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Do central banks respond to exchange rate movements? A structural investigation ☆

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Abstract

We estimate a small-scale, structural general equilibrium model of a small open economy using Bayesian methods. Our main focus is the conduct of monetary policy in Australia, Canada, New Zealand and the UK. We consider generic Taylor-type rules, where the monetary authority reacts in response to output, inflation, and exchange-rate movements. We perform posterior odds tests to investigate the hypothesis whether central banks do target exchange rates. The main result of this paper is that the central banks of Australia and New Zealand do not, whereas the Bank of Canada and the Bank of England do include the nominal exchange rate in its policy rule. This result is robust for various specification of the policy rule. We also find that terms-of-trade movements do not contribute significantly to domestic business cycles.

JEL classification: C32; E52; F41

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1. Introduction

The New Keynesian framework has been the focus of much recent research on the theory and practice of monetary policy. While not an unqualified empirical success, its parsimony and theoretical consistency lends itself easily to theoretical and empirical policy analysis. Recently, this framework has been applied to study monetary policy in the open economy. An important question in this area is to what extent central banks respond to exchange rate movements when setting monetary policy (see Taylor, 2001). We address this issue by estimating a dynamic stochastic general equilibrium (DSGE) model of a small open economy (SOE) for several countries that potentially differ in their approaches to and experiences with monetary policy.

Our theoretical framework is based on Galí and Monacelli (2005) who extend the benchmark New Keynesian DSGE model described, for instance, in Woodford (2003) to a SOE setting. Open economies can engage in intertemporal as well as intratemporal trade for the purposes of smoothing consumption above and beyond what is possible in a closed economy. At the same time, foreign shocks, such as the terms of trade, can alter domestic business cycle fluctuations which may lead the monetary authority to explicitly take into account international variables. Like its closed-economy counterpart, the model consists of a forward-looking (open economy) IS-equation and a Phillips curve relationship. The former is derived from a consumption Euler equation taking into account that households consume not only domestically produced but also imported goods. The latter is obtained from the optimal price setting decisions of domestic producers. Monetary policy is described by an interest rate rule, while the exchange rate is introduced via the definition of the consumer price index (CPI) and under the assumption of purchasing power parity (PPP).

Rather than estimating policy reaction functions in a univariate setting we pursue a multivariate approach by estimating the entire structural model. The full-information likelihood-based approach optimally adjusts the estimation of the policy rule coefficients for the endogeneity of the right-hand-side variables. Moreover, we are able to exploit cross-equation restrictions that link agents' decision rules to the policy parameters. We assign prior distributions to reaction function specifications and the remaining model parameters and conduct Bayesian inference. Posterior probabilities are used to assess the adequacy of various policy rules. While this methodology has been applied to various economic questions before, we believe that our paper is the first to address the issue of open economy policy rules. Consequently, our paper presents a departure from—and a fairly straightforward alternative to—the single equation approach prevalent in the literature. To illustrate the information gain due to the DSGE model's restrictions we compare our model-based estimates to Bayesian instrumental variable estimates.

We apply our estimation technique to four small open economies, Australia, Canada, New Zealand and the UK and focus on the estimates of the monetary policy rule. Australia and Canada are both large natural resource exporters (as is the UK, but to a smaller degree) so that domestic business cycle fluctuations likely to have a substantial international relative price component. Central banks in these countries therefore may have a specific interest in explicitly reacting to and smoothing exchange rate movements as a predictor of domestic volatility. The Bank of Canada specifically acknowledged this point in that it developed a monetary condition index (MCI) that encompasses both interest rate and exchange rate information as a more comprehensive indicator of the monetary stance.

The main empirical finding in this paper is that the central banks of Australia and New Zealand did not explicitly target exchange rates over the last two decades. The Bank of Canada and the Bank of England, on the other hand, did. This finding is robust over different specifications of the monetary policy reaction function. We also find that in our framework the terms of trade have a fairly small impact on domestic fluctuations, which is significantly at odds with most calibrated business cycle models.

The paper is organized as follows. The following section presents a structural SOE model which we proceed to estimate. In Section 3 we discuss our econometric approach and contrast our structural approach with instrumental variable estimation. Section 4 contains our estimation results, and Section 5 concludes.

2. A simple, structural open economy model

Our model is a simplified version of Galí and Monacelli (2005) to which we refer for details on the derivation of the reduced form equations. Like its closed-economy counterpart, the model consists of a forward-looking (open economy) IS-equation and a Phillips curve. Monetary policy is described by an interest rate rule, while the exchange rate is introduced via the definition of the CPI and under the assumption of PPP. Specifically, the evolution of the SOE is determined by the following equations.

The consumption Euler equation can be rewritten as an open economy IS-curve:

$$\begin{aligned} y_t = {} & E_t y_{t+1} - [\tau + \alpha(2-\alpha)(1-\tau)](R_t - E_t\pi_{t+1}) - \rho_z z_t \\ & - \alpha[\tau + \alpha(2-\alpha)(1-\tau)]E_t\Delta q_{t+1} + \alpha(2-\alpha)\,\frac{1-\tau}{\tau}E_t\Delta y^*_{t+1}, \end{aligned} \tag{1}$$

where $0<\alpha<1$ is the import share, and τ the intertemporal substitution elasticity. Notice that the equation reduces to its closed economy variant when $\alpha = 0$. Endogenous variables are aggregate output y_t and the CPI inflation rate π_t. q_t are the terms of trade, defined as the relative price of exports in terms of imports. The terms of trade enter in first difference form since it is *changes* in (relative) prices that affect inflation (and ultimately the real rate) via the definition of the consumption based price index. y^*_t is exogenous world output, while z_t is the growth rate of an underlying non-stationary world technology process A_t. In order to guarantee stationary of the model, all real variables are therefore expressed in terms of percentage deviations from A_t.[1]

Optimal price setting of domestic firms leads to the open economy Phillips curve:

$$\pi_t = \beta E_t\pi_{t+1} + \alpha\beta E_t\Delta q_{t+1} - \alpha\Delta q_t + \frac{\kappa}{\tau + \alpha(2-\alpha)(1-\tau)}(y_t - \bar{y}_t), \tag{2}$$

where $\bar{y}_t = -\alpha(2-\alpha)(1-\tau)/\tau y^*_t$ is potential output in the absence of nominal rigidities. Again, the closed economy variant obtains when $\alpha = 0$. The slope coefficient $\kappa>0$ is a function of underlying structural parameters, such as labor supply and demand elasticities and parameters capturing the degree of price stickiness. Since we do not use any additional information from the underlying model we treat κ as structural.

[1]See Lubik and Schorfheide (2005) for further discussion of such a specification.

