Banco Central de Chile
Documentos de Trabajo

Central Bank of Chile
Working Papers

N° 352

Diciembre 2005

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MONETARY POLICY, EXCHANGE RATE AND INFLATION INERTIA IN CHILE: A STRUCTURAL APPROACH

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Resumen

En este documento se estima un modelo de equilibrio general dinámico para Chile durante el período de metas de inflación. En este contexto se investiga la manera en que la política monetaria ha sido diseñada, y se evalúa el desempeño de reglas óptimas bajo distintas preferencias que el Banco Central pueda tener. Finalmente, se estudia en qué medida la curva de Phillips puede ser representada por una especificación híbrida que considere persistencia inflacionaria. Se concluye que la conducción de la política monetaria en Chile durante los últimos quince años se caracteriza por una regla de política en la que la tasa de interés reacciona a los desvíos de la inflación contemporánea respecto de la meta. Adicionalmente, la tasa de política reacciona a desvíos del producto respecto de su nivel natural. Esta regla de política presenta un grado importante de inercia que, sin embargo, ha disminuido en los últimos cinco años. Además, las fluctuaciones del tipo de cambio nominal no generan una respuesta sistemática de la autoridad monetaria: la reacción de la tasa de interés a movimientos de esta variable no es distinta de cero. Un segundo set de resultados indica que la persistencia inflacionaria, que habitualmente está ausente de los modelos neokeynesianos tradicionales, es una característica de la economía chilena. En particular, un modelo en que la curva de Phillips contiene un término para la inflación rezagada es mejor que uno sin persistencia inflacionaria. Esta persistencia ha cambiado en los últimos años, y es hoy menos importante que a principios de los noventa. Por otro lado, desde 1999 los precios se han ido ajustando de forma menos frecuente: la probabilidad de reoptimizar precios ha disminuido. Por último, reglas óptimas de política para preferencias alternativas del Banco Central muestran que, además de reaccionar a desvíos de la inflación y del producto, al incorporar una respuesta de política al tipo de cambio no se produce una ganancia de bienestar.

Abstract

This paper estimates a DSGE model for Chile during the IT period using Bayesian techniques. In this setup, we investigate the way in which monetary policy has been designed. We also assess the performance of simple optimal rules under alternative preferences that the central bank may have, and we investigate whether the Phillips curve can be represented by a hybrid specification which considers inflation persistence. We conclude that the conduct of monetary policy in Chile during the last fifteen years can be characterized by a feedback rule in which the interest rate reacts to contemporaneous inflation misalignment from the target and to output deviations from its natural level. This policy presents an important degree of persistence, that has, however, declined in the past five years. Furthermore, fluctuations in the nominal exchange rate are not offset by the monetary authority: the interest rate response to movements in this variable is not different from zero. A second set of results indicates that inflation persistence, which is usually absent from the standard neo-Keynesian models, is a feature of the Chilean economy. In particular, a model in which the Phillips equation contains a lagged inflation term is preferred to an alternative one which does not consider inflation persistence. This inflation persistence has change in recent years, becoming less important than in the early nineties. On the other hand, since 1999 prices are adjusted less frequently: the probability of resetting prices has fallen. Finally, optimal simple rules for alternative preferences of the central bank show that, besides reacting to inflation and output, there are no welfare gains from reacting to exchange rate movements.

A previous version of this paper was presented at the First Monetary Policy Research Workshop in Latin America and the Caribbean on Monetary Policy Response to Supply and Asset Price Shocks, Santiago, Chile, November 17-18, 2005. The authors thank Klaus Masuch and Rodrigo Cerda for helpful comments and discussions. We are grateful to Juan Pablo Medina and Claudio Soto for fruitful discussions and suggestions. We thank Luis Felipe Céspedes, Igal Magendzo, Simon Potter, Klaus Schmidt-Hebbel and the rest of the conference participants for helpful comments. Finally, we express our gratitude to Thomas Lubik and Frank Schorfheide for sharing with us some of the GAUSS codes used in this paper. The views expressed in this paper are solely the responsibility of the authors and should not be interpreted as reflecting the views of the Board of the Central Bank of Chile. Any remaining mistake or misinterpretation is our own.
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1 Introduction

