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 U.K., as measured by nominal interest rate rules. We consider generic Taylor-type rules, where the monetary authority reacts in response to output, inflation, and exchange-rate movements. We perform posterior odds test to investigate the hypothesis whether central banks do respond to exchange rates. The main result of this paper is that the central banks of Australia, New Zealand and the U.K. do not, whereas the Bank of Canada does include the nominal exchange rate in its policy rule. This result is robust for various specification of the policy rule, among them an MCI-based rule. Additionally, we find that, based on variance decomposition of the estimated model, that terms-of-trade movements do not contribute significantly to domestic business cycles.

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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 do in fact include exchange rates in the process of formulating monetary policy (see Taylor, 2001).

The theoretical literature does not offer a clear-cut answer. Ball (1999) argues that monetary policy should react to exchange rate movements since they affect domestic inflation through a separably identifiable channel than domestic demand and supply shocks. Using a highly stylized model he finds that the central bank should optimally react to exchange rate movements with 1/10th of the weight on inflation. His results are echoed by Svensson (2000) in a richer modelling environment with forward-looking agents and some microfoundations. He cautions, however, that the welfare effects of exchange-rate targeting are small and may even be negative. Clarida, Galí, and Gertler (2001) take this point even further. In a fully specified, dynamic stochastic general equilibrium (DSGE) model they show that the monetary policy problem in an open economy is isomorphic to its closed economy counterpart. The policy objectives of smoothing output and inflation variations remain the same; what changes, however, is the structure of the reduced form. On the other hand, open economy policy rules have been studied empirically by Clarida, Galí, and Gertler (1998). They find that in the major industrialized countries the exchange rate did play a role in setting monetary policy, but its quantitative importance is small.

We address these issues by estimating a simple, structural 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 a straightforward extension of the New Keynesian monetary business cycle model. In its closed-economy variant inflation and output dynamics are jointly determined by a forward-looking IS-curve and a Phillips-curve type relationship. The former explains expected output growth as a function of the real rate of interest, whereas the latter relates the time path of inflation to the output gap. Monetary policy affects aggregate outcomes via a nominal interest rate reaction function.

We apply our estimation technique to four small open economies, Australia, Canada, New

Zealand and the U.K., that potentially differ in their approaches to monetary policy. 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 have a substantial international relative price components. 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 finding in this paper is that the central banks of Australia, New Zealand, and the U.K. did not explicitly respond to exchange rates over the last two decades. The Bank of Canada, on the other hand, did. This finding is robust over different specifications of the monetary policy reaction function, such as expected inflation targeting. The evidence also does not support the view that the central banks of Canada and New Zealand implemented MCI-based rules. In the case of Canada, the data favor strict inflation- and exchange rate-targeting. 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 methodological contribution of our paper is the structural estimation of monetary policy rules in a general equilibrium model of an open economy. 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 allows us to implicitly generate an optimal set of instruments for the coefficients of the reaction function. 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 rule. Our approach allows us to compare both nested and non-nested policy rules such as inflation versus expected inflation targeting. 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.

Our paper relates to the New Open Economy Macroeconomics originated by Obstfeld and

Rogoff (1995) and recently surveyed by Lane (2001). This literature developed micro-founded and optimization-based models that are usable for policy analysis in the open economy. It particularly highlighted the role of the terms of trade in the transmission of business cycles (see Corsetti and Pesenti, 2001). The traditional expenditure-changing effect of a domestic depreciation was found to be complemented by an expenditure-switching effect that could overcompensate the former and be welfare-reducing. However, most of the research in this area is of a theoretical nature. Notable exceptions are Bergin (2002, 2003), Dib (2003), Smets and Wouters (2002) and Ghironi (2000). The latter author estimates reduced-form equations by non-linear least squares, but does not fully incorporate the cross-equation restrictions from the structural model. His approach also does not allow him to fully characterize the stochastic properties of the model, such as variance decompositions. The papers by Bergin are closer to our approach. He uses classical maximum likelihood techniques to estimate a similar structural model for the same set of countries. His focus is on the ability of the theoretical model to adequately represent the data and does not specifically analyze the conduct of monetary policy. Bergin (2002) extends his approach to a two-country world. Smets and Wouters (2002) also use Bayesian technique, but they focus on the predictive ability of a much larger modeling framework for the U.S. and European economies. In a recent paper that builds on our methodology, Del Negro (2003) estimates monetary policy rules for Mexico and finds strong evidence in favor exchange rate targeting after the Peso crisis in 1994.

The paper is organized as follows. The next Section presents the structural small open economy model that we use for estimation. In Section 3 we discuss our econometric methodology. We also present GMM estimates of the open economy policy rule as a benchmark for out structural estimation. Section 4 contains our benchmark estimation results, while we test for the hypothesis of exchange rate targeting in Section 5. Section 6 contains robustness checks and further modifications of the benchmark model, where we focus on the interpretation an MCI as an open economy monetary policy rule. Section 7 concludes and offers suggestions for further research.

2 A Simple, Structural Open Economy Model

Our model is a simplified version of Galí and Monacelli (2002). For details on the derivation of the reduced form equations we refer to this paper. Like its closed-economy counterpart, the model consists of an (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 consumer price index (CPI) and under the assumption of purchasing power parity (PPP).

Introducing open economy features potentially expands the basic model in several dimensions. 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. International exposure may, however, affect an economy in a more indirect way by changing the structural relationships between aggregates.

Specifically, the evolution of the small open economy is determined by the following equations. The open economy IS curve is:

$$\begin{aligned} \widetilde{y}_t &= E_t\widetilde{y}_{t+1} - [\tau + \alpha(2-\alpha)(1-\tau)]\left(\widetilde{R}_t - E_t\widetilde{\pi}_{t+1}\right) - \rho_A dA_t \\ &\quad -\alpha\left[\tau + \alpha(2-\alpha)(1-\tau)\right]E_t\Delta\widetilde{q}_{t+1} + \alpha(2-\alpha)\frac{1-\tau}{\tau}E_t\Delta\widetilde{y}^*_{t+1}, \end{aligned} \tag{1}$$

where $0 < \alpha < 1$ is the import share and $\tau^{-1} > 0$ the intertemporal substitution elasticity. Notice that the equation reduces to its closed economy variant when $\alpha = 0$. Endogenous variables are aggregate output $\widetilde{y}_t$ and the CPI inflation rate $\widetilde{\pi}_t$, while y^*_t is exogenous world output, dA_t is technology growth, and $\widetilde{q}_t$ 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. This is in marked contrast, for instance, to the ad hoc specification in Ball (1999) who assumes that the lagged exchange rate enters in levels.

This form of the IS equation contains a problematic feature in that if $\tau = 1$ the world output shock drops out the system. 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.[1]

[1] 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.

The open economy Phillips curve is:

$$\widetilde{\pi}_t = \beta E_t \widetilde{\pi}_{t+1} + \alpha\beta E_t \Delta\widetilde{q}_{t+1} - \alpha\Delta\widetilde{q}_t + \frac{\kappa}{\tau + \alpha(2-\alpha)(1-\tau)}\left(\widetilde{y}_t - \widetilde{\widetilde{y}}_t\right), \tag{2}$$

where $\widetilde{\widetilde{y}}_t = -\alpha(2-\alpha)\frac{1-\tau}{\tau}\widetilde{y}^*_t$ is potential output in the absence of nominal rigidities and when technology is non-stationary. 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 measuring the degree of price stickiness. Since we do not use any additional information from the underlying model we treat it as structural.

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:

$$\widetilde{\pi}_t = \Delta\widetilde{e}_t + (1-\alpha)\Delta\widetilde{q}_t + \widetilde{\pi}^*_t, \tag{3}$$

where $\widetilde{\pi}^*_t$ is a world inflation shock.

We assume that monetary policy is described by an interest rate rule, where the central bank adjust its instrument in response to deviations of CPI inflation and output from their respective target levels of price stability and potential output. Moreover, we allow for the possibility of including nominal exchange rate depreciation $\Delta\widetilde{e}_t$ in the policy rule:

$$\widetilde{R}_t = \rho\widetilde{R}_{t-1} + (1-\rho)\left[\psi_1\widetilde{\pi}_t + \psi_2\left(\widetilde{y}_t - \widetilde{\widetilde{y}}_t\right) + \psi_3\Delta\widetilde{e}_t\right] + \varepsilon^R_t. \tag{4}$$

We assume that the policy coefficients $\psi_1, \psi_2, \psi_3 \geq 0$. In order to match the persistence in nominal interest rates, we include a smoothing term in the rule with $0 < \rho < 1$. ε^R_t is an exogenous policy shock and can be interpreted as the unsystematic component of monetary policy. One issue that we are interested in is whether monetary authorities include exchange rate terms in their reaction functions. We test this hypothesis by estimating the model separately under the restrictions $\psi_3 > 0$ and $\psi_3 = 0$. We reject one model specification in favor of the other by evaluating the posterior odds ratio.