In order to study exchange rate policies we introduce the nominal exchange rate e_t via the definition of the CPI. Assuming that relative PPP holds, we have

$$\pi_t = \Delta e_t + (1-\alpha)\Delta q_t + \pi_t^*, \tag{3}$$

where π_t^* is a world inflation shock which we treat as an unobservable.[2]

We assume that monetary policy is described by an interest rate rule, where the central bank adjusts its instrument in response to movements in CPI inflation and output. Moreover, we allow for the possibility of including nominal exchange rate depreciation Δe_t in the policy rule:

$$R_t = \rho_R R_{t-1} + (1-\rho_R)[\psi_1 \pi_t + \psi_2 y_t + \psi_3 \Delta e_t] + \varepsilon_t^R. \tag{4}$$

We assume that the policy coefficients $\psi_1, \psi_2, \psi_3 \geqslant 0$. In order to match the persistence in nominal interest rates, we include a smoothing term in the rule with $0 < \rho_R < 1$. ε_t^R is an exogenous policy shock which can be interpreted as the non-systematic component of monetary policy. Our primary interest is whether monetary authorities include exchange rate terms in their reaction functions. We evaluate this hypothesis by estimating the model separately under the restrictions $\psi_3 > 0$ and $\psi_3 = 0$ and computing a posterior odds ratio for the two specifications.

Instead of solving endogenously for the terms of trade, we add a law of motion for their growth rate to the system:

$$\Delta q_t = \rho_q \Delta q_{t-1} + \varepsilon_{q,t}. \tag{5}$$

This specification is not fully consistent with the underlying structural model. Since firms do have a certain modicum of market power, the prices of internationally traded products are not exogenous to the economy even if its size relative to the rest of the world goes to zero. The terms of trade are thus determined endogenously as the relative price that clears international goods markets. In terms of growth rates this relationship can be written as

$$[\tau + \alpha(2-\alpha)(1-\tau)]\Delta q_t = \Delta y_t^* - \Delta y_t. \tag{6}$$

An increase in world output raises demand for the domestically produced goods so that the terms of trade, i.e. its relative price improve, while a decline in domestic output has the opposite effect.

Estimation of the fully structural model turned out to be problematic, however. For most specifications, our numerical optimization routine had difficulties finding the maximum of the posterior density. Whenever our optimization did converge, we obtained implausible parameter estimates and low likelihood values. The apparent reason is that Eq. (6) implies a tight link between the terms of trade and output growth that the estimation procedure attempts to match. This creates a conflict with output and inflation dynamics as governed by the IS-equation and the Phillips-curve, which can at best only be resolved at the cost of implausible estimates. In other words, the model with fully endogenous terms of trade is too tightly restricted. We therefore decided to implement the model with terms of trade shocks with the added advantage that interpretation of their effects is straightforward.

[2]An alternative interpretation, as in Lubik and Schorfheide (2005), is that π_t^* captures misspecification, or deviations from PPP. Since the other variables in the exchange rate equation are observed, this relaxes the potentially tight cross-equation restrictions embedded in the model.

Eqs. (1)–(6) form a linear rational expectations model. We assume that y_t^* and π_t^* evolve according to univariate AR(1) processes with autoregressive coefficients ρ_{y*} and $\rho_{\pi*}$, respectively. The innovations of the AR(1) processes are denoted by $\varepsilon_{y*,t}$ and $\varepsilon_{\pi*,t}$. The model is solved using the method described in Sims (2002).

3. Estimation strategy and empirical implementation

We proceed with a discussion of our econometric methodology and explain the advantages of our model-based estimation of monetary policy rules. We then describe the construction of the data sets that are used for the empirical work and present our choice of prior distributions for the Bayesian analysis.

3.1. Econometric methodology

This paper focuses on the estimation of the monetary policy rule (4) and in particular the magnitude of ψ_3 which determines the extent to which central banks respond to exchange rate movements. The policy rule cannot be consistently estimated by ordinary least squares because the regressors are endogenous, that is, $\mathbb{E}[\varepsilon_t^R|\pi_t, y_t, \Delta e_t] \neq 0$. System-based estimation methods correct for the endogeneity by adjusting for the non-zero conditional expectation of the monetary policy shock. The monetary policy rule is implicitly replaced by the following equation:

$$\begin{aligned} R_t = {} & \mathbb{E}[\varepsilon_t^R|\pi_t, y_t, \Delta e_t] + \rho_R R_{t-1} + (1-\rho_R)[\psi_1\pi_t + \psi_2 y_t + \psi_3 \Delta e_t] \\ & + (\varepsilon_t^R - \mathbb{E}[\varepsilon_t^R|\pi_t, y_t, \Delta e_t]). \end{aligned} \tag{7}$$

We use the likelihood function associated with the DSGE model discussed in Section 2 to generate the correction term $\mathbb{E}[\varepsilon_t^R|\pi_t, y_t, \Delta e_t]$ and impose all the rational expectations cross-coefficient restrictions to exploit potential efficiency gains.

The policy rule parameters of the DSGE model are collected into the 4×1 vector $\psi = [\psi_1, \psi_2, \psi_3, \rho_R]'$ and the non-policy parameters and the shock standard deviations are stacked in the 13×1 vector θ. Under the assumption that all the structural shocks are normally distributed and uncorrelated with each other at all leads and lags we can obtain a joint probability distribution for the endogenous model variables. In the empirical analysis the vector of observables Y_t will be composed of annualized interest rates, annualized inflation rates, output growth, depreciation rates, and terms of trade changes and assume that the observations are demeaned. The vector of observations is related to the model variables according to

$$Y_t = [4R_t, 4\pi_t, \Delta y_t + z_t, \Delta e_t, \Delta q_t]'.$$

Recall that the model variable y_t is defined as the ratio of output and world productivity A_t. Hence, observed output growth corresponds to Δy_t adjusted by productivity growth z_t.

We adopt a Bayesian approach and place a prior distribution with density $p(\psi, \theta) = p(\psi)p(\theta)$ on the structural parameters. The data are used to update the prior through the likelihood function. We denote the likelihood function associated with the DSGE model by $\mathscr{L}_{\mathrm{D}}(\psi, \theta|Y^{\mathrm{T}})$, where $Y^{\mathrm{T}} = \{Y_1, \ldots, Y_T\}$. According to Bayes Theorem the posterior distribution of the parameters is of the form

$$p_{\mathrm{D}}(\psi, \theta|Y^{\mathrm{T}}) = \frac{\mathscr{L}_{\mathrm{D}}(\psi, \theta|Y^{\mathrm{T}})p(\psi)p(\theta)}{\int \mathscr{L}_{\mathrm{D}}(\psi, \theta|Y^{\mathrm{T}})p(\psi)p(\theta)d(\psi, \theta)}. \tag{8}$$

Draws from this posterior can be generated through Bayesian simulation techniques described in detail in Schorfheide (2000) and An and Schorfheide (2005). The procedure has the advantage that we are not just estimating the policy rule parameters, but also the non-policy parameters which are of independent interest. Moreover, we are able to study the propagation and relative importance of structural shocks through impulse response functions and variance decompositions.