In recent years, a large number of small open economies have adopted inflation targeting as a framework for monetary policy. In this context, one of the issues that still remains unsettled is the role of the exchange rate in a monetary regime characterized by a flexible exchange rate, an inflation target and a monetary-policy rule (Taylor 2001). Recent theoretical models find no role for the exchange rate. In particular, Clarida *et al* (2001) show that the optimal monetary policy reaction function should not react to the exchange rate. In this model, the representative household welfare criterion depends on the variance of three elements, domestic inflation, the output gap and the real exchange rate. However, because the real exchange rate is proportional to the output gap, such a criterion depends, in the end, only on domestic inflation and the output gap variances. As a consequence, the real exchange rate becomes irrelevant for monetary policy decisions. Similar conclusions are found in Gali and Monacelli (2005).

Despite the theoretical prescriptions, the empirical evidence on the role of the exchange rate is mixed. Using a single equation approach, Clarida *et al* (1998) show that the monetary authorities in some European countries and Japan respond to exchange rate misalignments. This reaction is statistically significant although it is not quantitatively important. Using a similar approach for Chile, Schmidt-Hebbel and Tapia (2002) and Caputo (2005) find that the relative size of this response is larger. Similarly, Calvo and Reinhart (2002) conclude that many emerging economies use the interest rate as the preferred means of smoothing exchange rate fluctuations. In this case, the "fear of floating" induces many central banks to move interest rates aggressively in response to exchange rate fluctuations. On the other hand, Lubik and Schorfheide (2003) using a Bayesian approach to estimate a small open economy structural model, find that the central banks of Australia, New Zealand and the UK did not explicitly respond to exchange rate. On the contrary, the bank of Canada did.

In this line of research, some studies investigate the optimally of policy rules that react to exchange rate in the context of small macromodels (usually calibrated) for open economies. In general this approach involves the estimation of a particular policy rule that is then introduced in a calibrated model. In this setup, a policy rule that generates a lower volatility in the variables of interest is considered to be "better". In this line of research, Batini *et al* (2003) conclude that an optimal policy rule for the UK should contain a response to the real exchange rate, but only marginal gains are derived from responding to it. In calibrated models for small open economies, Leitemo and Sodestrom (2003) show that responding to the exchange rate brings only marginal gains. One drawback of the previous approach is that it does not take into account the cross-correlation between the policy reaction function and the rest of the equations in the economy. In particular, the policy rule coefficients derived from a single equation estimation are not necessarily coherent with the calibrated - or estimated- coefficients that characterize the rest of the economy. This makes it difficult to assess the plausibility of alternative policy rule specifications. In order to overcome this limitation, recent studies have jointly estimated the coefficients of simple Taylor-type rules along with the structural coefficients and shocks that characterize the economy.

In this context, the objective of this paper is twofold. First, we estimate the monetary policy rule in a general equilibrium model in order to understand the way in which monetary policy has been design in Chile during the IT period. In particular, we want to know: i) whether the interest rate has reacted to lagged or expected inflation, ii) what is the role of output and iii) to what extent the Chilean central bank has reacted to exchange rate movements. In this setup, we also investigate the importance of inflation inertia. Second, we derive monetary policy frontiers in order to assess whether reacting to the exchange rate movement contributes to attenuate the volatility of output and inflation in the face of different structural shocks.

In order to estimate the policy reaction function in a general equilibrium framework, we follow Lubik and Schorfheide (2003) and estimate a dynamic stochastic general equilibrium model for Chile during the targeting period. The advantage of this multivariate approach is that it enables us to exploit the cross-equation restrictions that link agents´ decision rules to the policy coefficients. In doing so, we use Bayesian estimation technics, in particular we assign prior distributions to reaction function and other structural parameters and conduct Bayesian inference. Posterior probabilities are used to assess the adequacy of alternative policy rules and Phillips curve specifications. This approach allows us to compare both nested and non-nested policy rules such as inflation versus expected inflation targeting with and without an explicit reaction to exchange rate. It makes also possible to compare alternative specifications for the Phillips equation. In particular, we can test whether the standard forward-looking specification is preferred to the hybrid model that considers inflation persistence. There are other studies that estimate DSGE models for Chile using Bayesian technics. One is Caputo et al (2005), and the other is Medina and Soto (2005). Although we follow the same econometric methodology, our model is much more compact and is focuses on some structural parameters of the policy reaction function and the Phillips equation.