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

$$\Delta\widetilde{q}_t = \rho_q\Delta\widetilde{q}_{t-1} + \epsilon_{q,t}. \tag{5}$$

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. Proper classification is therefore that of a semi-small open economy. When we attempted to estimate the model with endogenous terms of trade, we encountered several problems which convinced us to implement this alternative version. We will return to this issue later.

Equations (1)-(9) form a linear rational expectations model in the variables $\left[\widetilde{y}_t, \widetilde{\pi}_t, \widetilde{R}_t, \Delta\widetilde{e}_t, \Delta\widetilde{q}_t\right]$. We assume that $\widetilde{y}_t^*$ and $\widetilde{\pi}_t^*$ evolve according to univariate AR(1) processes with autoregressive coefficients ρ_{y^*} and ρ_{π^*}, respectively. The innovations of the AR(1) processes are denoted by $\epsilon_{y^*,t}$ and $\epsilon_{\pi^*,t}$. The model is solved using the method described in Sims (2002).

3 Estimation Strategy and Empirical Implementation

3.1 Econometric Methodology

Let y_t be the 5×1 vector of observables and $Y^T = \{y_1, \ldots, y_T\}$. The parameters of the structural model are collected in the 17×1 vector θ. The linear rational expectations model provides a state-space representation for y_t. Under the assumption that all the structural shocks are normally distributed and uncorrelated over time we obtain a likelihood function $L(\theta|Y^T)$ that can be evaluated using the Kalman filter.

We adopt a Bayesian approach and place a prior distribution with density $p(\theta)$ on the structural parameters. The data Y^T are used to update the prior through the likelihood function. According to Bayes Theorem the posterior distribution of θ is of the form

$$p(\theta|Y^T) = \frac{L(\theta|Y^T)p(\theta)}{\int L(\theta|Y^T)p(\theta)d\theta}. \tag{6}$$

Draws from this posterior can be generated through Bayesian simulation techniques described in detail in Schorfheide (2000). Posterior draws of impulse response functions and variance decompositions can be obtained by transforming the θ draws accordingly.

In the subsequent empirical analysis we are interest in examining the hypothesis that central banks do not react systematically to exchange rate movements. In the context of the reaction function specification this corresponds to $H_0 : \psi_3 = 0$. Let $\pi_{0,0}$ be the prior probability associated with this hypothesis. The posterior odds of H_0 versus $H_1 : \psi_3 > 0$ are given by[2]

$$\frac{\pi_{0,T}}{\pi_{1,T}} = \left(\frac{\pi_{0,0}}{\pi_{1,0}}\right)\left(\frac{p(Y^T|H_0)}{p(Y^T|H_1)}\right). \tag{7}$$

The first factor is the prior odds ratio in favor of $\psi_3 = 0$. The second term is called the Bayes factor and summarizes the sample evidence in favor of H_0. The term $p(Y^T|H_i)$ is called Bayesian data density and defined as

$$p(Y^T|H_i) = \int L(\theta|Y^T, H_i)p(\theta|H_i)d\theta. \tag{8}$$

The logarithm of the marginal data density can be interpreted as maximized log-likelihood function penalized for model dimensionality, see, for instance, Schwarz (1978). We use a numerical technique known as modified harmonic mean estimation to approximate (8).

3.2 Data Description and Choice of the Prior

The SOE model is fitted to data on output growth, inflation, nominal interest rates, exchange rate changes, and terms of trade changes. We consider data from the United Kingdom, Canada, Australia, and New Zealand. All data are seasonally adjusted and at quarterly frequencies for the period 1983:1 to 2002:3 for the UK and Canada, 1983:1 to 2001:4 for Australia and 1987:1 to 2001:4 for New Zealand. The series were obtained from the DRI International database. The output series is real GDP in per-capita terms, inflation is computed using the CPI. The nominal interest rate is a short-term rate for each country. As nominal exchange rate variable we use a nominal trade-weighted exchange rate index, whereas the terms of trade are measured as the (log-) ratio of export and import price indices. Unlike the latent variables world output and inflation, we treat the terms of trade as an endogenous variable since data are readily available, both to central banks and econometricians, and since they are a sharper defined concept than the former.[3] We de-mean the data prior to estimation.

[2]According to Jeffreys (1961) the posterior odds may be interpreted as follows: $\pi_{0,T}/\pi_{1,T} > 1$ null hypothesis is supported; $1 > \pi_{0,T}/\pi_{1,T} > 10^{-1/2}$ evidence against H_0 but not worth more than a bare mention; $10^{-1/2} > \pi_{0,T}/\pi_{1,T} > 10^{-1}$ substantial evidence against H_0; $10^{-1} > \pi_{0,T}/\pi_{1,T} > 10^{-3/2}$ strong evidence against H_0; $10^{-3/2} > \pi_{0,T}/\pi_{1,T} > 10^{-2}$ very strong evidence against H_0; $10^{-2} > \pi_{0,T}/\pi_{1,T}$ decisive evidence against H_0.

[3]See Bergin (2003) for an alternative approach. He constructs world interest and inflation series as weighted averages of the G7.

We choose priors for the structural parameters to be estimated based on several considerations. Table 1 provides information about the distributional form, means and 90% confidence intervals. 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 at the boundary of the determinacy region.[4] Our initial prior assigns approximately 5% probability to indeterminacy. Mean and confidence intervals are then calculated for the truncated version of the prior.

We use fairly wide confidence intervals 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. The confidence interval for the smoothing parameter ranges from 0.17 to 0.83. The determinacy region is likely to be more complicated than in the closed-economy version. In particular, low values of ψ_1 ($\psi_1 < 1$) lead to determinacy *ceteris paribus* if ψ_3 is sufficiently large. The confidence interval for ψ_1 reported in the table therefore extends to 0.8.

The model is parameterized in terms of the steady state home real interest rate r, rather than the discount factor β. r is annualized so that the conversion is $\beta = exp[-r/400]$. Our prior confidence interval for the real rate ranges from 0.9 to 4 percent. 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, Gali and Gertler, 1999, and Sbordone, 2002). Its mean is set at 0.5, but we allow it to vary widely in the unit interval.[5] The prior for the import share α is centered at 0.3 with a plausible degree of variation for the countries in question.

We choose identical priors for the parameters of each model economy with one exception: we allow for country specific variation in the exogenous shock processes to capture possibly different macroeconomic histories. To specify the priors we used a pre-sample from 1973:1 to 1982:4.[6] We

[4]Lubik and Schorfheide (2003) estimate the simple closed economy version of the present model allowing for the possibility of indeterminacy and sunspot driven business cycle fluctuations.

[5]Note that in the estimation we put a prior on the coefficient $\kappa^* = \frac{\nu-1}{\tau\varphi\pi^2}$ instead of the individual structural parameters ν, τ, φ.

[6]Table 1 only reports priors for the United Kingdom. The priors for the other countries are largely similar. Details are available from the authors.

fitted an AR(1) to terms of trade data to obtain priors for the Δq_t processes, and to output growth to get prior for the technology processes. Priors for the rest-of-world shocks π_t^* and y_t^* are obtained by estimating an AR(1) for US inflation and the ratio of US output to domestic output. The priors are by and large consistent with the estimates, although we increased their variance.

3.3 A Fully Structural (Non-) Alternative Specification

We pointed out earlier that our baseline model is not fully consistent with the underlying structural model. Since domestic firms have market power over their exports their relative price can no longer be treated as exogenous. 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:

$$[\tau + \alpha(2-\alpha)(1-\tau)]\,\Delta\widetilde{q}_t = \Delta\widetilde{y}_t^* - \Delta\widetilde{y}_t. \tag{9}$$

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

Empirical implementation of the fully structural model turned out to be problematic, however. A first hurdle is that endogenizing the terms of trade results in the loss of an exogenous shock in the equation system, namely the innovation to the terms of trade in Eq. (5). The covariance matrix of the system and, consequently, of the posterior distribution would then be singular. A convenient remedy is to introduce 'measurement error' in one of the structural relationships. Although not as appealing as identifying an additional disturbance with economic content (such as terms of trade variations), this allows the researcher to capture unobservable, yet plausible deviations from the strict, benchmark modeling assumptions. Del Negro (2003), for instance, adds measurement error to his counterpart of Eq. (3). The intuitively appealing rationale is that this captures non-systematic deviations from purchasing power parity. However, this modification creates identification problems in our set-up. The decomposition of CPI inflation already includes an exogenously determined shock process, world inflation $\widetilde{\pi}_t^*$, which does not appear in any of the other equations. Without further cross-equation restrictions we would not be able to identify the two shocks separately.