Since we are interested in the hypothesis that central banks do not react systematically to exchange rate movements, we estimate a version $\mathcal{M}_1$ of the DSGE model in which $\psi_3 > 0$ and a second version $\mathcal{M}_0$ in which ψ_3 is restricted to be zero. The posterior odds of $\mathcal{M}_0$ versus $\mathcal{M}_1$ are given by

$$\frac{\pi_{0,T}}{\pi_{1,T}} = \frac{\pi_{0,0}}{\pi_{1,0}} \cdot \frac{p(Y^{\mathrm{T}}|\mathcal{M}_0)}{p(Y^{\mathrm{T}}|\mathcal{M}_1)}. \tag{9}$$

The first factor is the prior odds ratio in favor of $\mathcal{M}_0$. The second term is called the Bayes factor and summarizes the sample evidence in favor of $\psi_3 = 0$. The term $p(Y^{\mathrm{T}}|\mathcal{M}_i)$ is called marginal data density and appears as normalizing constant in the denominator of (8). The logarithm of the marginal data density can be interpreted as maximized log-likelihood function penalized for model dimensionality, e.g. Schwarz (1978).

To assess the role of the cross-coefficient restrictions in the estimation of the policy rule we compare the DSGE model estimates to Bayesian instrumental variable (IV) estimates that do not impose the model restrictions on the law of motion of inflation, output, and exchange rates. We report Bayesian IV estimates instead of classical generalized method of moments (GMM) estimates, e.g. Clarida et al. (1998), to highlight the information about policy parameters generated by the cross-coefficient restrictions of the DSGE model and adjust the IV estimates for the information introduced through the prior distribution $p(\psi)$. Both types of estimates are based on the same prior distribution of the policy parameters and on the assumption that the exogenous shocks are normally distributed.

According to our formulation of the policy rule (4) the central bank responds to deviations of output from the stochastic trend induced by the random walk technology process A_t. In the estimation of the DSGE model, the deviations y_t can be treated as latent variable and the Kalman filter is used to infer y_t based on the observables. On the other hand, the IV estimation requires all right-hand-side variables in (4) to be observable. Hence, for the IV analysis we change the definition of Y_t as follows:

$$Y_t = [4R_t, 4\pi_t, y_t^G, \Delta e_t]'.$$

Thus, we replace output growth by an observable proxy for y_t. In the subsequent empirical analysis, we detrend the observed level of output with the HP-filter to obtain y_t^G. Moreover, to keep the relationship between the endogenous regressors and the IVs fairly parsimonious we exclude the terms of trade changes from Y_t.

We partition $Y_t' = [Y_{1,t}, Y_{2,t}']$, where $Y_{1,t}$ corresponds to the nominal interest rate. Moreover, let $X_t' = [Y_{t-1}', \ldots, Y_{t-p}']'$. The monetary policy rule can be represented as

$$Y_{1,t} = X_t' M \beta_1(\psi) + Y_{2,t}' \beta_2(\psi) + \varepsilon_t^R. \tag{10}$$

Here M selects R_{t-1}, $\beta_1(\psi) = \rho_R$, and $\beta_2(\psi) = [(1-\rho_R)\psi_1, (1-\rho_R)\psi_2, (1-\rho_R)\psi_3]'$. The system is completed by the following reduced form of equations for $Y_{2,t}$:

$$Y_{2,t}' = X_t' \Psi + u_{2,t}'. \tag{11}$$

Let Σ_{22} denote the covariance matrix of $u_{2,t}$ and $\Sigma_{2R} = \mathbb{E}[u_{2,t}\varepsilon_{R,t}]$. The endogeneity correction in this IV setup is provided by

$$\mathbb{E}[\varepsilon_t^R | Y'_{2,t}] = (Y'_{2,t} - X'_t\Psi)\Sigma_{22}^{-1}\Sigma_{2R}, \tag{12}$$

and does not use cross-coefficient restrictions derived from the DSGE model outlined in Section 2. We are essentially exploiting the exclusion restriction built into (4) that lagged inflation rates, output gaps, and exchange rates do not affect current monetary policy.

Define $u_{1,t} = u'_{2,t}\beta_2(\psi) + \varepsilon_t^R$ and let $u'_t = [u_{1,t}, u'_{2,t}]$ with covariance matrix Σ. We combine (10) and (11) to obtain a likelihood function $\mathscr{L}_{\mathrm{IV}}(\psi, \Psi, \Sigma | Y)$ for the IV model by assuming that u_t is normally distributed. We use a prior distribution of the form $p(\psi, \Psi, \Sigma) \propto p(\psi)|\Sigma|^{-(n+1)/2}$, where $\propto$ denotes proportionality, $p(\psi)$ is the same prior that was used in the DSGE model analysis, and $|\Sigma|^{-(n+1)/2}$ is an uninformative prior for Ψ and Σ. As above, the application of Bayes Theorem leads to

$$p_{\mathrm{IV}}(\psi, \Psi, \Sigma | Y^{\mathrm{T}}) = \frac{\mathscr{L}_{\mathrm{IV}}(\psi, \Psi, \Sigma | Y^{\mathrm{T}})p(\psi)p(\Psi, \Sigma)}{\int \mathscr{L}_{\mathrm{IV}}(\psi, \Psi, \Sigma | Y^{\mathrm{T}})p(\psi)p(\Psi, \Sigma)d(\psi, \Psi, \Sigma)}. \tag{13}$$

A Gibbs sampler can be used to generate draws from this posterior distribution.[3]

There is a growing literature highlighting parameter identification problems associated with New Keynesian DSGE models, e.g. Beyer and Farmer (2004), Canova and Sala (2005), and Lubik and Schorfheide (2004, 2005). In some cases, the rational expectations solution of the DSGE model implies that a subset of structural parameters disappear from the reduced form law of motion of the observables, in other cases the estimation objective function may have little curvature in some directions. Straightforward manipulations of Bayes Theorem can be used to show that priors are not updated in directions of the parameter space in which the likelihood function is flat, e.g. Poirier (1998). Hence, in our empirical analysis we will conduct careful comparisons of priors and posteriors to characterize the information extracted from the sample.

In a classical GMM or IV framework weak or lack of identification alters the sampling distribution of estimators and test statistics. As a consequence standard limit theory provides a poor approximation of the sampling distribution and naive large sample confidence intervals are unreliable. An excellent survey of the weak instrument literature is provided by Stock et al. (2002). While the early empirical literature on policy rule estimation and, more generally, on the single-equation estimation of equilibrium relationships did not pay careful attention to the identification problems, more recent papers such as Ma (2002), Dufour et al. (2006), and Nason and Smith (2005) employ identification robust inference procedures to obtain valid confidence sets for DSGE model parameters of interest.

