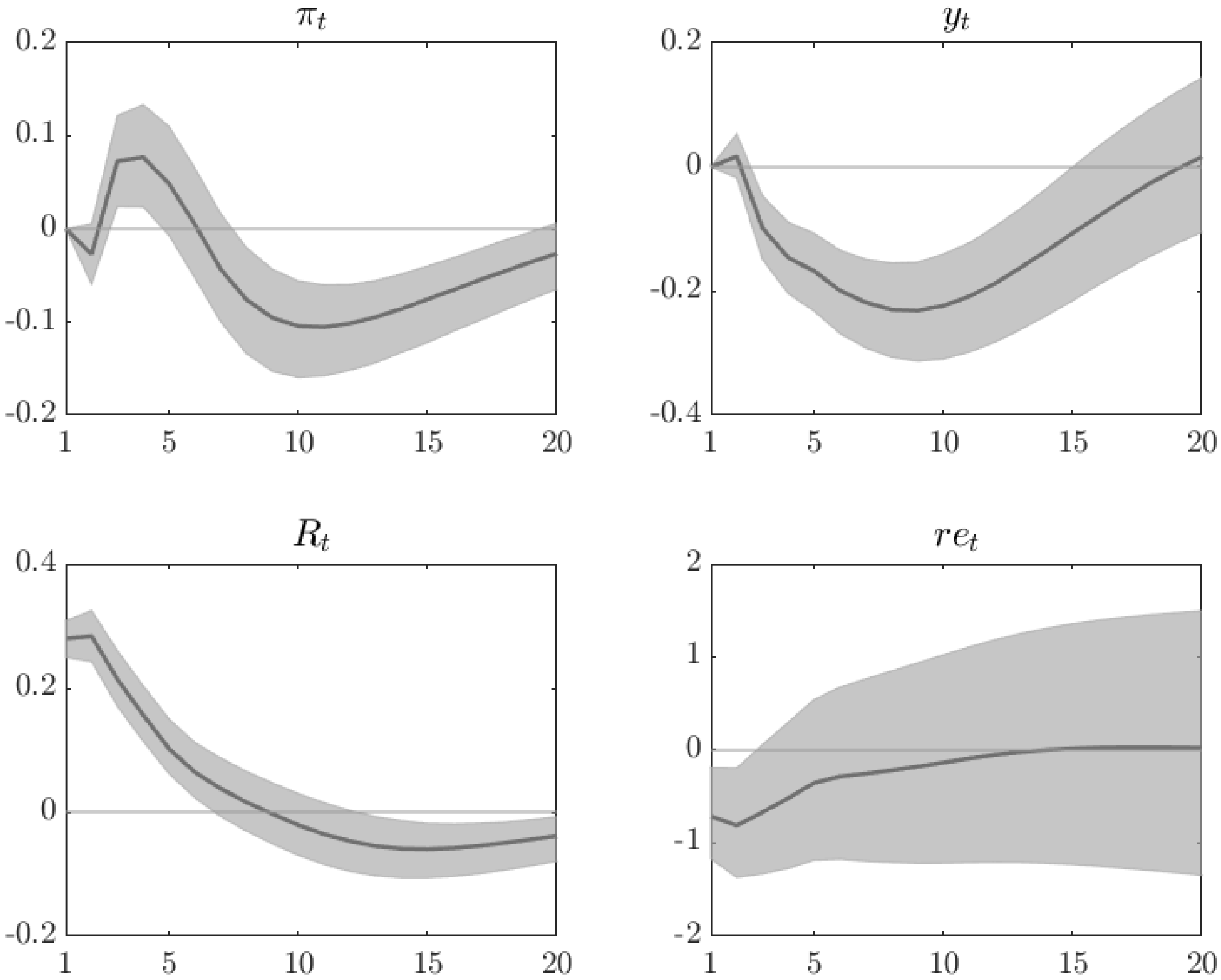

Figure 21: Sweden – impulse responses to a monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