Once the model has been estimated, we derive monetary policy frontiers. In particular, we follow Levin et al (1999) and obtain the coefficients in the policy reaction function that minimize a welfare criterion that the central bank may have. In this way, it is possible to assess the performance of alternative policy rules and to see whether it is optimal to introduce and explicit policy response to the exchange rate.

The main conclusions of the paper can be summarized as follows. First, there is no evidence that the central bank of Chile has reacted, besides inflation and output, to exchange rate movements. In particular, models that introduce such a response have a worse fit than restricted models that set this response to be zero. The only period in which the central bank reacted to exchange rate movements is 1998. Second, inflation persistence is a relevant feature of the Chilean economy, but it has declined in the period 1999 tom 2005 where inflation has stabilized in a one digit level. Third, price stickiness has increased since 1999. In particular, the probability of resetting prices has fallen in the 1999 to 2005 period. Fourth, an optimal simple rule considers a policy reaction to inflation and output that is different from zero, however, the optimal response to asset price movements -exchange rate fluctuations- is zero. This result holds independently of the relative importance that the central bank gives to output and inflation volatility. In other words, reacting to exchange rate, in addition to the policy response to inflation and output, does not attenuate inflation and output volatility.

This paper is organized as follows. In Section 2, a structural model for a small open economy is specified for Chile. This is a rational expectations model based on Gali and Monacelli (2005) and Lubik and Schorfheide (2003). This simple model is derived from first principles. We extend the basic setup to consider inflation inertia as in Cespedes et al (2005). Section 3 presents the estimation strategy and the empirical implementation. Section 4 describes the data used in this exercise and discusses the choice of priors. Section 5 presents the results and test whether the Chilean central bank has reacted to exchange rate misalignments and to what extent the hybrid Phillips equation is supported by the data. Section 6 derives monetary policy frontiers for simple policy rules. Finally, section 7 concludes the paper.

2 A Structural Model for a Small Open Economy

As is noted by Dennis (2003), most of the micro founded models used in empirical research are calibrated, not estimated. Moreover, these models are tailored to reflect the characteristics of developed countries, limiting their applicability to small and emerging economies. In this section, we present a microfounded model that is then estimated for the Chilean economy. As is noted by Batini *et al* (2003), using a microfounded model enables the researcher to identify the structural shocks that the economy has faced[1]. In addition, since the model has microfoundations, the structural coefficients that characterize the economy are independent from the monetary policy.

Following Gali and Monacelli (2005) and Lubik and Schorfheide (2005) we lay down a model for a small open economy that is coherent with microfoundations. In particular, the aggregate demand and supply equations could be derived from the optimizing behavior of consumers and firms. As is common in the literature[2], some of the exogenous processes are allowed to follow an autorregressive process of order one. Finally, the model is closed with the introduction of a monetary policy reaction function. The model is represented by the following equations.

The open economy IS curve is :

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

where variables are expressed as deviation from the steady state. The variable y_t represents the domestic output, R_t is the nominal interest rate, π_t is the inflation rate, A_t is an exogenous

[1] In a less structural macro-model, the observed shocks may be a combination of structural shocks. Hence, it is difficult to analyze them.

[2] See Svensson (2000) and Leitemo and Soderstrom(2003).

technological shock, y_t^* is the foreign output and q_t represents the real exchange rate, defined as the ratio between foreign and domestic prices, both expressed in units of the domestic currency. . As noted by Lubik and Schorfheide (2005), the real exchange rate enters in first differences form since it is changes in relative prices that affect inflation via the definition of the consumption based price index. On the other hand, $0 < \alpha < 1$ is the import share and $\tau^{-1} > 0$ represents the intertemporal substitution elasticity. When $\alpha = 0$, equation (1) collapses to its closed economy variant.