3.2. Data description

We use observations on real output growth, inflation, nominal interest rates, exchange rate changes, and terms of trade changes in our empirical analysis. All data are seasonally adjusted and at quarterly frequencies for the period 1983:1 to 2002:4, except for New Zealand where the sample starts in 1988:1. Most of the series are obtained from the DRI

[3]A technical appendix that describes the posterior simulations in detail is available from the authors upon request.

(Global Insight) International Database. Output growth rates are computed as log differences of GDPR (Australia, New Zealand, UK) and GDPRC (Canada), respectively, and scaled by 100 to convert them into quarter-to-quarter percentages. Inflation rates are defined as log differences of the consumer price indices CPINS and multiplied by 400 to obtain annualized percentage rates. The series TOTNS (Australia, New Zealand, and UK) and TOT (Canada) are converted in log differences (scaled by 100) to obtain percentage changes in the terms of trade. The Reserve Bank of Australia has targeted the Interbank Cash Rate in recent years which is contained in the DRI series RMOCSH. For New Zealand we are using RMBANK which corresponds to the Overnight Interbank Cash Rate. The Bank of England targets the Repo Rate (Base Rate) which is available through the DRI series RM. Finally, we are using the Overnight Money Market Rate obtained from Statistics Canada. Trade weighted nominal exchange rate indices for the four countries were obtained from the International Monetary Fund.[4] We take log differences (scaled by 100) to convert the indices into depreciation rates. Our IV estimation requires observations on output detrended by the level of technology that are not readily available. We use HP-filtered log output (scaled by 100) instead. All series are demeaned prior to estimation.

3.3. Choice of prior

We choose priors for the structural parameters to be estimated based on several considerations. Table 1 provides information about the prior for Canada. Prior distributions are assumed to be independent. Size restrictions on the parameters, such as non-negativity, are implemented either by truncating the distribution or properly redefining the parameters actually to be estimated. Since the solution of the linear rational expectations model may be non-existent or exhibit multiple equilibria, we truncate the joint prior distribution used for the DSGE model estimation at the boundary of the determinacy region.[5] Our non-adjusted benchmark prior assigns approximately 5% probability to indeterminacy. The prior of the policy coefficients in the IV estimation is almost identical to the one used in the DSGE model analysis. Since in the reduced-form IV model the concept of indeterminacy is not well-defined, the prior for the IV estimation is not truncated.

We use fairly loose priors for the parameters of the policy rule. The priors for ψ_1 and ψ_2 are centered at the values commonly associated with the Taylor-rule. Our rule also allows for interest rate smoothing with a prior mean of 0.5 with a standard deviation of 0.20. The prior mean of the exchange coefficient is set at 0.25. The model is parameterized in terms of the steady state real interest rate r, rather than the discount factor β. r is annualized such that $\beta = \exp[-r/400]$. Its mean is chosen to be 2.5% with a large standard deviation. The prior for the slope coefficient κ in the Phillips curve is consistent with values reported in the literature (see, for instance, Rotemberg and Woodford, 1997; Galí and Gertler, 1999; Sbordone, 2002). Its mean is set at 0.5, but we allow it to vary widely. The prior for the preference parameter α, import share, is tightly centered at 0.2. Since the model has a

[4]We are grateful to Alessandro Rebucci for providing us with the data.

[5]Lubik and Schorfheide (2004) estimate the simple closed economy version of the present model allowing for the possibility of indeterminacy and sunspot driven business cycle fluctuations.

Table 1
Prior distributions for Canada

Name	Domain	Density	Benchmark		Alternative	
			P(1)	P(2)	P(1)	P(2)
ψ_1	$\mathbb{R}^+$	Gamma	1.50	0.50	1.50	0.60
ψ_2	$\mathbb{R}^+$	Gamma	0.25	0.13	0.75	0.30
ψ_3	$\mathbb{R}^+$	Gamma	0.25	0.13	0.75	0.30
ρ_R	$[0, 1)$	Beta/uniform	0.50	0.20	0.00	1.00
α	$[0, 1)$	Beta	0.20	0.05		
r	$\mathbb{R}^+$	Gamma	2.50	1.00		
κ	$\mathbb{R}^+$	Gamma	0.50	0.25		
τ	$[0, 1)$	Beta	0.50	0.20		
ρ_q	$[0, 1)$	Beta	0.40	0.20		
ρ_z	$[0, 1)$	Beta	0.20	0.05		
ρ_{y*}	$[0, 1)$	Beta	0.90	0.05		
$\rho_{\pi*}$	$[0, 1)$	Beta	0.80	0.10		
σ_R	$\mathbb{R}^+$	InvGamma	0.50	4.00		
σ_q	$\mathbb{R}^+$	InvGamma	1.50	4.00		
σ_z	$\mathbb{R}^+$	InvGamma	1.00	4.00		
σ_{y*}	$\mathbb{R}^+$	InvGamma	1.50	4.00		
$\sigma_{\pi*}$	$\mathbb{R}^+$	InvGamma	0.55	4.00		

Notes: P(1) and P(2) list the means and the standard deviations for beta, gamma, and normal distributions; the upper and lower bound of the support for the uniform distribution; s and v for the inverse gamma distribution, where $p_{\mathcal{IG}}(\sigma|v,s) \propto \sigma^{-v-1}e^{-vs^2/2\sigma^2}$. The effective prior is truncated at the boundary of the determinacy region.

singularity at $\tau = 1$ as the world output shock disappears from the IS-equation (1) and thus drops out from the system, we restrict $0 < \tau < 1$ with a prior mean of 0.5.[6]

To specify the priors for the exogenous shocks we conduct a pre-sample analysis using data from 1970:1 to 1982:4. We fit an AR(1) process to US CPI inflation in order to set the prior for $\Delta\pi_t^*$: $\rho_{\pi*}$ is centered at 0.8 and $\sigma_{\pi*}$ at 0.55. Priors for the rest-of-world output shock y_t^* are obtained by estimating an AR(1) for the ratio of US GDP to domestic GDP. Point estimates of the autoregressive coefficient range from 0.80 (Australia) to 0.95 (UK). Point estimates for the innovation standard deviation range from 1.2 (Canada) to 1.6 (UK). We center the prior for ρ_{y*} at 0.9 and use 1.5 to center the prior of the standard deviation.

We choose identical priors for the parameters of each model economy with the exception that we allow for country specific variation in the technology and terms of trade processes to capture possibly different macroeconomic histories. The priors for Australia, New Zealand and the UK are reported in Table 2. We fit AR(1) processes to domestic output growth rates. The point estimate is 0.3 for Canada, and slightly negative for the UK and

[6] World output shocks can still influence the economy if they are correlated with the terms of trade. However, we cannot identify the independent contribution of y_t^* since the model imposes no further restrictions. From a theoretical point of view, this is a useful benchmark case. It depends on the assumptions of perfect international risk sharing and the equality of intertemporal and intratemporal substitution elasticities. In this case, the trade balance is identically equal to zero for all time periods, and the economy is isolated from world output fluctuations.