We therefore added measurement error to Eq. (9). Since world output determines the domestic natural rate, the model provides in principle enough independent restrictions to identify the measurement error as deviation from endogenous terms of trade dynamics. The latter are derived under the assumption of international asset market completeness. As in Cole and Obstfeld (1991) adjustment in the terms of trade serves as an endogenous risk-sharing mechanism. Alternatively, and at the price of more complexity, we could have assumed incomplete financial markets such that current account dynamics and foreign asset accumulation play a role. Measurement error might therefore capture deviations from perfect risk-sharing or, as in Bergin (2002), deviations from uncovered interest parity.

Estimation of the thus modified model proved to be difficult. Most specifications did not converge, and those that did resulted in implausible parameter estimates or did not attain local maxima. In particular, the Phillips curve parameter κ could not be estimated, while the preference parameter α took on values in excess of 0.5. The apparent reason is that Eq. (9) 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 they are straightforward to interpret.

3.4 GMM Estimation of the Monetary Policy Rule

We first estimate the open economy monetary policy rule (4) by GMM to establish a point of comparison with our structural estimates. GMM estimation has been advocated by Clarida, Galí, and Gertler (1998, 2000) since it potentially incorporates information from an underlying economic structure. On the other hand, OLS estimation of monetary policy rules, for instance as in Orphanides (2001) or Ball and Tchaidze (2002), has to rely on implausible identification assumptions to avoid any endogeneity bias.[7]

GMM estimates are reported in Table 2. The sample periods for the individual countries are as reported above. Our list of instruments includes lagged values of the variables in the regression equation. We performed estimation by using two measures of the output gap, as deviation from

[7] Lubik and Schorfheide (2003) present a simple example where OLS estimates of the Taylor-rule are inconsistent and exhibit a severe downward bias to the extent that the estimate of the inflation coefficient converges to zero even if the true value is above unity.

a linear trend and as HP-detrended output. The results were roughly similar, so that we report estimates only under the latter specification. The estimates for the UK are in line with the conventional wisdom on Taylor rules. The inflation coefficient is almost 1.5, and there is a moderate degree of output targeting and interest rate smoothing, although the former is not statistically significant. The coefficient on the nominal depreciation rate is small with a negative sign which implies that the Bank of England loosens policy when the Pound depreciates.

The GMM estimates are more dramatic for Canada and Australia. The output gap coefficient in Australia is exceedingly large, but insignificant. Proxying the output gap by deviations from a linear trend reduces its coefficient substantially, but it is still not statistically different from zero. Initial estimates of the policy rule for Canada revealed a negative smoothing coefficient, which we restricted to zero in subsequent estimation. Table 2 presents the estimates under this parameter restriction. Strikingly, the inflation and exchange rate coefficients are very large, while the output coefficient is almost zero. The same picture emerges when the linearly detrended output gap measure is used.

We believe that our estimation procedure carries several distinct advantages over single equation estimation procedures even if they utilize implicit information from an underlying structural model (as in Clarida, Galí, and Gertler, 2000). Viewed from a classical perspective, the likelihood-based estimation procedure uses the optimal set of instruments if the model is correctly specified. GMM estimation is less accurate, since it uses suboptimal instruments and a weighting matrix for the moment conditions has to be constructed from a fairly short sample of observations.[8] Even if the model is to some extent misspecified, we can expect that the instruments that are implicitly used in the likelihood procedure are approximately valid and more effective than the instruments typically used in GMM estimations.

Moreover, we obtain estimates for the other model parameters and are able to generate impulse response functions and variance decompositions which is not possible in a univariate framework. Through the prior distributions we can also formally incorporate information in the analysis that is not contained in the sample from 1983:1 to 2002:3, such as information about the behavior of world

[8]The additional advantage of a Bayesian approach is that its applicability is not hindered by sample size. While the estimates will be dominated to some extent by the prior in small samples, the posterior summarizes the information contained in the data in a consistent fashion, whose validity does not depend on convergence arguments in large samples.

inflation and output. Monetary policy is usually specified as reacting to the output gap defined as the deviation of actual from potential output. Since measures of the latter are typically difficult to come by, it is incorporated as a linear or HP-filtered trend. Instead, our approach treats potential output as a latent variable.

4 Bayesian Estimation Results

The Bayesian estimates of the structural parameters can be found in Table 3. In addition to 90% posterior probability intervals we report posterior means as point estimates. While the estimates for the four countries are reasonable and broadly similar, a closer inspection reveals some notable differences. All countries pursue strict anti-inflationary policies with inflation coefficients ψ_1 ranging from 1.84 (UK) to 2.49 (New Zealand). Similarly, significant emphasis is put on output gap targeting, with New Zealand being the most aggressive ($\hat{\psi}_2^{NZ} = 0.26$). There is also a high degree of interest-smoothing with an average estimate of 0.70. These estimates are broadly in line with the commonly used values in the monetary policy literature as well as empirical studies of closed economies (Clarida, Galí, and Gertler, 2000, Lubik and Schorfheide, 2003). We also find significant estimates for the exchange rate coefficient in the three economies policy rule: $\hat{\psi}_3^{UK} = 0.07$, $\hat{\psi}_3^{CAN} = 0.09$, $\hat{\psi}_3^{AUS} = 0.08$, $\hat{\psi}_3^{NZ} = 0.24$. The respective central banks thus respond to a nominal depreciation with a monetary tightening of between 7% and 24% of the initial percentage-point exchange rate movement.

The combined estimates of the inflation and exchange rate coefficients suggest that the monetary authorities pursue strong anti-inflationary policies.[9] Differences in the estimates may reflect central bank preferences, but also underlying structural characteristics of the economies. Exchange rate movements affect domestic aggregates via import price changes that have real effects under price stickiness. Ceteris paribus, exchange rates have stronger domestic effects the more open the economy is and the more distorted, in which case the central bank may optimally chose a more anti-inflationary policy.[10]

[9] However, our estimates tend to be overall lower than those found optimal in the *theoretical* literature; see our discussion above and Taylor (2001). In a framework with an optimizing authority that chooses response coefficients in a linear rule, Kollmann (2003) suggests that the inflation coefficient should be in the two digits.

[10] An additional consideration is whether firms engage in producer- or consumer-currency pricing. Bergin (2003) finds strong support for a model in which, contrary to ours, prices are sticky in the currency of the buyer.

We can gain some insight into this question by comparing estimates of the import share α and the slope parameter of the Phillips-curve κ across countries. For each country κ is estimated to be around 0.5 which is consistent with values found in closed economy studies. There are differences, however, with respect to the import share parameter. Estimates are 0.07 for the U.K., 0.16 for Canada, 0.21 for Australia and New Zealand. At the same time Australia's growth rate of the terms of trade exhibits the highest degree of autocorrelation and the highest innovation variance among the four. This suggests that the Reserve Bank of Australia (RBA) pursues a strong policy of 'leaning against the wind' since its economy is relatively open and buffeted by more persistent and volatile international relative price shocks. A similar argument can be made for New Zealand, which compared to the other countries has a narrower export base, and its economy is more dependent on agriculture and natural resources. It can therefore be expected that the Reserve Bank of New Zealand (RBNZ) will be more inclined towards exchange rate targeting. On the other hand, the U.K.'s terms of trade are the least persistent and volatile. It is also perhaps surprising that the economy of the U.K., the proverbial small open economy in many empirical studies, is relatively closed to the rest of the world. An explanation might be that the other countries are large resource exporters and therefore subject to international commodity price fluctuations. Since identification of the import share parameter is achieved from the relationship between the terms of trade and domestic output and inflation, the relatively low volatility of the former may affect its value.

In order to gauge the importance of the individual shocks we compute variance decompositions. The results are reported in Table 4. In all countries domestic output is largely driven by technology and to a lesser degree by (latent) world output shocks. The contribution of monetary policy innovations is slightly below 10% which is in line with evidence from VAR studies. Interestingly, world output also contributes significantly to inflation and interest rate volatility. A possible explanation is the reliance of the model on complete international asset markets. Introducing current account dynamics would be an interesting extensions of the model.

Exchange rate movements on the other hand are largely determined by foreign inflation, which proxies for foreign monetary policy shocks, and to a smaller degree by the terms of trade. In particular, their contribution to exchange rate volatility is 19% and 32% for Australia and Canada, respectively; and 8% and 5% for the U.K. and New Zealand. Although terms of trade movements do not play a substantial role in domestic business cycles, this finding lends support to the conclusion that central banks respond to exchange rates to smooth the impact of international relative price movements. On a final note, it is worth pointing out that the marginal role of the terms of trade stands in contrast to the international real business cycle literature. For instance, in a calibration analysis with 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%.