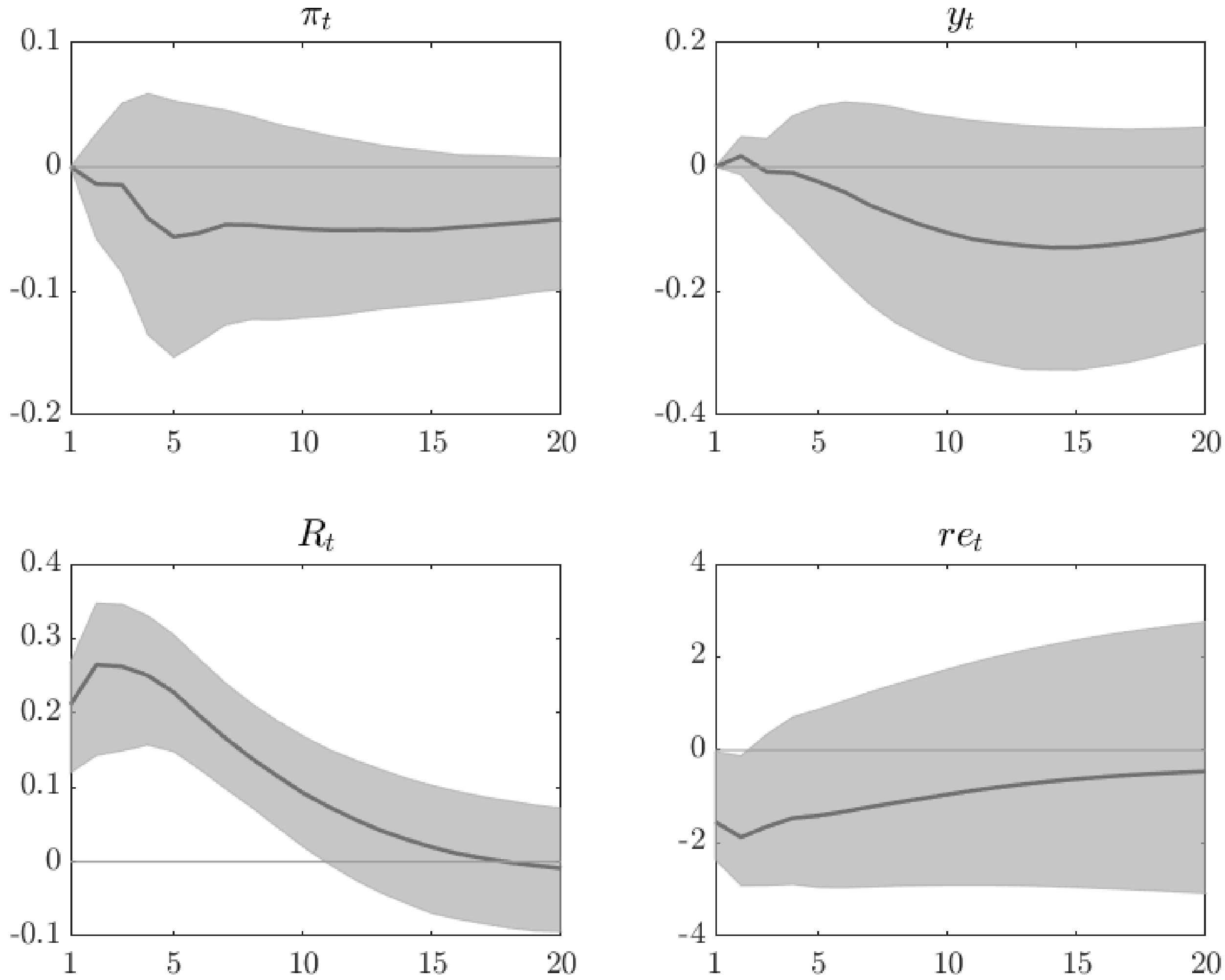

Figure 22: UK – impulse responses to a monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

C Robustness Results

C.1 Impulse Responses

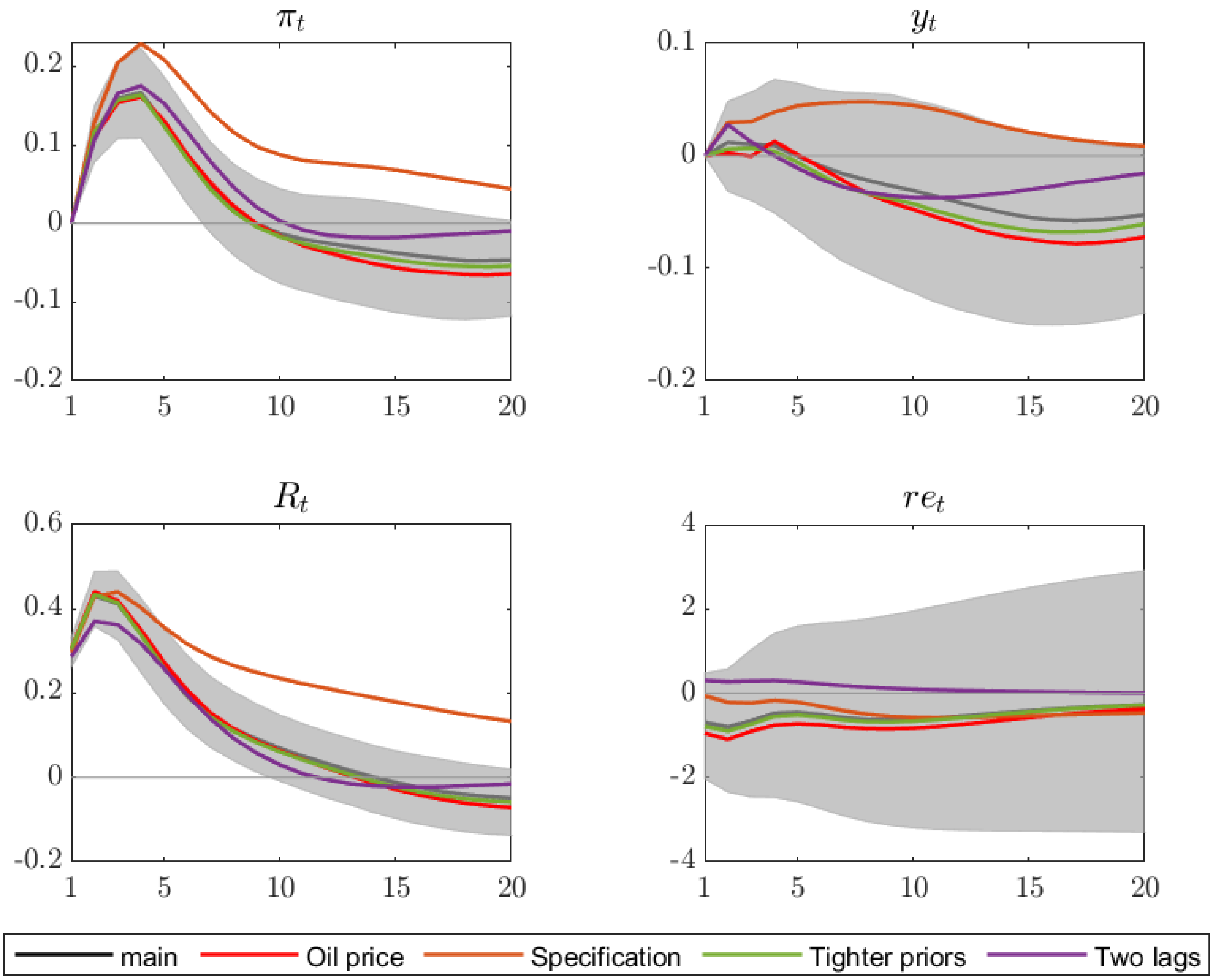

Figure 23: Australia – impulse responses to a monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