Table 2
Prior distributions for other countries

Name	Domain	Density	Australia		New Zealand		UK	
			P(1)	P(2)	P(1)	P(2)	P(1)	P(2)
ρ_q	[0, 1)	Beta	0.20	0.10	0.50	0.20	0.40	0.20
ρ_z	[0, 1)	Beta	0.20	0.10	0.20	0.10	0.20	0.10
σ_q	$\mathbb{R}^+$	InvGamma	3.50	4.00	4.00	4.00	1.50	4.00
σ_z	$\mathbb{R}^+$	InvGamma	1.50	4.00	1.50	4.00	1.50	4.00

Notes: See Table 1.

Australia. We thus choose prior means for ρ_z of 0.2 in all countries. Our terms of trade data start in 1970 for New Zealand and Australia, and 1980 for Canada and the UK. Hence, we rely on the former for the prior resulting in an autoregressive coefficient of 0.2 for Australia and 0.5 for New Zealand. For Canada and the UK we roughly match the unconditional standard deviation.

4. Estimation results

We begin by fitting our SOE DSGE model to Canadian data and discuss the resulting parameter estimates and implied model dynamics. We then compare the DSGE model-based estimates of the monetary policy rule to the IV estimates. Finally, we re-estimate the DSGE model for Australia, New Zealand, and the UK, and report findings from several robustness exercises.

4.1. DSGE model estimation

The Bayesian estimates of the structural parameters for Canada can be found in Table 3. In addition to 90% posterior probability intervals we report posterior means as point estimates. The results are by and large consistent with the previous literature. We find that the Bank of Canada pursues a moderately anti-inflationary policy ($\psi_1 = 1.30$) and demonstrates concern for output ($\psi_2 = 0.23$) and exchange rate movements ($\psi_3 = 0.14$). There is also a reasonably high degree of interest-smoothing with an estimate of $\rho_R = 0.69$. Although the posterior means do not differ markedly from their priors, the data appear informative as the posterior distributions for the policy parameters are more concentrated.

The estimates of the structural parameters fall within plausible ranges. The preference parameter α is estimated to be lower than the observable Canadian import share. However, it is widely recognized (Lubik and Schorfheide, 2005; Justiniano and Preston, 2005) that this interpretation is tenuous at best. Instead, the estimation procedure attempts to chose α such as to reconcile the volatility of the terms of trade and of CPI inflation in Eq. (3) and obey the cross-coefficient restrictions embedded in Eqs. (1)–(2). Similar reasoning applies to the estimates of the Phillips-curve parameter κ and the intertemporal substitution elasticity τ, which both nevertheless attain plausible values consistent with alternative evidence (e.g. Ostry and Reinhart, 1992). The estimates of the stochastic processes reflect the substantial degree of persistence found in the data, most of which is captured by the

Table 3
Parameter estimation results, Canada (Output rule, $\psi_3 \geqslant 0$)

	Prior				Posterior			
	Benchmark		Alternative		Benchmark		Alternative	
	Mean	90% Interval	Mean	90% Interval	Mean	90% Interval	Mean	90% Interval
ψ_1	1.54	[0.82, 2.27]	1.51	[0.59, 2.42]	1.30	[0.98, 1.60]	1.84	[1.23, 2.43]
ψ_2	0.25	[0.06, 0.43]	0.75	[0.26, 1.20]	0.23	[0.09, 0.36]	0.54	[0.24, 0.82]
ψ_3	0.25	[0.06, 0.44]	0.75	[0.29, 1.22]	0.14	[0.06, 0.21]	0.26	[0.13, 0.40]
ρ_R	0.50	[0.17, 0.82]	0.50	[0.10, 1.00]	0.69	[0.61, 0.77]	0.78	[0.70, 0.86]
α	0.20	[0.12, 0.28]			0.11	[0.06, 0.15]	0.10	[0.05, 0.14]
r	2.50	[0.90, 3.99]			2.52	[0.92, 4.05]	2.49	[0.91, 4.03]
κ	0.50	[0.12, 0.87]			0.32	[0.17, 0.47]	0.29	[0.15, 0.43]
τ	0.50	[0.17, 0.83]			0.31	[0.20, 0.42]	0.29	[0.18, 0.39]
ρ_q	0.40	[0.06, 0.71]			0.31	[0.15, 0.48]	0.30	[0.14, 0.46]
ρ_z	0.20	[0.12, 0.28]			0.42	[0.38, 0.46]	0.43	[0.39, 0.48]
ρ_{y*}	0.90	[0.83, 0.98]			0.97	[0.94, 0.99]	0.97	[0.94, 0.99]
$\rho_{\pi*}$	0.80	[0.64, 0.95]			0.46	[0.34,0.58]	0.44	[0.33, 0.56]
σ_R	0.62	[0.27, 1.00]			0.36	[0.29, 0.42]	0.38	[0.31, 0.45]
σ_q	1.90	[0.80, 2.99]			1.25	[1.09, 1.41]	1.25	[1.09, 1.41]
σ_z	1.25	[0.51, 1.95]			0.84	[0.68, 1.00]	0.85	[0.69, 1.01]
σ_{y*}	1.89	[0.79, 2.96]			1.29	[0.74, 1.80]	1.21	[0.73, 1.68]
$\sigma_{\pi*}$	0.69	[0.29, 1.09]			2.00	[1.74, 2.26]	2.00	[1.74, 2.25]

high degree of autocorrelation in technology growth ($\rho_z = 0.42$) and the foreign demand shock ($\rho_{y*} = 0.97$).

In order to gauge the importance of the individual shocks we compute variance decompositions. The results are reported in Table 4. Canadian GDP is largely driven by the technology shock and to a lesser degree by (latent) world output. The contribution of monetary policy innovations is slightly below 10% which is in line with evidence from VAR studies. Interestingly, foreign output shocks also contribute significantly to inflation and interest rate volatility. This is likely the outcome of model misspecification as the unobserved process might also pick up the effects of foreign interest rate movements. Exchange rate movements on the other hand are largely determined by foreign inflation, and to a smaller degree by the terms of trade. If the latent variable π_t^* in Eq. (3) is interpreted as measurement error designed to capture deviations from PPP (as in Lubik and Schorfheide, 2005), then we can conclude that the model is able to explain roughly 40% of Canadian exchange rate movements.