Estimated impulse response functions can be found in Figure 1. Since the dynamic behavior of the four countries is qualitatively very similar we only report the results from Canada. The model contains only weak endogenous propagation mechanisms 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 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 a depreciation. This result stems from the fact that world output shocks lower domestic potential output under the estimate of τ =0.4 (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 domestic currency, but raise inflation since the central bank reacts to movements in the exchange rate and subsequently relaxes policy.

As mentioned above, the estimated dynamic behavior of the four economies is almost identical. Only New Zealand's response to foreign output shocks stands out. Percentage-wise, the contraction in output is more than twice the size of the other countries. Since foreign shocks enter directly via the output gap and feed back through exchange rate targeting, this response crucially depends on two structural characteristics: the economy's openness and its degree of exchange rate targeting. The larger α and the more the central bank reacts to exchange rate movements, the more volatile domestic output becomes. Openness exposes an economy to foreign disturbances. In our framework, however, responding to the exchange rate exacerbates this since it introduces an additional transmission channel. Since this is likely not the preferred outcome for policymakers, we consequently ask the question to what extent the four central banks in our sample *do* in fact respond to exchange rates movements.

5 Do Central Banks Respond to Exchange Rates?

We reestimate the model for the four countries under the restriction that the exchange rate does not play an independent role in the monetary policy rule, that is, we set $\psi_3 = 0$. The results can be found in Table 5. The parameter estimates for the UK do not change dramatically. Inflation and output coefficients decrease and increase, respectively, while α and κ rise, but there is substantial overlap of the confidence intervals in these cases. A similar pattern can also be observed for Canada and Australia, although for the latter the differences are more pronounced. The estimation results for New Zealand go in the opposite direction. Under the restriction $\psi_3 = 0$, inflation targeting is estimated to be more aggressive with a response coefficients of 2.79 compared to 2.49 in the unrestricted model. Clearly, when monetary policy cannot react to exchange rate movements, these variations have to be picked up by a stronger inflation response. This suggests a tight link between the exchange rate and the CPI and domestic output in New Zealand. Similarly, the output coefficient drops from 0.26 to 0.17, while α and κ also decline. Impulse response functions are similar (see Figure 2) to the unrestricted case, with the exception of 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 channel. In the absence of exchange rate targeting, a domestic appreciation absorbs the inflation shock.

We now formally test the hypothesis $H_0 : \psi_3 = 0$ against the alternative $H_1 : \psi_3 > 0$ by computing the posterior odds ratio. The results are reported in Table 6. The posterior odds of the null hypothesis of no exchange rate targeting is 767:1 for the U.K and more than 1500:1 for New Zealand and the UK which leads us to strongly reject the null. In the case of Canada, on the other hand, the posterior odds ratio is zero, thus favoring exchange rate targeting.[11] 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 an MCI in gauging nominal demand pressures.

Neither the Bank of England nor the Reserve Bank of Australia announce an exchange rate target nor do they explicitly compute an MCI. In fact, the RBA has been pursuing a policy of strict inflation targeting (Macfarlane, 2000), which is in line with our empirical analysis. Official statements suggest that the exchange rate is nevertheless an important variable in informing the central bank's thinking about the structure and state of the economy (Stevens, 1999). International exposure changes the responsive of the economy to interest rate changes, while exchange rate movements contain information on inflation. Directly responding to exchange rate fluctuations may not be the optimal policy since it creates an additional transmission mechanism for foreign shocks.[12]

The result for the U.K. may be puzzling in light of the fact that over the period 1989-1992 it was part of the ERM. In order to study the relationship between German and British monetary policy during the ERM, we could include the Pound-Deutschmark exchange rate in the interest-rate equation. A high coefficient would reveal adherence to a fixed-exchange rate regime. In a study using GMM to estimate policy rules, Clarida, Galí, and Gertler (1998) find, however, that this coefficient is comparatively small. On the other hand, a short-term German interest rate is highly significant in the U.K.'s policy rule even before official membership in the ERM. We could properly analyze this policy episode in a regime-switching model. However, implementation in our structural model is not a trivial matter and beyond the scope of the present paper. The results for the U.K. may therefore be qualified as averages over two different policy regimes, and may indeed pick up elements of the German Bundesbank's behavior.[13]

We did, however, estimate the model for the sample period after the collapse of the ERM

[11] This is in contrast to Dib's (2003) finding that the exchange rate is not statistically significant in his specification of the Canadian rule, although he does not report estimates nor formally tests the two specifications.

[12] Dennis (2003) points out that (real) exchange rate targeting is only optimal in models of the type studied if policymakers target CPI-inflation.

[13] Even the 'hard' ERM was not a perfectly fixed regime. Exchange rates were allowed to fluctuate between narrow bands. To what extent this allowed individual central banks to pursue independent policies during normal economic conditions is a matter open to debate. When economic conditions were no longer normal due to fiscal and monetary pressures arising from German unification, the ERM all but crumbled.

from 1992:4 on. Interestingly, policy parameter estimates increase substantially for both model specifications when compared to the full sample. For instance, in the unrestricted (restricted) model $\hat{\psi}_1^{UK}$ increases to 2.64 (2.30) from 1.84 (1.74). Post-ERM the Bank of England apparently shifted to a much a more anti-inflationary stance which is borne out by other empirical studies (see Batini and Turnbull, 2002).

Interestingly, we can strongly reject an exchange rate response for New Zealand despite the fairly large coefficient in the unrestricted model. While a single-equation study might conclude from this that the RBNZ pursues an exchange rate policy, our structural general equilibrium analysis reveals that the restricted model offers a better description of the data. It bears notice that New Zealand is the only case where the restricted model implies a higher inflation coefficient. This may be explained by relative openness and a high degree of pass-through. Exchange rate movements thereby quickly affect inflation which the central bank counters aggressively. Although the RBNZ has never formally endorsed an exchange rate target, it has famously shifted emphasis throughout 1997 towards an MCI that includes the exchange rate. Since this policy was formally de-emphasized in 1999, this time span is presumably too short to change the results from the posterior odds test. To investigate this issue further, we will, however, re-estimate the model under an MCI-rule in the next section.

In summary, estimation of our simple structural model under different specifications of the monetary policy rule provides evidence that the central banks of Australia, New Zealand, and the U.K. do not respond to exchange rates while the Bank of Canada (BoC) does. This is consistent with official statements from the BoC which emphasize exchange rate movements in determining aggregate demand conditions. Where the BoC went beyond other central bank is that it formulated policy with explicit reference to the exchange rate as a measure of aggregate monetary conditions for an extended period of time.

6 Robustness Analysis

Our empirical results so far have shown to what extent, and if at all, the central banks in the sample countries pursued a policy of systematically responding to exchange rate movements. Naturally, the analysis is conditional on the modelling framework. In this Section we therefore investigate whether our conclusions are robust to changes in the model's structure. We do not modify the equations derived from the underlying structural model, but consider alternative specifications for the ad-hoc policy rule as well as different data and subsamples. While it is possible that our conclusions can be altered in more elaborate models with richer exchange rate and current account dynamics, we regard our parsimonious model as sufficient for the questions we seek to answer. We begin by considering alternative specifications for the policy rule such as expected inflation targeting. We then look at a type of policy rule that utilizes a monetary conditions index in determining the monetary stance. Finally, we estimate the model allowing for real exchange rate targeting.

6.1 Expected Inflation Targeting

We now estimate the model for different specifications of the monetary policy reaction function. Both policymakers and academic researchers emphasize the forward-looking nature of the monetary policy process. 'Leaning against the wind' allows the monetary authority to strike preemptively against future inflationary dangers. We therefore re-estimate our baseline model with an interest-rate rule that targets expected inflation:

$$\widetilde{R}_t = \rho\widetilde{R}_{t-1} + (1-\rho)\left[\psi_1 E_t\widetilde{\pi}_{t+1} + \psi_2\left(\widetilde{y}_t - \widetilde{\overline{y}}_t\right) + \psi_3\Delta\widetilde{e}_t\right] + \varepsilon_t^R. \tag{10}$$

The model is then estimated under the restriction ψ_3=0. We perform posterior odds tests for the hypothesis of no exchange rate targeting versus its alternative, and for the hypothesis of current inflation targeting against expected inflation targeting.