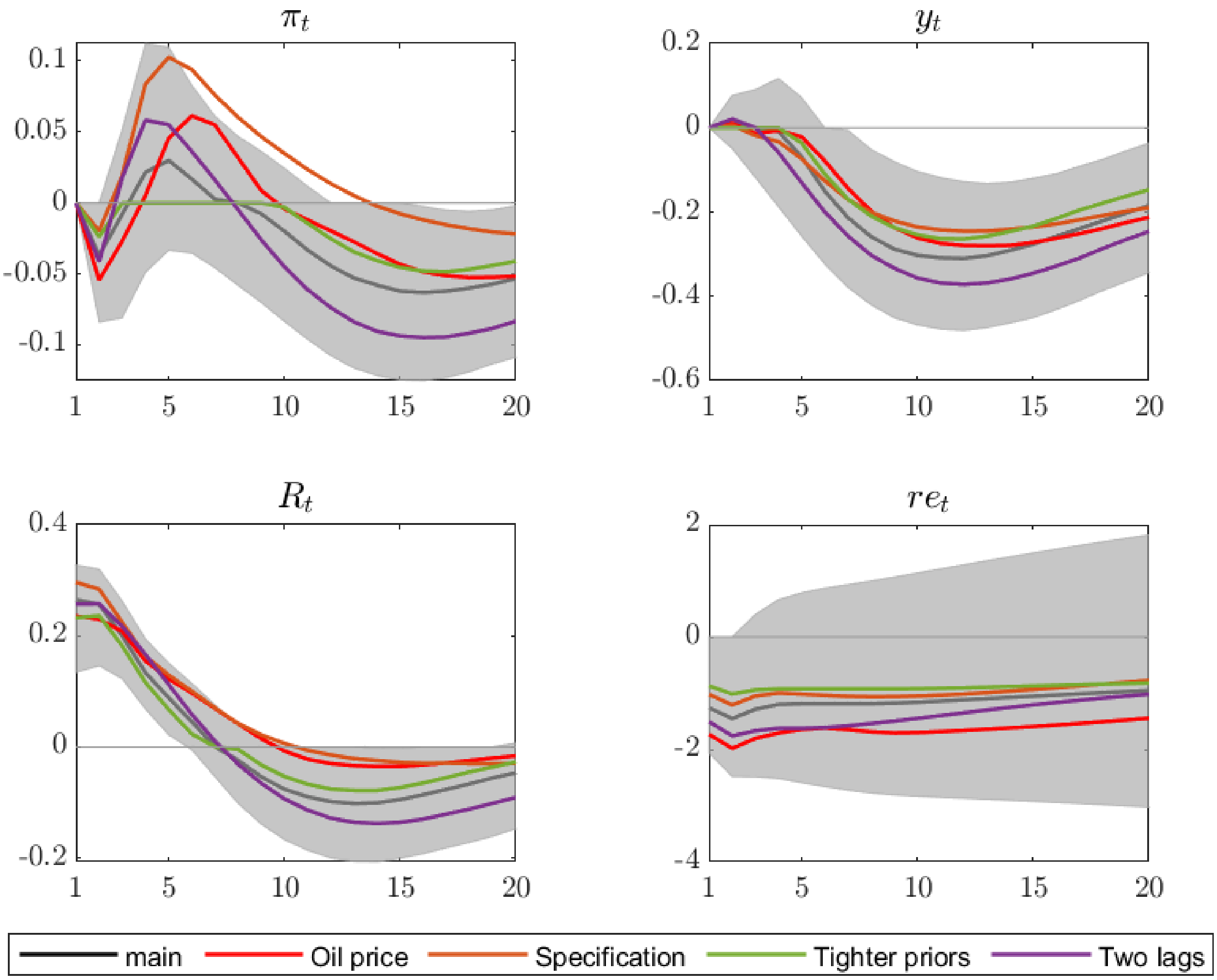

Figure 24: Canada – impulse responses to a monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

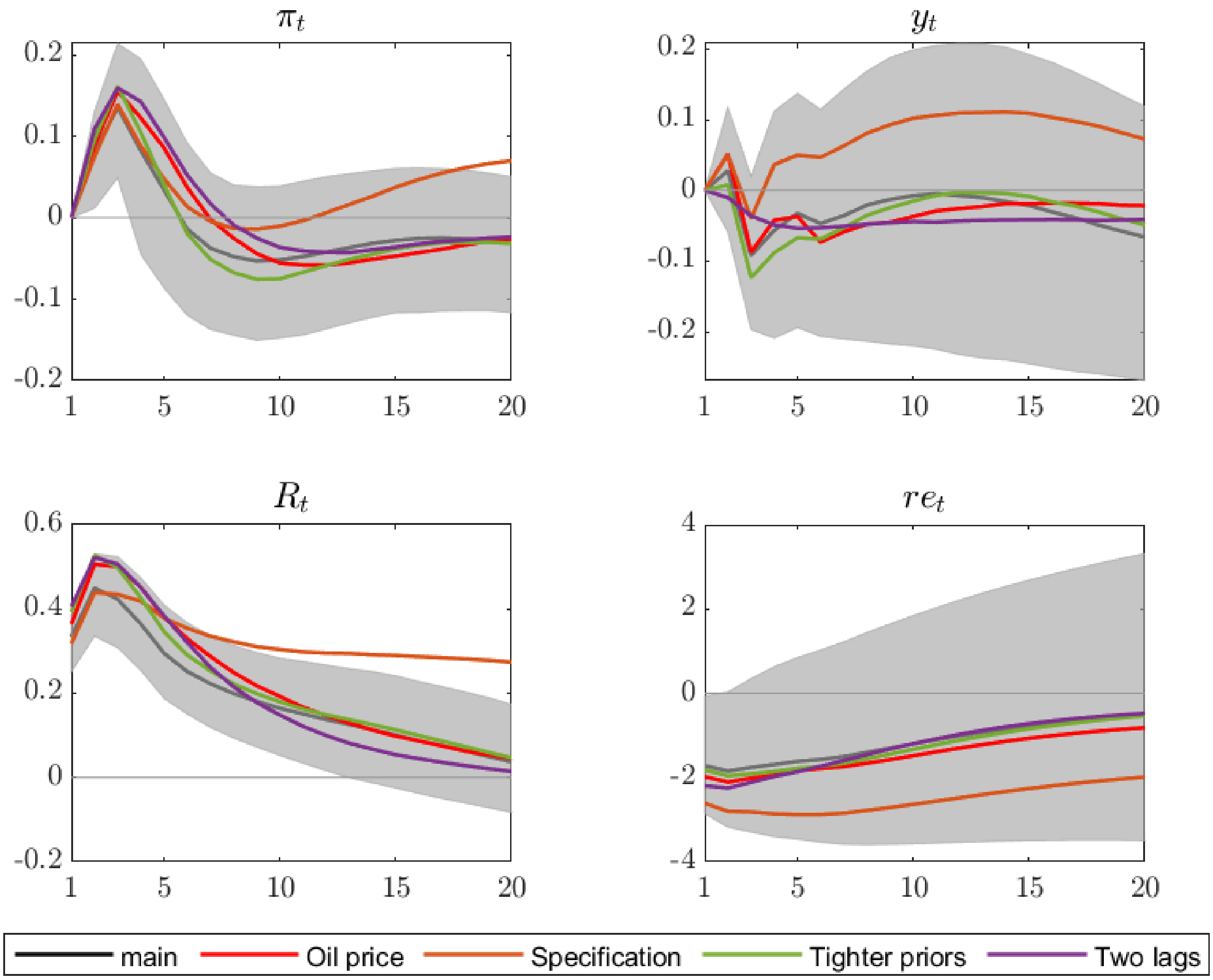

Figure 25: New Zealand – impulse responses to a monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