Although the terms of trade do not play a substantial role in domestic business cycles, the fact that they explain 20% of the exchange rate would indicate support for the view that the Bank of Canada responds to exchange rates to smooth the impact of international relative price movements. On a final note, it is worth pointing out that the minor role of the terms of trade stands in contrast to much of the international real business cycle literature. For instance, in a calibration analysis of a much richer framework Mendoza (1995) attributes up to 50% of domestic GDP fluctuations to the terms of trade, while Kose (2002) even presents evidence for 90%. On the other hand, Lubik and Teo (2005) in an

Table 4
Variance decomposition, Canada (Output rule, $\psi_3 \geqslant 0$)

	Output	Inflation	Interest rate	Exchange rate
Policy	0.10	0.10	0.08	0.02
	[0.06, 0.13]	[0.03, 0.17]	[0.02, 0.14]	[0.01, 0.02]
Terms of trade	0.005	0.002	0.007	0.21
	[0.001, 0.008]	[0.000, 0.005]	[0.001, 0.013]	[0.14, 0.27]
Technology	0.71	0.09	0.08	0.01
	[0.63, 0.79]	[0.03, 0.15]	[0.01, 0.13]	[0.008, 0.023]
World output	0.18	0.78	0.84	0.14
	[0.11, 0.26]	[0.66, 0.93]	[0.73, 0.97]	[0.05, 0.22]
World inflation	0.011	0.024	0.001	0.63
	[0.002, 0.019]	[0.003, 0.046]	[0.000, 0.003]	[0.53, 0.73]

Notes: The table reports posterior means and 90% probability intervals (in brackets) based on the benchmark prior.

estimated international real business cycle model find that the explanatory power of the terms of trade is below 10%.

The model dynamics can be further studied by computing impulse response functions, which are selectively reported in Fig. 1. The model contains only weak endogenous propagation so that the shape of the responses mirrors those of the underlying shock. Contractionary monetary policy appreciates the currency and lowers inflation and output. An improvement in the terms of trade raises output and lowers inflation on impact via a nominal appreciation. The decline in the exchange rate prompts the central bank to loosen policy, which has an additional expansionary effect on production. Since we assume that technology is difference stationary productivity innovations have permanent effects on output. Positive technology shocks lower inflation and interest rates and thereby appreciate the currency.

The behavior of the economy with respect to demand shocks from the rest of the world deserves special mention. Domestic output declines along with an increase in inflation and the exchange rate. This result arises since world output shocks lower domestic potential output under the estimate of $\tau = 0.31$ (see Eq. (2)). The subsequent 'excess demand' stimulates inflation and leads the central bank to raise nominal rates. The expansionary effect of a foreign demand shock on output is not strong enough to compensate for the contractionary policy. However, this pattern depends crucially on the value of the coefficient of relative risk aversion τ. Since it is below unity, domestic and foreign goods are substitutes, which implies countercyclicality of domestic and world output. The dependence of aggregate dynamics on a single preference parameter can be easily broken in a richer modelling framework. Shocks to import price inflation appreciate the currency, but raise inflation since the central bank reacts to movements in the exchange rate and subsequently relaxes policy.

We now address the question whether the Bank of Canada consistently responded to exchange rate movements over the sample period. We re-estimate the model under the restriction $\psi_3 = 0$.[7] The parameter estimates are virtually identical as are the variance

[7]For the sake of brevity we do not report parameter estimates. They are available upon request.

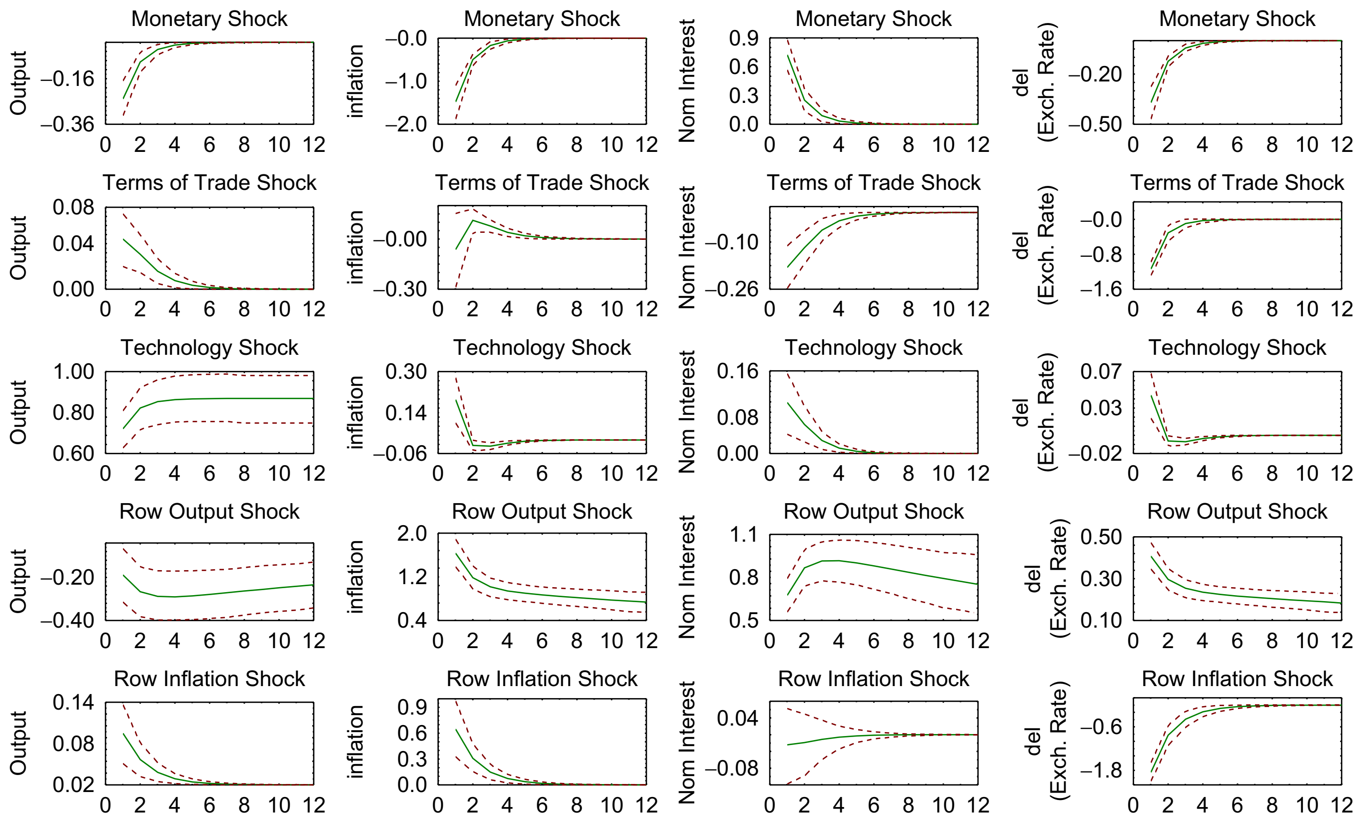

Fig. 1. Impulse responses, Canada (Output rule, $\psi_3 \geqslant 0$). Figure depicts posterior means (solid lines) and pointwise 90% posterior probability intervals (dashed lines) for impulse responses of output, inflation, and exchange rate changes to one-standard deviation structural shocks.