Regarding the estimates of the policy parameters we find that the inflation coefficients $\widehat{\psi}_1$are considerably larger than under the current-inflation rule, ranging from 2.51 (Australia) to 3.06 (U.K.). This implies a much more aggressive policy stance, in particular for the Bank of England whose inflation coefficient increased by a factor of 1.7. At the same time, there is less emphasis on interest-rate smoothing with the average coefficient dropping to 0.65. Setting the exchange rate coefficient equal to zero reduces the inflation coefficients, but they are still larger than the corresponding ones under the current-looking rule. For all four countries the slope coefficient $\widehat{\kappa}$ in the Phillips-curve (2) is estimated to be around 0.3, whereas under current inflation targeting $\widehat{\kappa} = 0.5$. Central banks can achieve a higher degree of inflation smoothing when responding to its expected level. Since this results in more output volatility, our estimation procedure fits the model to the data by returning a lower value of the Phillips-curve coefficient.

We find that for all four countries expected inflation rate targeting results in a worse model fit, as measured by the marginal data densities, than the previously accepted specification. That is, all central banks in question are more likely to respond to current inflation. The ranking of policy rules with respect to exchange rate targeting remains unaffected. The Bank of Canada's rule is the only one for which this is still the preferred specification.

6.2 Reinterpreting Interest Rate Rules as MCIs

Much of the recent monetary policy literature focuses on the use of interest rate-based rules, in which the 'instrument', typically a short-term money rate over which the monetary authority has some control, is adjusted in response to deviations of selected macroeconomic variables from their 'targets'. Although no central bank has acknowledged explicitly the use of interest-rate rules, there is ample evidence from a wide range of empirical studies that the joint behavior of interest rates and target variables approximates these rules quite well. Instead of implementing, say, a Taylor-rule the Bank of Canada and the Reserve Bank of New Zealand have prominently and explicitly experimented with a Monetary Conditions Index.

An MCI is a weighted average of interest rates and exchange rates. Its usage in the formulation and implementation of monetary policy has been advocated, for instance, by Ball (1999). The basic idea is that interest rate and exchange rate movements affect domestic demand via potentially different transmission mechanisms. While monetary policy changes the level of aggregate demand through the real interest rate channel, exchange rates affect international relative prices and the composition of demand. Moreover, changes in the exchange can have a direct effect on inflation through import prices. The 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 the central bank earlier warning of future inflationary conditions.

The Bank of Canada uses the MCI "[...] *as the operational target of policy, in much the same way as interest rates were used in the past*" (Freedman, 1995, p. 53). Following the lead of the BoC the Reserve Bank of New Zealand implemented an MCI in June 1997, but abandoned the practice of stating a 'desired level' of monetary conditions in early 1999 after a deep disinflation in the wake of the East Asian Crises. In an almost natural experiment Australia escaped the economic turmoil in the region relatively unscathed. Unlike New Zealand, which adopted an MCI-based rule shortly before the onset of the crisis, Australia's Reserve Bank outspokenly opposed its use (see Stevens, 1998). The Bank of England does not use an MCI either, although various financial and international institutions as well as central bank researchers have suggested alternative MCIs, for instance Batini and Turnbull (2002).

MCI's are constructed in practice from estimated reduced-form equations, whereby the weighting coefficients in the index represent the effect that interest and exchange rates have on aggregate demand. For instance, the Bank of Canada's MCI puts a weight of 1/3 on the depreciation rate, which implies that a 1% increase in the nominal rate has the same effect as a 3% appreciation of the Canadian Dollar. Similarly, the exchange rate weight in the RBNZ is roughly 1/2. Our estimates from the baseline model, however, show values for ψ_3 that are close to zero. Since the earlier analysis cannot deliver conclusions to what extent an MCI may have served as a monetary instrument, we therefore re-estimate our model with an explicit MCI-based rule.

We can construct an MCI in our model in various ways. Essentially any rearrangement of the monetary policy rule (4) with the interest rate and the exchange rate on the left-hand side of the equation constitutes an MCI relative to some base period. For instance, rewrite the policy rule in the following way (assuming that $\rho = 0$):

$$MCI_t = \widetilde{R}_t - \psi_3 \Delta \widetilde{e}_t = \psi_1 \widetilde{\pi}_t + \psi_2 \left(\widetilde{y}_t - \widetilde{\widetilde{y}}_t \right) + \varepsilon_t^{MCI}. \tag{11}$$

This MCI-rule is similar to the specification found to be optimal by Ball (1999), and it coincides with the simple Taylor-rule with respect to its targeting variables. The crucial difference, however, is that policymakers now have a choice of adjusting instruments. The policy is implemented by adjusting actual monetary conditions, measured by MCI_t to match desired conditions, the weighted average of inflation and the output gap. Here, the distinction between these rules is purely semantic since rearranging the variables in the policy rule does not affect the reduced-form representation of the estimated model.[14]

The situation is different in a specification that allows for instrument smoothing since this introduces additional lag terms. In order to differentiate the assumed MCI-based rule from earlier interest rate rules, we estimate the following specification:

$$MCI_t = \rho MCI_{t-1} + (1 - \rho) \left[\psi_1 \widetilde{\pi}_t + \psi_2 \left(\widetilde{y}_t - \widetilde{\widetilde{y}}_t \right) \right] + \varepsilon_t^{MCI}, \tag{12}$$

[14] Ball (2000) gives an argument why this distinction is not entirely semantic in practice.

where:

$$MCI_t = (1 - \psi_3)\,\widetilde{R}_t - \psi_3 \Delta \widetilde{e}_t, \tag{13}$$

preserving the notation from before. Normalizing on the interest rate coefficient allows us to rewrite the MCI as an interest rate rule with a stricter emphasis on exchange rate smoothing. Restricting the coefficients in the MCI to sum up to zero also allows comparison with related studies, e.g. Ball (1999).

The estimates for the modified model can be found in Table 7. We only report the policy parameters since the structural parameter estimates do not differ significantly from the benchmark specification, nor do the variance decompositions. Yet, there is also no marked change for the coefficients in the policy rule under an MCI-specification. Only Australia shows a sizeable decrease in $\widehat{\psi}_1$, but the confidence interval largely overlap. It is noticeable, however, that the weight on the exchange rate drops in all countries (and precipitously so in Canada). This is indicative of the notion that the central banks prefer to smooth interest rates rather than exchange rates. These findings also point towards the earlier conclusion that the exchange rate does not play an explicit role in setting monetary policy.

We can adjudicate this assertion formally by testing the MCI-rule against other interest-rate rules. The former can be soundly rejected against the previously accepted specification; that is, there is no evidence of exchange rate targeting in Australia, New Zealand and the U.K., whereas the data favor a positive exchange rate coefficient in the Bank of Canada's rule, but not an MCI-based rule. This may be surprising in light of the BoC's stated reliance on an MCI, but official statements tend to fall short of endorsing an MCI-rule.

6.3 Real Exchange Rate Targeting

Real exchange rate targeting is often advocated as a monetary policy for developing countries and natural resource exporters.[15] Since Australia, Canada, New Zealand and, to a smaller extent, the U.K. have large primary sectors, we are therefore interested to what extent monetary policy in these countries can be described as real exchange rate targeting. While we found no evidence of systematic nominal exchange rate responses, it may very well be that the central banks attempt to

[15] Calvo, Reinhart and Végh (1995) give an overview of the theoretical and empirical arguments, while Uribe (2003) argues that such rules can lead to macroeconomic instability.

smooth international relative price movements that could adversely affect primary exporters. We investigate this issue by estimating the model under policy rule (4) with $\Delta\widetilde{q}_t$ in place of the nominal exchange rate. An increase in the terms of terms implies a real appreciation so that we expect the central bank to loosen its monetary stance.

Overall, we find that the exchange rate coefficients in the policy rule are large and significant, ranging from -0.13 in the U.K to -0.24 in Australia.[16] Other parameter estimates are broadly similar to our benchmark specification which allows for a nominal exchange rate response. Inflation coefficients are smaller, while output coefficients larger. This reflects the transmission channel of terms of trade movements into domestic output, which the monetary authorities attempt to accommodate. Since inflation does not move one to one with the terms of trade, a stronger response to the latter is required. The posterior odds tests clearly reject real exchange rate targeting for Australia as well as for New Zealand. In Canada, the data favor a policy rule that responds to real exchange rates over one that includes nominal rates.[17] The results for the U.K are surprising. The posterior odds in favor of real exchange rate targeting over a conventional Taylor-rule are 765:1, although we were able to reject a nominal rate response. This result is puzzling in that the U.K. is not a large resource exporter. A possible explanation is that the Bank of England, before becoming formally independent in 1997, was concerned about international competitiveness, whereas nominal exchange rate movements were largely as driven by financial flows. We leave this issue open for further research.

7 Summary and Concluding Remarks

We briefly summarize our findings, highlight the shortcomings of our approach and point towards further research. We formally estimate a small-scale, structural general equilibrium model of an

[16]New Zealand's terms of trade coefficient enters with a positive sign, implying that the RBNZ tightens with a real appreciation. Since ceteris paribus a real appreciation increases prices of domestically produced goods, the RBNZ may be more concerned with inflationary than with competitive pressures. However, this specification is rejected against the previously preferred, i.e. no exchange response.