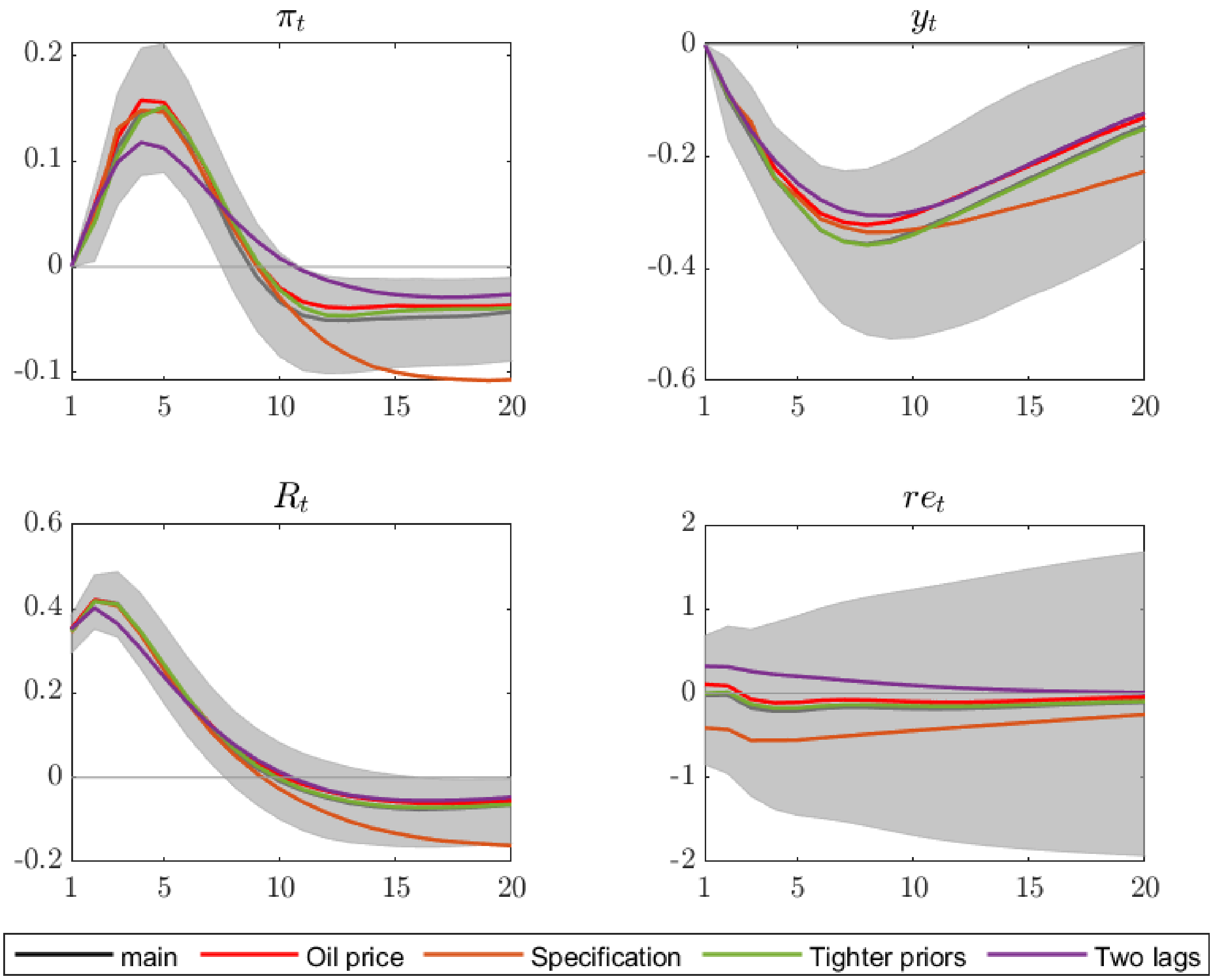

Figure 26: Norway – impulse responses to a monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