In this IS equation, the world output shock drops out of the system when $\tau = 1$, and as noted by Lubik and Schorfheide (2005) this a useful benchmark case. In particular, it depends on the assumptions of international risk sharing and the equality of the of intertemporal and intratemporal substitution elasticities. In this case, the trade balance is equal to zero for all time periods, and the economy is isolated from world output fluctuations.

The open economy Phillips curve is:

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

where $\bar{y}_t = -\alpha(2-\alpha)(1-\tau)\frac{1-\tau}{\tau}y_t^*$ is potential output in the absence of nominal rigidities and when technology is stationary. On the other hand, $\beta = \exp(-r/400)$ where r is the steady state home real interest rate. As before, when $\alpha = 0$ we obtain the close economy variant. The slope coefficient, $k > 0$, is a function of underlying structural parameters, such as labor supply and demand elasticities and parameters measuring the degree of price stickiness. Following Lubik and Schorfheide (2005) we treat this as structural parameter since we do not have additional information from the underlying model.

We introduce the nominal exchange rate, e_t, via the definition of the CPI. In particular, assuming PPP holds, we can express inflation as a function of the nominal exchange rate, the real exchange rate an the level of foreign inflation as follows:

$$\pi_t = \Delta e_t - \Delta q_t + \pi_t^* \tag{3}$$

We assume the monetary policy is described by an interest rate rule, where the central bank adjusts its instrument in response to deviations of CPI inflation from target and output from its potential level. In addition, we allow for the possibility of including exchange rate considerations into the policy reaction function. Finally, we introduce a smoothing coefficient that reflects the degree of persistence in the policy instrument. The policy rule can be expressed as:

$$R_t = \rho R_{t-1} + (1-\rho)[\psi_1\pi_t + \psi_2\left(y_t - \bar{y}_t\right) + \psi_3\Delta e_t] + \varepsilon_t^R \tag{4}$$

where ε_t^R is a monetary policy shock. It is assumed that the policy coefficients, $\psi_1, \psi_2, \psi_3 \geqslant 0$. On the other hand, the persistence coefficient, ρ, is between zero and one. A question we address empirically is to what extent the Chilean central bank has reacted to exchange rate on top of its response to output and inflation. This is equivalent to testing the null hypothesis that $\psi_3 = 0$. In this setup this test is performed by evaluating the Bayes factor which is an indicator that enables us to compare alternative models (more on this in section 3). We will also consider alternative specifications in which the monetary authority reacts to expected inflation, and to the real exchange rate.

On the other hand, as in Lubik and Schorfheide (2005), we treat the real exchange rate as an exogenous variable. In particular, it can be expressed as an autoregressive process of order one

$$\Delta q_t = \rho_q \Delta q_{t-1} + \varepsilon_t^q \tag{5}$$

An alternative to this specification is to express the real exchange rate as proportion of output growth differentials:

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

This specification, despite being theoretically sound, imposes a very tight condition on the dynamic of the real exchange rate. In this case, the exchange rate may become redundant as information variable: the exchange rate behavior will be determined by the output dynamics which is already considered in the policy reaction function. Hence, by construction the real exchange rate maybe redundant in both the estimated policy reaction function and the optimal policy response that determines the policy frontier. It is because of this that we introduce the exchange rate as an exogenous variable. Finally, in this setup it is assumed that the exogenous variables, follow an autoregressive process of order one:

$$A_t = \rho_A A_{t-1} + \varepsilon_t^A \tag{7}$$

$$y_t^* = \rho_{y^*} y_{t-1}^* + \varepsilon_t^{y^*} \tag{8}$$

$$\pi_t^* = \rho_{\pi^*} \pi_{t-1}^* + \varepsilon_t^{\pi^*} \tag{9}$$

While the above assumptions about the shocks processes may seem arbitrary, the Bayesian approach used here enables us to test whether the structural shocks follow an autoregressive process like the one suggested here.