[17]When we estimated the model with bilateral U.S. Dollar exchange rates, nominal exchange rate targeting was the preferred policy over all alternatives. Since the U.S. is by far the largest import and export market for Canada, it appears that the terms of trade variable picks up this variation. this is reminiscent of the argument in Schmitt-Grohé (1998).

open economy using Bayesian methods. Our main focus is the conduct of monetary policy in selected economies as described by nominal interest rate rules. The main finding of this paper is that the central banks of Australia, New Zealand and the U.K. do not respond to exchange rate movements, whereas the Bank of Canada does. This finding is robust to various alternative specifications for the policy rule with the possible exception of real exchange rate targeting in the U.K. This is not to say that the exchange rate is not part of the decision-making process; openness changes the structure of the economy and its reaction to monetary policy. Moreover the exchange rate may carry important information about future demand conditions. However, we do not find evidence that central banks alter their interest rate instrument directly in response to a depreciation.

Naturally, our results have to be qualified with respect to the specific structural model employed. In particular, we highlighted the problems we encountered with a fully structural model that turned out to be too restrictive for the data. While the New Keynesian framework has been reasonably successful in describing the behavior of aggregate output and inflation its forecasting performance and ability to match the data pales in comparison to vector autoregressions. A richer model environment will be a step in the right direction.

We would find it useful to extend the paper in several directions. We maintain throughout the assumption of perfect pass-through of nominal exchange rate changes into domestic import prices. However, there is overwhelming empirical evidence across countries of a far lower degree of pass through.[18] The theoretical model could be modified along the lines in Monacelli (2003), where producers set their prices in terms of consumers' currency. This assumption alters the central bank's policy trade-off and introduces a potentially much larger role for exchange rate stabilization. In our framework, however, the policy problems in closed and open economies are isomorphic. If the central bank solved an optimal policy problem it need not pay attention to exchange rates in a pricing–to-market framework. In this case, the exchange rate coefficient in the properly defined policy rule would be zero, whereas under imperfect pass-through it most likely would not be. Moreover, this modification would allow us to estimate the degree of pass-through in a theory-consistent general equilibrium framework.

[18] Estimating a two-country model for the U.S. and the other G-7 countries, Bergin (2002) finds that a fairly low degree of pass-through results in a more adequate model fit.

Secondly, the terms of trade play an almost negligible role in aggregate fluctuations. This finding 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.

Finally, it is well known that the use of interest-rate rules can lead to equilibrium indeterminacy such that aggregate dynamics are influenced by non-fundamental shocks. We rule out this possibility by restricting our estimation to the region in the parameter space for which a unique equilibrium exists. However, central banks may (inadvertently) implement indeterminate rules, in which case our empirical model is misspecified. Moreover, extrinsic belief shocks, for instance, may help explain inflation dynamics. Lubik and Schorfheide (2003) develop a methodology to estimate models of this type.

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Table 1: Prior Distribution (United Kingdom)

Name	Range	Density	Mean	90% Interval
ψ_1	$I\!R^+$	Gamma	1.54	[0.80, 2.24]
ψ_2	$I\!R^+$	Gamma	0.25	[0.05, 0.43]
ψ_3	$I\!R^+$	Gamma	0.25	[0.05, 0.43]
ρ_R	$[0, 1)$	Beta	0.50	[0.17, 0.83]
α	$[0, 1)$	Beta	0.30	[0.14, 0.46]
r	$I\!R^+$	Gamma	2.51	[0.92, 4.03]
κ	$I\!R^+$	Gamma	0.50	[0.12, 0.87]
τ	$I\!R^+$	Gamma	0.50	[0.19, 0.80]
ρ_q	$[0, 1)$	Beta	0.40	[0.06, 0.71]
ρ_A	$[0, 1)$	Beta	0.20	[0.04, 0.35]
ρ_{y^*}	$[0, 1)$	Beta	0.90	[0.83, 0.98]
ρ_{π^*}	$[0, 1)$	Beta	0.70	[0.48, 0.95]
σ_R	$I\!R^+$	InvGamma	1.25	[0.55, 1.99]
σ_q	$I\!R^+$	InvGamma	2.50	[1.06, 3.94]
σ_A	$I\!R^+$	InvGamma	1.89	[0.80, 3.00]
σ_{y^*}	$I\!R^+$	InvGamma	1.89	[0.80, 3.00]
σ_{π^*}	$I\!R^+$	InvGamma	1.89	[0.80, 3.00]

Notes: The Inverse Gamma priors are of the form $p(\sigma|\nu, s) \propto \sigma^{-\nu-1}e^{-\nu s^2/2\sigma^2}$, where $\nu = 4$ and s equals 1, 2, 1.5, 1.5, and 1.5, respectively. The prior is truncated at the boundary of the determinacy region.

Table 2: GMM Estimation Results

	ψ_1	ψ_2	ψ_3	ρ
Australia	1.33	2.97	0.29	0.85
	(0.57)	(3.67)	(0.61)	(0.16)
Canada	2.38	0.09	1.57	0
	(0.51)	(0.47)	(1.03)	0
New Zealand	1.61	0.43	-0.42	0.79
	(0.24)	(0.61)	(0.31)	(0.04)
UK	1.49	0.14	-0.10	0.54
	(0.23)	(0.40)	(0.37)	(0.23)

Notes: The table reports GMM point estimates and standard errors in parentheses for the parameters of the policy rule (4). Interest rates, inflation, and exchange rate changes are demeaned. We use HP-detrended output as a measure of the output gap. Instruments are $\widetilde{R}_{t-1}$, $\widetilde{\pi}_{t-1}$, $\Delta\widetilde{e}_{t-1}$, and the first lag of the output gap.

Table 3: Parameter Estimation Results, $\psi_3 \geq 0$

	United Kingdom		Canada		Australia		New Zealand	
	Mean	90% Interval	Mean	90% Interval	Mean	90% Interval	Mean	90% Interval
ψ_1	1.84	[1.47, 2.17]	2.24	[1.80, 2.62]	2.10	[1.60, 2.56]	2.49	[1.78, 3.32]
ψ_2	0.15	[0.03, 0.27]	0.21	[0.08, 0.39]	0.16	[0.06, 0.25]	0.29	[0.05, 0.47]
ψ_3	0.07	[0.03, 0.11]	0.09	[0.03, 0.15]	0.08	[0.04, 0.11]	0.24	[0.11, 0.34]
ρ_R	0.65	[0.59, 0.72]	0.73	[0.69, 0.79]	0.73	[0.67, 0.79]	0.73	[0.63, 0.81]
α	0.07	[0.04, 0.10]	0.16	[0.10, 0.20]	0.21	[0.15, 0.27]	0.21	[0.07, 0.34]
r	2.17	[1.22, 3.30]	2.41	[1.19, 4.00]	2.70	[1.24, 4.47]	2.40	[0.93, 3.76]
κ	0.48	[0.33, 0.69]	0.55	[0.39, 0.73]	0.59	[0.36, 0.78]	0.41	[0.27, 0.60]
τ	0.36	[0.28, 0.43]	0.50	[0.38, 0.57]	0.44	[0.36, 0.49]	0.46	[0.34, 0.60]
ρ_q	0.14	[0.06, 0.22]	0.29	[0.17, 0.42]	0.29	[0.18, 0.41]	0.26	[0.09, 0.42]
ρ_A	0.41	[0.36, 0.45]	0.48	[0.44, 0.54]	0.53	[0.48, 0.59]	0.46	[0.39, 0.53]
ρ_{y^*}	0.94	[0.91, 0.96]	0.88	[0.85, 0.91]	0.91	[0.88, 0.93]	0.95	[0.93, 0.97]
ρ_{π^*}	0.34	[0.24, 0.46]	0.36	[0.23, 0.52]	0.30	[0.17, 0.40]	0.28	[0.15, 0.36]
σ_R	0.41	[0.33, 0.47]	0.41	[0.35, 0.46]	0.50	[0.43, 0.58]	0.46	[0.38, 0.54]
σ_g	1.37	[1.21, 1.56]	1.30	[1.15, 1.44]	2.15	[1.96, 2.35]	2.22	[1.95, 2.54]
σ_A	0.76	[0.64, 0.89]	0.76	[0.67, 0.85]	0.91	[0.76, 1.04]	1.02	[0.85, 1.19]
σ_{y^*}	1.25	[0.78, 1.56]	1.29	[0.81, 1.78]	1.20	[0.82, 1.59]	2.20	[1.12, 2.98]
σ_{π^*}	3.92	[3.60, 4.27]	2.45	[2.20, 2.76]	5.31	[4.79, 5.89]	3.59	[3.17, 4.01]

Notes: The table reports posterior means and 90 percent probability intervals (in brackets). The posterior summary statistics are calculated from the output of the posterior simulator.