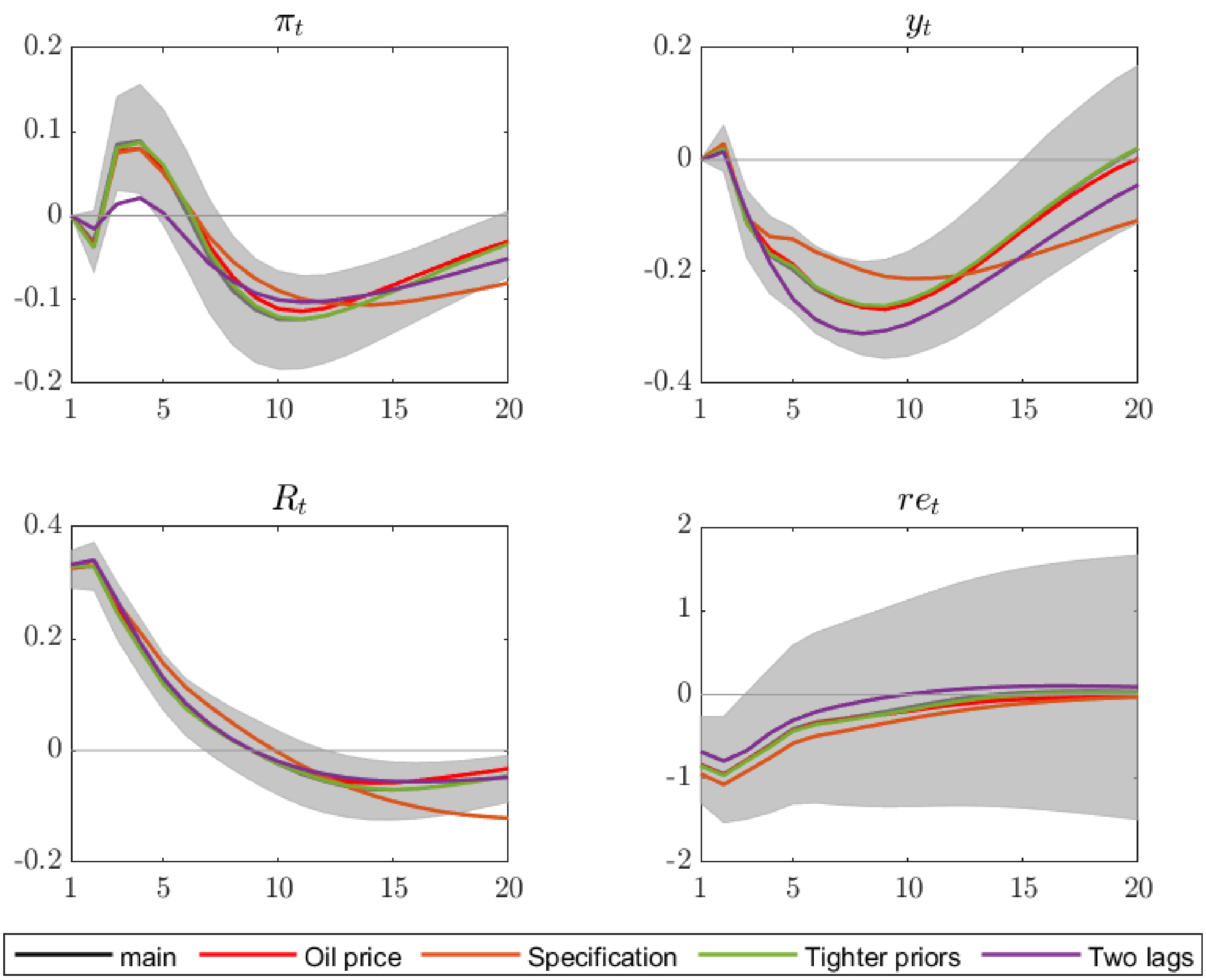

Figure 27: Sweden – impulse responses to a monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

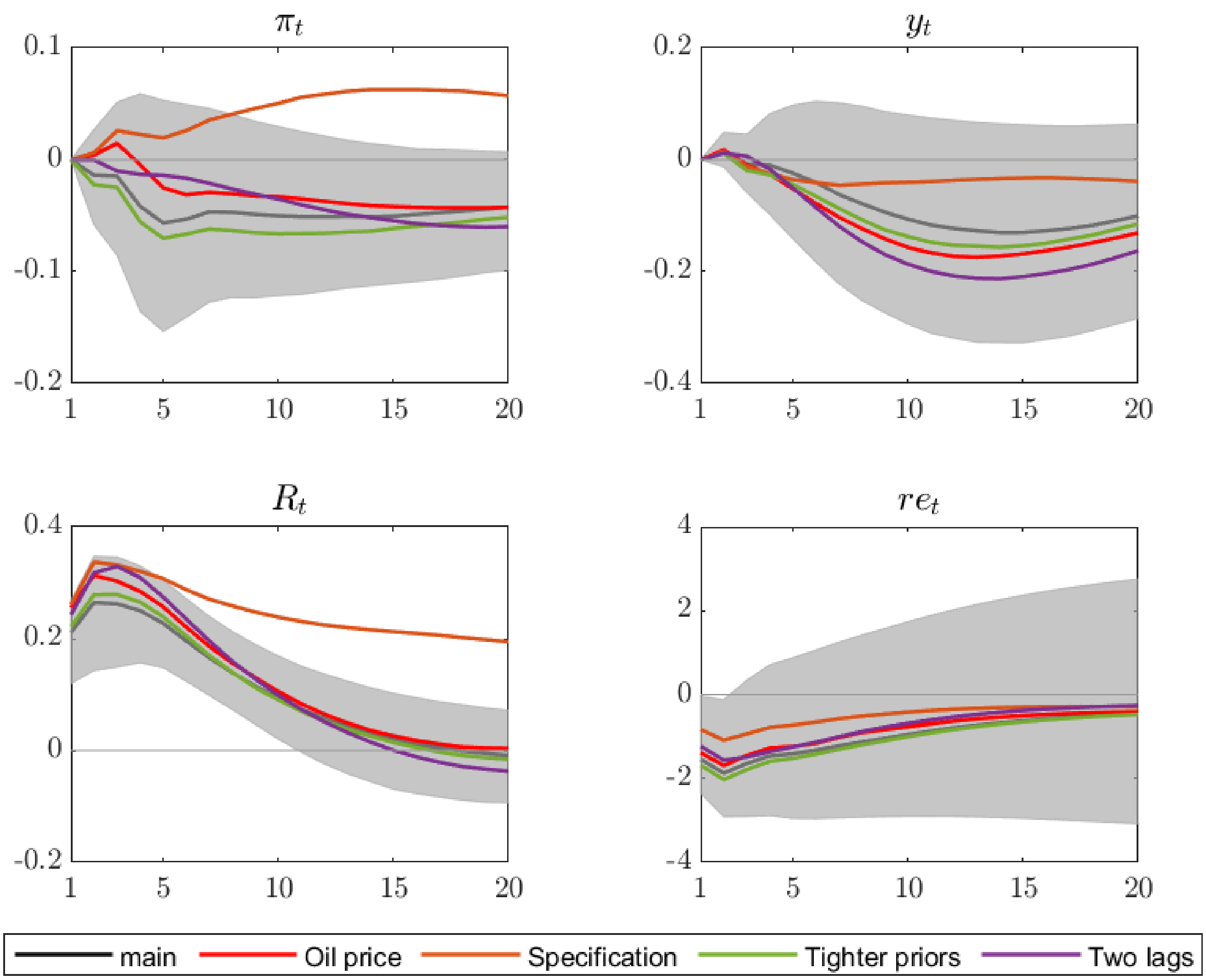

Figure 28: UK – impulse responses to a monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