Table 5
Posterior odds

Country	Log marginal	Data densities	Odds
	$\psi_3 = 0$	$\psi_3 > 0$	
Output rule			
Canada (benchmark prior)	−705.37	−702.85	0.0802
Canada (alternative prior)	−708.00	−708.07	1.0705
Australia	−867.17	−870.49	27.644
New Zealand	−889.47	−896.02	702.34
UK	−795.30	−792.10	0.0409
Output gap rule			
Canada	−756.83	−748.06	0.0002
Australia	−912.46	−914.62	8.6659
New Zealand	−896.44	−900.72	72.561
UK	−835.26	−826.48	0.0002

Notes: The table reports posterior odds of the hypothesis $\psi_3 = 0$ versus $\psi_3 > 0$, assuming that the prior odds are one.

decomposition and impulse response functions. An exception is the response to foreign inflation. Since an innovation in π^* feeds directly into the nominal exchange rate, including the latter in the policy rule opens another transmission channel. In the absence of an exchange rate response, a domestic appreciation absorbs the inflation shock while leaving the other variables unaffected.

We can now assess the hypothesis $\psi_3 = 0$ against the alternative $\psi_3 > 0$ by computing the posterior odds ratio. The results are reported in Table 5. The marginal data density of the benchmark model is 2.5 larger on a log-scale which translates into a posterior odds ratio of almost zero. This leads us to favor a consistent Canadian exchange rate response. This result is in line with the notion that the Bank of Canada pays close attention to exchange rate movements on account of its pioneering use of a MCI in gauging nominal demand pressures.[8]

4.2. Bayesian IV estimation

Insights on the importance of using model-embedded cross-equation restrictions can be gleaned from Fig. 2. The panels depict draws from the prior and posterior distributions of the policy parameters for the benchmark DSGE model (upper panels) and the IV specification (lower panels). Visual inspection reveals that the priors are widely dispersed around the respective means, whereas in particular the DSGE model-based posteriors are more concentrated. In other words, the data are informative with respect to the parameters. This is particularly evident for the exchange rate and persistence coefficients.

[8]An MCI is a weighted average of interest rates and exchange rates. It is based on the idea that interest rate and exchange rate movements affect domestic demand via different transmission channels. The presumed advantage of an MCI is that it presents a broader picture of pressures on the economy than the nominal interest rate does. Including the exchange rate in the policy instrument may give a central bank earlier warning of future inflationary conditions. The Bank of Canada has been using the MCI on and off as an operational target of policy over the sample period. Further discussion can be found in the working paper version.

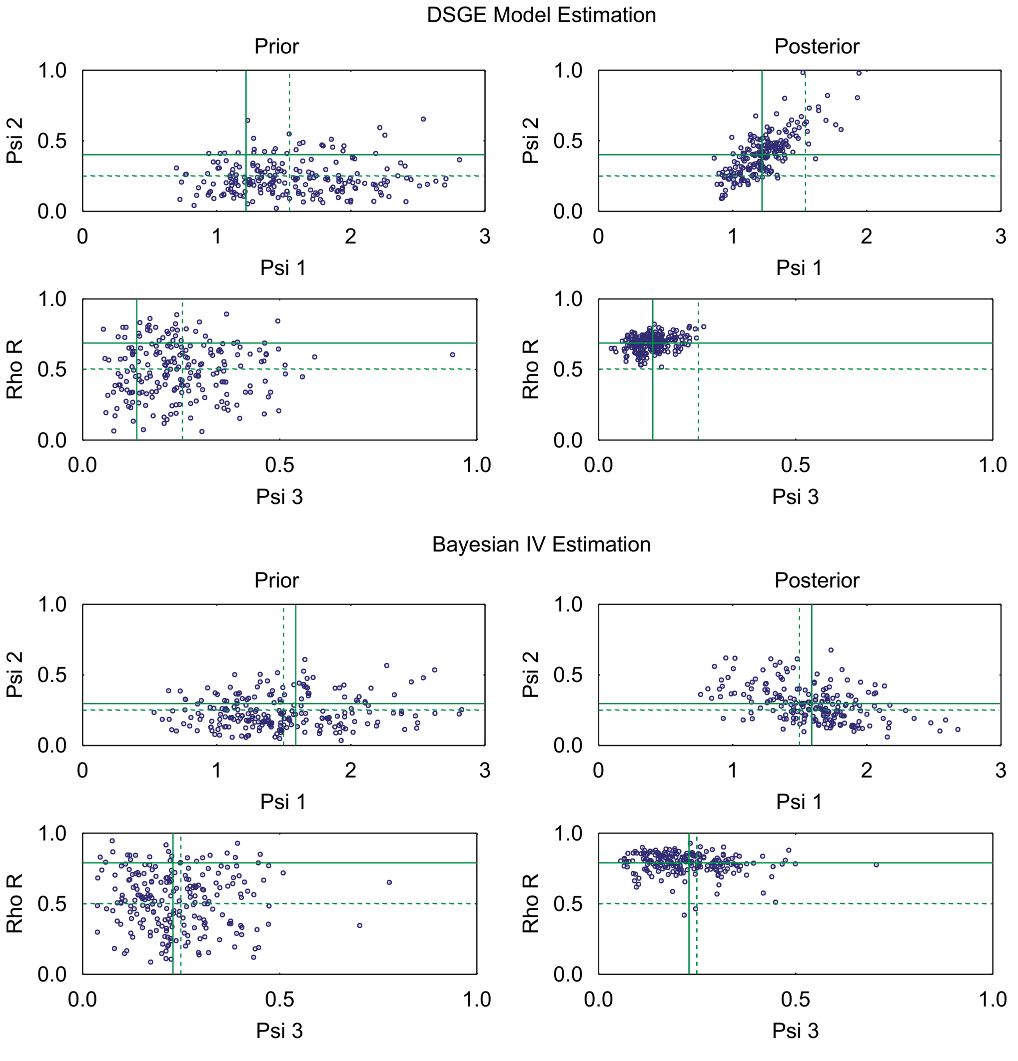

Fig. 2. Output rule, benchmark prior, Canada. The panels depict 200 draws from prior and posterior distributions. Intersections of lines signify prior (dashed) and posterior (solid) means.

The concentration occurs since the model provides restrictions on the volatility and comovement of the variables to be consistent with the data. For instance, a too aggressive inflation response ($\psi_1 > 2$) would imply highly volatile interest rate and inflation dynamics. This induces output volatility via the IS-curve Eq. (1) that conflicts with the volatility implied by the Phillips-curve Eq. (2). Consequently, the posterior assigns very low density to this part of the parameter space.