Table 4: VARIANCE DECOMPOSITION (CANADA)

	Output	Inflation	Interest Rate	Exchange Rate
Policy	0.11	0.34	0.08	0.02
	[0.07, 0.14]	[0.27, 0.45]	[0.05, 0.11]	[0.01, 0.02]
Terms of Trade	0.01	0.03	0.03	0.18
	[0.00, 0.02]	[0.00, 0.07]	[0.01, 0.05]	[0.14, 0.21]
Technology	0.60	0.23	0.22	0.01
	[0.52, 0.68]	[0.18, 0.29]	[0.16, 0.28]	[0.01, 0.02]
World Output	0.28	0.37	0.68	0.02
	[0.21, 0.35]	[0.28, 0.49]	[0.60, 0.76]	[0.01, 0.03]
World Inflation	0.00	0.02	0.00	0.77
	[0.00, 0.01]	[0.02, 0.04]	[0.00, 0.00]	[0.73, 0.81]

Notes: The table reports posterior means and 90 percent probability intervals (in brackets). The posterior summary statistics are calculated from the output of the posterior simulator.

Table 5: PARAMETER ESTIMATION RESULTS, $\psi_3 = 0$

	United Kingdom		Canada		Australia		New Zealand	
	Mean	90% Interval	Mean	90% Interval	Mean	90% Interval	Mean	90% Interval
ψ_1	1.74	[1.46, 2.05]	2.09	[1.64, 2.58]	1.83	[1.57, 2.17]	2.79	[2.05, 3.76]
ψ_2	0.16	[0.04, 0.32]	0.24	[0.13, 0.38]	0.21	[0.07, 0.31]	0.17	[0.07, 0.31]
ψ_3	0.00	[0.00, 0.00]	0.00	[0.00, 0.00]	0.00	[0.00, 0.00]	0.00	[0.00, 0.00]
ρ_R	0.60	[0.51, 0.66]	0.70	[0.62, 0.76]	0.73	[0.66, 0.78]	0.74	[0.66, 0.81]
α	0.08	[0.05, 0.11]	0.18	[0.10, 0.29]	0.22	[0.16, 0.29]	0.13	[0.08, 0.19]
r	2.49	[1.03, 4.12]	2.39	[1.11, 3.54]	2.04	[1.16, 3.18]	2.48	[0.80, 4.34]
κ	0.56	[0.35, 0.76]	0.70	[0.44, 1.16]	0.55	[0.41, 0.69]	0.46	[0.26, 0.69]
τ	0.35	[0.27, 0.44]	0.45	[0.37, 0.52]	0.51	[0.41, 0.64]	0.40	[0.32, 0.49]
ρ_q	0.11	[0.01, 0.21]	0.25	[0.16, 0.36]	0.32	[0.23, 0.47]	0.18	[0.07, 0.30]
ρ_A	0.41	[0.35, 0.47]	0.50	[0.45, 0.58]	0.54	[0.49, 0.59]	0.44	[0.38, 0.52]
ρ_{y^*}	0.95	[0.93, 0.97]	0.90	[0.84, 0.94]	0.93	[0.91, 0.96]	0.95	[0.93, 0.96]
ρ_{π^*}	0.35	[0.23, 0.47]	0.37	[0.21, 0.51]	0.37	[0.22, 0.49]	0.39	[0.27, 0.56]
σ_R	0.40	[0.34, 0.46]	0.43	[0.37, 0.49]	0.47	[0.42, 0.53]	0.46	[0.39, 0.51]
σ_g	1.39	[1.25, 1.54]	1.35	[1.15, 1.54]	2.20	[1.96, 2.44]	2.14	[1.84, 2.38]
σ_A	0.72	[0.58, 0.92]	0.71	[0.55, 0.86]	0.92	[0.76, 1.10]	0.99	[0.79, 1.20]
σ_{y^*}	1.25	[0.89, 1.76]	1.10	[0.72, 1.40]	1.42	[0.84, 2.00]	1.75	[0.84, 2.41]
σ_{π^*}	4.03	[3.41, 4.83]	2.37	[2.08, 2.61]	5.38	[4.77, 6.03]	3.73	[3.26, 4.47]

Notes: The table reports posterior means and 90 percent probability intervals (in brackets). The posterior summary statistics are calculated from the output of the posterior simulator.

Table 6: POSTERIOR ODDS TEST

	Marginal Data Densities		Posterior Odds
	H_0	H_1	
Australia	-867.37	-871.31	51.41
Canada	-710.33	-707.39	0.05
New Zealand	-621.15	-624.20	21.11
UK	-764.02	-765.50	4.39

Notes: The Table reports posterior odds tests of the hypothesis $H_0 : \psi_3 = 0$ against the alternative $H_1 : \psi_3 > 0$. The posterior probabilities are calculated based on the output of the Metropolis algorithm. Marginal data densities are approximated by Geweke's (1999) harmonic mean estimator.

Table 7: Policy Parameter Estimation Results

	United Kingdom		Canada		Australia		New Zealand	
	Mean	90% Interval	Mean	90% Interval	Mean	90% Interval	Mean	90% Interval
				Expected Inflation, $\psi_3 \geq 0$				
ψ_1	3.06	[2.50, 3.60]	2.95	[2.22, 3.91]	2.51	[2.02, 3.16]	3.03	[2.24, 3.99]
ψ_2	0.35	[0.19, 0.54]	0.26	[0.11, 0.40]	0.29	[0.06, 0.45]	0.23	[0.11, 0.40]
ψ_3	0.04	[0.02, 0.06]	0.23	[0.14, 0.30]	0.09	[0.04, 0.15]	0.06	[0.02, 0.09]
ρ_R	0.63	[0.56, 0.72]	0.65	[0.57, 0.73]	0.70	[0.63, 0.76]	0.63	[0.49, 0.76]
				Expected Inflation, $\psi_3 = 0$				
ψ_1	2.44	[1.89, 2.83]	2.82	[2.11, 3.46]	2.42	[1.80, 3.02]	3.06	[2.46, 3.71]
ψ_2	0.22	[0.12, 0.34]	0.28	[0.12, 0.41]	0.23	[0.11, 0.33]	0.19	[0.10, 0.28]
ψ_3	0.00	[0.00, 0.00]	0.00	[0.00, 0.00]	0.00	[0.00, 0.00]	0.00	[0.00, 0.00]
ρ_R	0.59	[0.48, 0.71]	0.69	[0.60, 0.78]	0.68	[0.61, 0.73]	0.66	[0.59, 0.72]
				MCI Rule				
ψ_1	1.80	[1.36, 2.17]	2.24	[1.71, 2.87]	1.93	[1.55, 2.29]	2.42	[2.08, 2.93]
ψ_2	0.16	[0.04, 0.25]	0.21	[0.10, 0.36]	0.18	[0.07, 0.26]	0.16	[0.06, 0.26]
ψ_3	0.02	[0.01, 0.02]	0.06	[0.04, 0.07]	0.02	[0.01, 0.03]	0.03	[0.01, 0.05]
ρ_R	0.60	[0.47, 0.69]	0.72	[0.59, 0.81]	0.72	[0.65, 0.78]	0.71	[0.65, 0.76]
				Real Exchange Rate Targeting				
ψ_1	1.79	[1.36, 2.29]	2.21	[1.85, 2.81]	1.98	[1.55, 2.47]	2.19	[1.71, 2.58]
ψ_2	0.18	[0.09, 0.29]	0.23	[0.14, 0.33]	0.29	[0.11, 0.46]	0.15	[0.05, 0.22]
ψ_3	-0.13	[-0.21, -0.07]	-0.24	[-0.34, -0.12]	-0.14	[-0.20, -0.08]	0.13	[0.06, 0.23]
ρ_R	0.59	[0.48, 0.69]	0.68	[0.62, 0.75]	0.70	[0.61, 0.79]	0.72	[0.63, 0.81]