C.2 Conditional Excess Returns

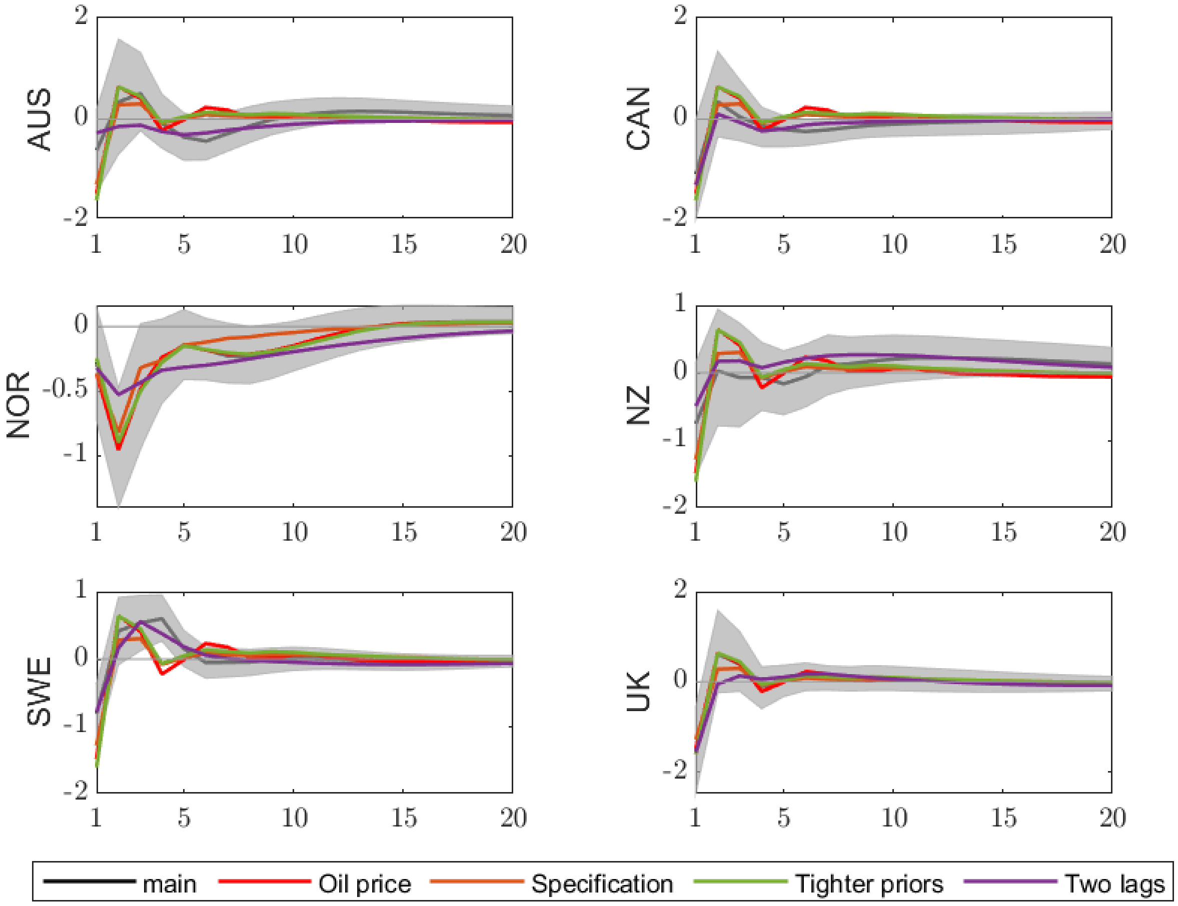

Figure 29: Robustness checks for conditional excess returns.

C.3 Forecast Error Variance Decompositions

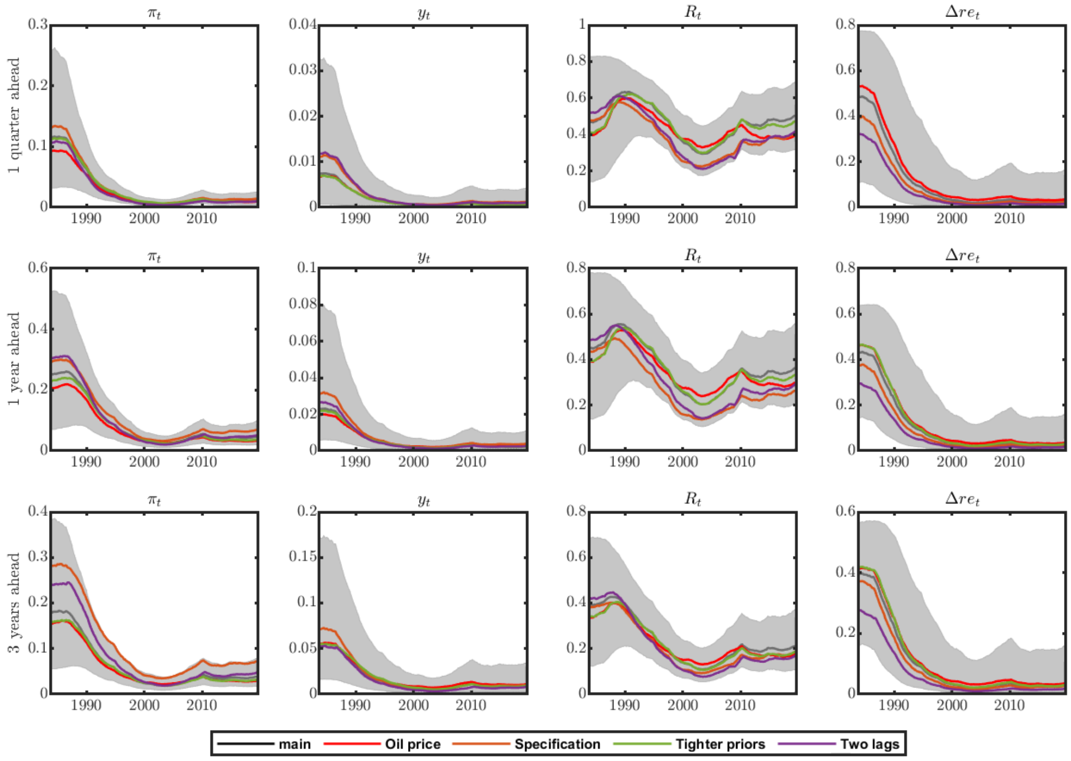

Figure 30: Australia – FEVD for monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