These cross-equation restrictions are not present in the Bayesian IV estimation. Instead we solely rely on exclusion restrictions implied by the specific form of the monetary policy rule to correct for the endogeneity of the regressors. The IV estimates differ slightly from the benchmark and have larger standard deviations. Notably, the coefficients on inflation

Table 6
Parameter estimation results (Output rule, $\psi_3 \geqslant 0$)

	Australia		New Zealand		UK	
	Mean	90% Interval	Mean	90% Interval	Mean	90% Interval
ψ_1	1.41	[1.04, 1.77]	1.69	[1.24, 2.13]	1.30	[0.96, 1.62]
ψ_2	0.24	[0.09, 0.39]	0.25	[0.13, 0.37]	0.20	[0.07, 0.32]
ψ_3	0.07	[0.03, 0.12]	0.04	[0.01, 0.08]	0.13	[0.07, 0.19]
ρ_R	0.76	[0.69, 0.83]	0.63	[0.53, 0.72]	0.74	[0.66, 0.81]

($\psi_1^{\mathrm{IV}} = 1.59$), exchange rate ($\psi_3^{\mathrm{IV}} = 0.23$), and lagged interest rate ($\rho_R^{\mathrm{IV}} = 0.79$) are higher than the DSGE-model based estimates. This is also evident from the lower panels of Fig. 2. However, the posterior of the policy coefficients appear not to be markedly different from the prior. The exception is the coefficient ρ_R^{IV} which picks up the persistence in the interest rate. This suggests that the IV estimation suffers from an identification problem that is overcome by the use of cross-equation restrictions in the DSGE-based estimates. As for the question of the presence of the exchange rate in the policy rule a researcher would come unequivocably to an affirmative conclusion but this need not be the case as the results for the other countries show.

4.3. Other countries

The structural parameter estimates for the other countries in our sample are very similar. A notable exception is the UK where a Phillips-curve parameter estimate of $\kappa = 0.65$ suggests a lower degree of price stickiness than in the other countries. Moreover, the process for the terms of trade is estimated to be less persistent ($\rho_q = 0.09$) and less volatile. A likely reason is that the UK is less of a commodity exporter. The estimates of the policy parameters are reported in Table 6. All countries are found to pursue strict anti-inflationary policies with inflation coefficients ψ_1 ranging from 1.30 (UK) to 1.69 (New Zealand). Similarly, significant emphasis is put on output targeting, with New Zealand being the most aggressive ($\psi_2^{\mathrm{NZ}} = 0.25$). There is also a high degree of interest-smoothing with an average estimate of 0.70.

Estimates of the exchange rate coefficients are lower than for Canada, the lowest being New Zealand with $\psi_3^{\mathrm{NZ}} = 0.04$.[9] The posterior odds reported in Table 5 imply that Australia and New Zealand do not respond to exchange rate movements, whereas there is evidence that the Bank of England raises interest rates in response to exchange rate depreciations.

4.4. Robustness

Since the chief focus of this paper is the structural estimation of open economy policy rules, we assess the robustness of the benchmark results by relaxing the priors on the policy

[9] This may be surprising in light of the Reserve Bank of New Zealand's experiments with MCI-targeting and the openness of the economy, but is consistent with recent evidence of richer modelling frameworks, e.g. Lubik (2005).

parameters. In particular, we impose a uniform prior on the smoothing parameter ρ_R, while increasing the standard deviation of the response coefficients. Since ψ_2 and ψ_3 are restricted to be non-negative, we also raise the prior mean when making the prior more diffuse. The estimates of the response coefficients are all higher than in the benchmark case. This reflects the influence of the increased prior mean in the case of the output and exchange rate coefficients ψ_2 and ψ_3. At the same time, the data are informative as the posterior estimates are clearly pulled away from the prior. Interestingly, the structural parameter estimates remain unchanged which suggests that the restrictions between the policy rule and the structural equations are fairly weak. Under this alternative prior the marginal data densities reported in Table 5 deteriorate. While the in-sample fit improves slightly the more diffuse prior relaxes some of the parameter restrictions and leads to a larger penalty for model complexity. It turns out that the increase in the penalty term dominates and causes the marginal data density to fall. Under the alternative prior the posterior odds are essentially one suggesting that the data provide neither evidence in favor nor against the hypothesis that the Bank of Canada responds to exchange rates.

A second robustness check concerns the specification of the monetary policy rule. We re-estimated the model under an output gap rule where the central bank responds to deviations of actual from potential output $\bar{y}_t$, where potential output here is the level of output that would prevail in the absence of nominal rigidities in the domestic economy. While this specification has some theoretical appeal, it is unlikely to have been followed in practice. Parameter estimates reveal subtle differences compared to the benchmark specification; in particular, higher policy coefficients overall, and estimates of the preference parameters α and τ that are small and not entirely plausible as the model attempts to minimize the effects of foreign output movements on potential output. This suggests that the restriction imposed by the output gap does not hold in the data. This is confirmed by the marginal data densities which are considerably lower for all countries than the output rule. It is interesting, however, that the previous conclusions regarding policy behavior remain unaffected: the central banks of Canada and the UK respond to the exchange rate, while the Reserve Banks of Australia and New Zealand do not.

We also estimated the model under an expected inflation rule. We found that for all four countries this resulted in a worse model fit, as measured by the marginal data densities, than the benchmark specification. The ranking of policy rules with respect to exchange rate targeting remained unaffected. We obtained similar results when we used an MCI-based policy rule which introduces additional persistence in the model. This specification was rejected in favor of the benchmark. The working paper version of this paper discusses this and further robustness checks in more detail.

5. Summary and concluding remarks

We specify and estimate a small-scale, structural general equilibrium model of an open economy using Bayesian methods. Our main finding is that the central banks of Australia and New Zealand do not respond to exchange rate movements, whereas the Bank of Canada and the Bank of England do. This result is robust to various alternative specifications. This is not to say that the exchange rate is not part of the decision-making process in Australia and New Zealand. Openness changes the structure of the economy and its reaction to monetary policy. However, we do not find evidence that central banks alter their interest rate instrument systematically in response to depreciations. Moreover, we emphasize the methodological point that a fully-specified DSGE model provides cross-equation restrictions that can help overcome the identification problems often encountered in single-equation studies. We illuminate this issue by contrasting our structural estimates with those obtained from an IV procedure.

Our results have to be qualified with respect to the structural model employed as it may be misspecified. One issue is our assumption of exogenous terms of trade movements, another is the lack of imperfect pass-through of nominal exchange rate changes into domestic import prices. Overall model misspecification is of concern as it can lead to biased parameter estimates, prevent identification of the true structural parameters and may imply incorrect model selection. Moreover, our finding that the terms of trade play an almost negligible role in aggregate fluctuations is puzzling. This is at odds with studies based on vector autoregression and, in particular, calibration studies. A richer economic environment could reconcile these different results and allow the model to be fit with endogenous terms of trade determination. Our model contains only a very weak endogenous transmission mechanism. Introducing capital accumulation, different production sectors and internationally incomplete asset markets will generate richer model dynamics and a potentially larger role for terms-of-trade fluctuations.

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