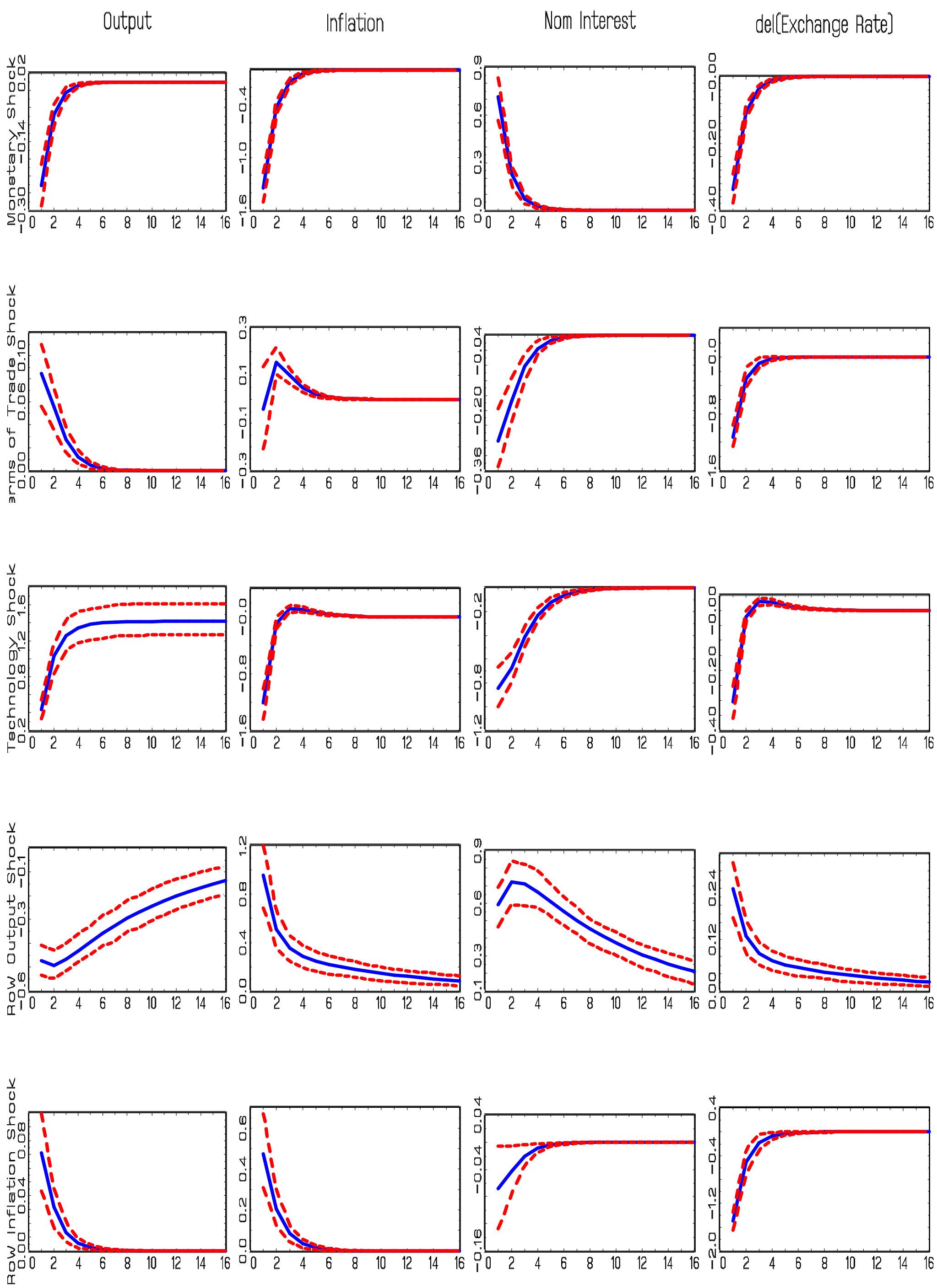
Notes: The table reports posterior means and 90 percent probability intervals (in brackets). The posterior summary statistics are calculated from the output of the posterior simulator.

Table 8: Marginal Data Densities and Posterior Odds Tests

	Marginal Data Densities				
	Pref. Rule	Exp. Infl. $\psi_3 \geq 0$	Exp. Infl. $\psi_3 = 0$	MCI	R.E.R.
Australia	-867.37	-872.18	-869.04	-871.99	-871.42
Canada	-707.39	-710.98	-715.26	-711.47	-705.72
New Zealand	-621.15	-624.37	-623.74	-623.52	-623.80
UK	-764.02	-769.55	-765.26	-768.10	-762.80

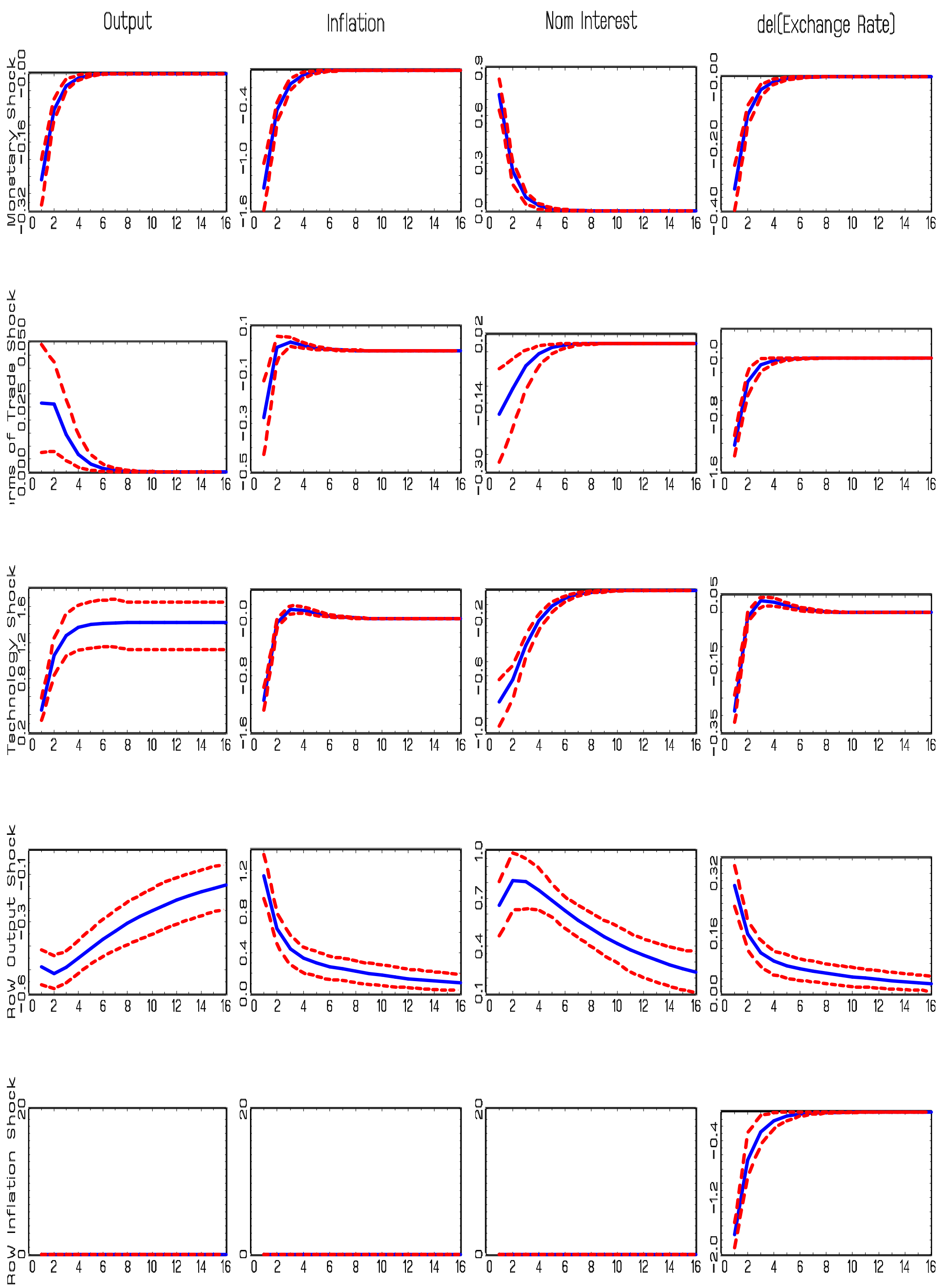
Notes: The Table reports posterior odds tests of the previously accepted policy rule against the most likely alternative. The posterior probabilities are calculated based on the output of the Metropolis algorithm. Marginal data densities are approximated by Geweke's (1999) harmonic mean estimator.

Figure 1: Impulse Responses for Canada: $\psi_3 \geq 0$



Notes: Figure depicts posterior means (solid lines) and pointwise 90% posterior probability intervals (dashed lines) for impulse responses of output, inflation, the nominal interest rate, and exchange rate changes to one-standard deviation structural shocks.

Figure 2: Impulse Responses for Canada: $\psi_3 = 0$



Notes: Figure depicts posterior means (solid lines) and pointwise 90% posterior probability intervals (dashed lines) for impulse responses of output, inflation, the nominal interest rate, and exchange rate changes to one-standard deviation structural shocks.