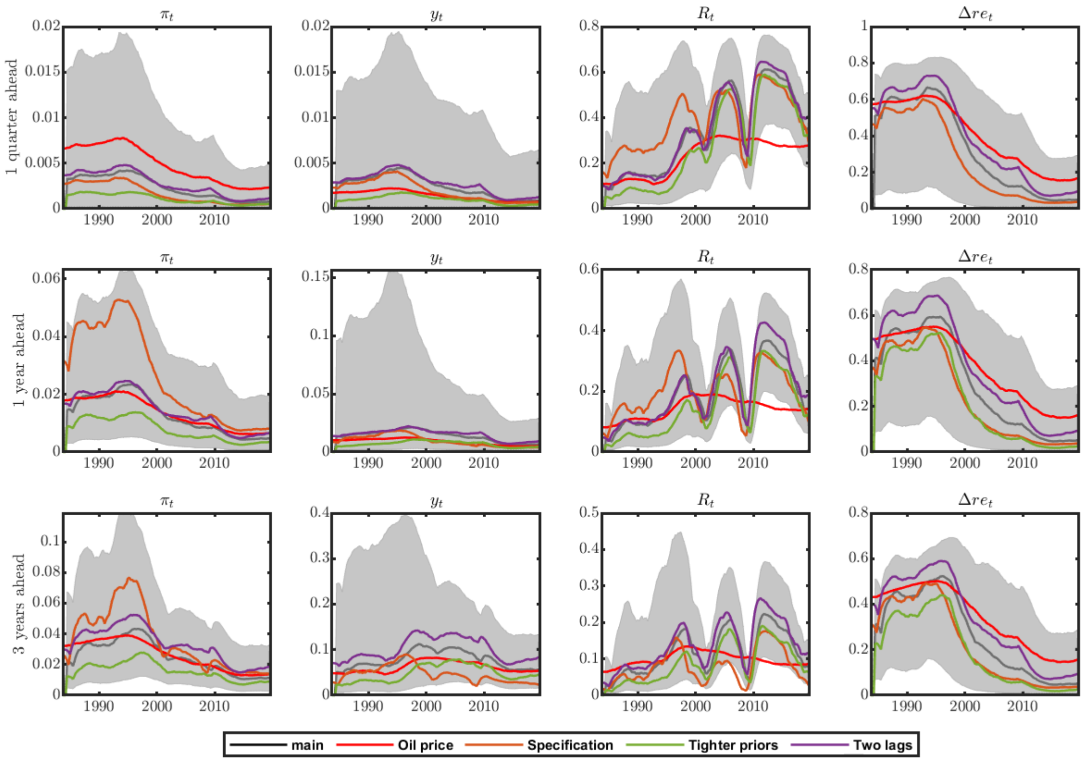

Figure 31: Canada – FEVD for monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

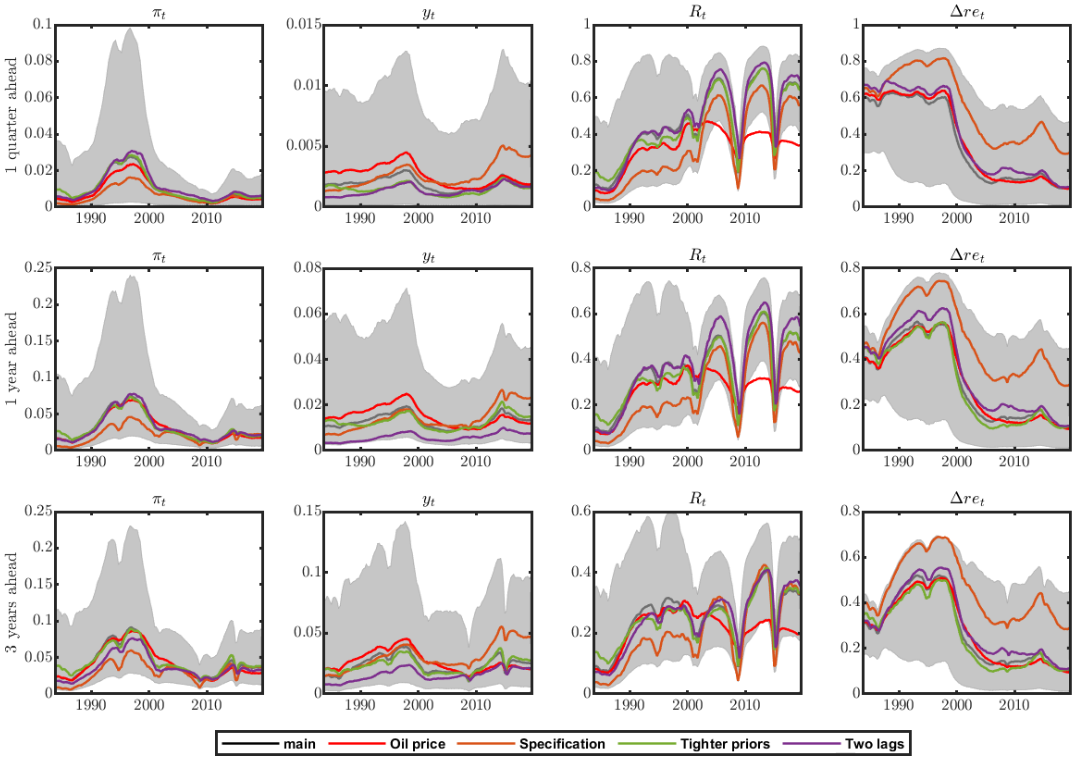

Figure 32: New Zealand – FEVD for monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

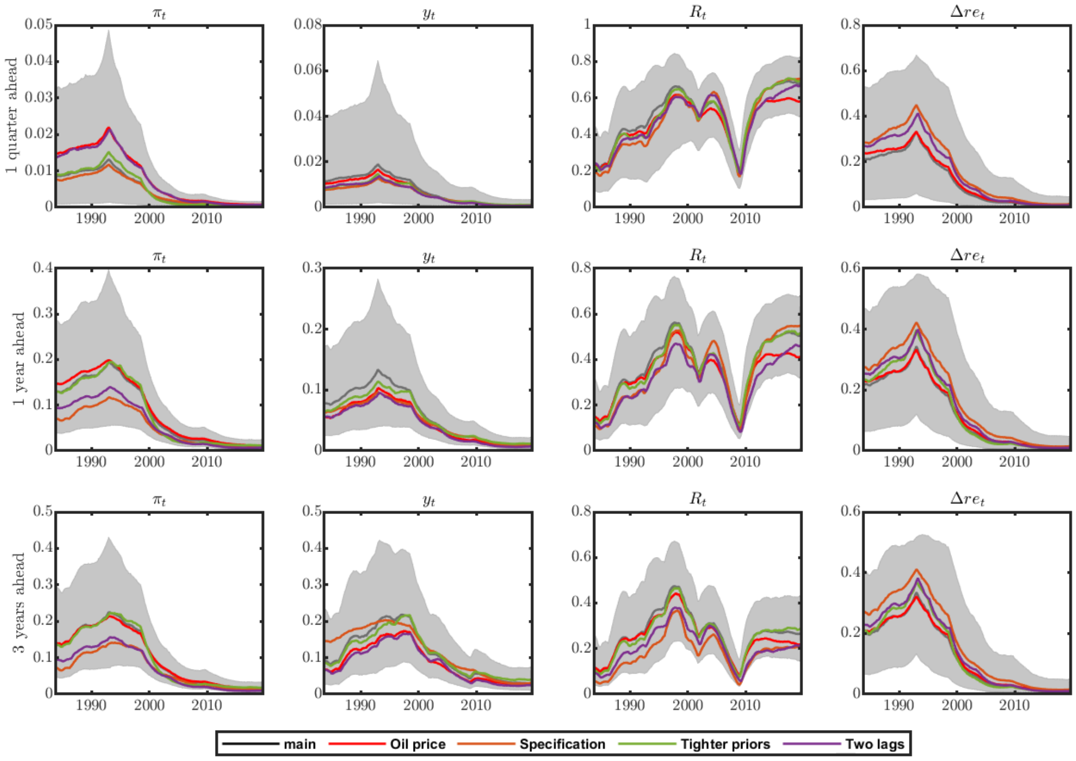

Figure 33: Norway – FEVD for monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

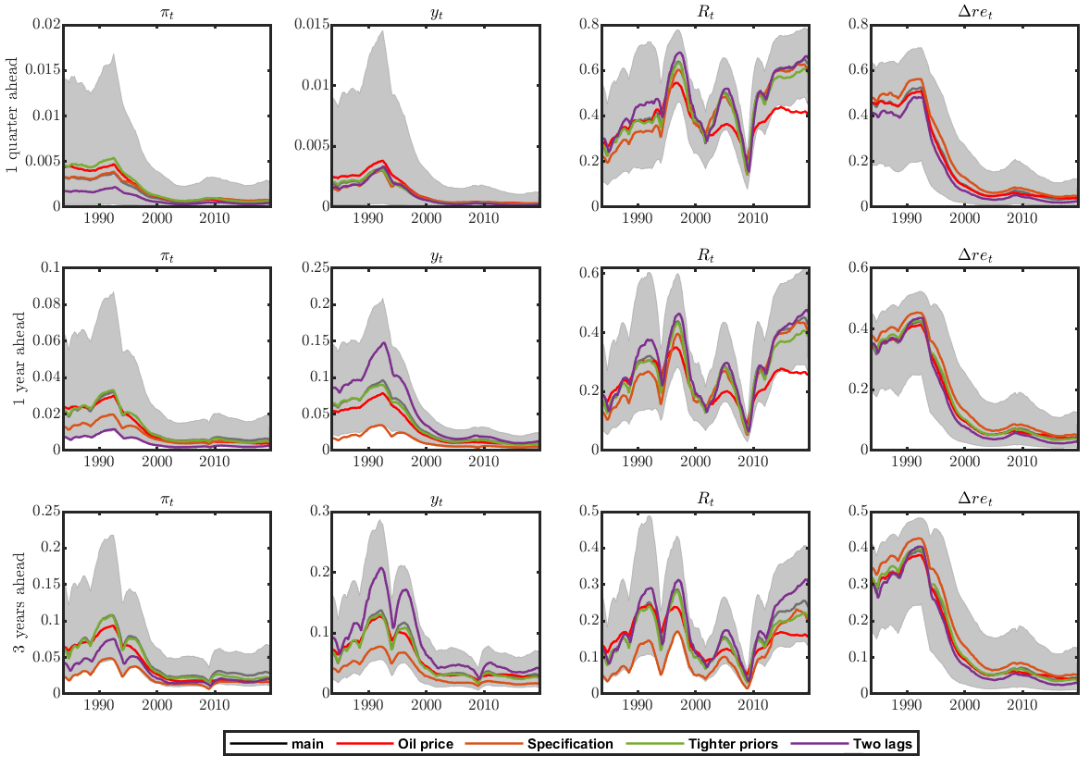

Figure 34: Sweden – FEVD for monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

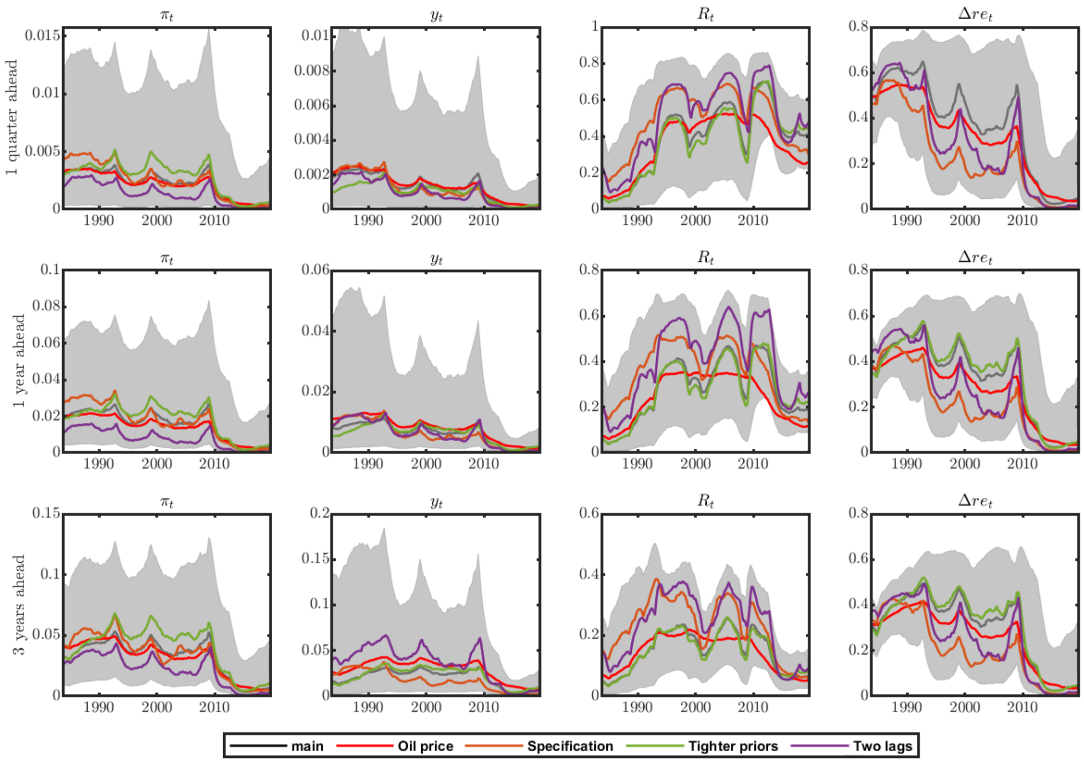

Figure 35: United Kingdom – FEVD for monetary policy shock from the VAR-SV model. Solid lines depict the posterior median estimates while the gray shaded area represents 68% posterior credible intervals around the posterior median.

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