2.1 Alternative Specifications

Inflation Persistence The theoretical model presented so far does not consider inflation inertia. However, inflation persistence seems to be a feature that characterizes the inflation process not only in Chile - as documented by Cespedes et al (2005)- but also in the United States - see Clarida and Gertler (1999). A way in which we take into account the inflation inertia is by modifying the Phillips equation in (2). In particular, we follow Cespedes et al (2005) and consider that there is a fraction of firms that set prices according to past inflation. In this way equation (2) can be reformulated as follows

$$\pi_t = \frac{\beta}{(1+\beta k_1)} E_t \pi_{t+1} + \frac{k_1}{(1+\beta k_1)} \pi_{t-1} + \frac{\left[-\frac{\alpha\beta\Delta q_{t+1}}{(1-\alpha)} + \frac{\alpha \Delta q_t}{(1-\alpha)} + \frac{k}{\tau+\alpha(2-\alpha)(1-\tau)} \left(y_t - \overline{y}_t \right) \right]}{(1+\beta k_1)} \tag{10}$$

were the k_1 coefficient represents the share of backward looking firms that do not adjust their prices looking at the inflation target. Whether equation (10) is a better specification for the inflation process in Chile is an empirical question. This issue is address by computing the Bayes factor in order to see whether the null $H_0 : k_1 = 0$ is rejected by the data.

Inflation Forecast Based Taylor Rule It has been argued that, instead of reacting to current inflation, central banks move their instrument in reaction to expected inflation -see Batini and Haldane (1999)-. The rationale for this type of reaction is that central banks cannot affect output nor inflation in the short-run. Hence, in order to have the desired effect on output and inflation it is better to anticipate movements on those variables. This type of reaction can be synthesized by the following expression:

$$R_t = \rho R_{t-1} + (1-\rho)[\psi_1 E_t(\pi_{t+1}) + \psi_2 \left(y_t - \overline{y}_t \right) + \psi_3 \Delta e_t] + \varepsilon_t^R \tag{11}$$

where ψ_2 and ψ_3 capture the response of the central bank to movements in output and exchange rate over and above their impact on future inflation. Therefore, as suggested by Clarida (2001), ψ_2 and ψ_3 can be interpreted as genuine responses to output and exchange rate misalignments.

3 Econometric Methodology

Equations (1) to (9) form a linear rational expectation model in the variables $s_t = [y_t, \pi_t, R_t, \Delta e_t, \Delta q_t, u_t^q, A_t, y_t^*, \pi_t^*]$. Following Sims (2002), the log-linearized DSGE model can be written as a system of the form

$$\Gamma_0(\theta) s_t = \Gamma_1(\theta) s_{t-1} + \Gamma_\epsilon(\theta)\epsilon_t + \Gamma_\eta(\theta)\eta_t \tag{12}$$

where θ is the vector of structural coefficients, ϵ_t stacks the innovations of the exogenous processes and η_t is composed of rational expectation forecast errors. The solution to (12) can be expressed as

$$s_t = \Phi_1(\theta) s_{t-1} + \Phi_\epsilon(\theta)\epsilon_t \tag{13}$$

A measurement equation then relates the model variables s_t to a vector of observables, y_t :

$$y_t = A(\theta) + B s_t \tag{14}$$

I our case, the vector of observable variables is given by $y_t = [y_t, \pi_t, R_t, \Delta e_t, \Delta q_t]$ whereas the non observable variables are u_t^q, A_t, y_t^* and π_t^*. Now, given $Y^T = \{y_1, ..., y_T\}$ we obtain the likelihood function $L(\theta/Y^T)$ that can be evaluated using the Kalman filter.

As in Lubik and Schorfheide (2005), we adopt a Bayesian approach and place a priori distribution with density $p(\theta)$ on the structural parameters. Simply stated, the Bayesian approach works as follows. First, it places a priori distribution with density $p(\theta)$ on the structural parameters, θ. Then, the data Y^T are used to update the prior through the likelihood function, $L(\theta/Y^T)$, in order to obtain the posterior distribution of θ. According to the Bayes Theorem, this later distribution, $p(\theta/Y^T)$, takes the form

$$p(\theta/Y^T) = \frac{L(\theta/Y^T)p(\theta)}{p(Y^T)} \tag{15}$$

Draws from this posterior can be generated through Bayesian simulation techniques (more on this in section 3.1). Based on these draws it is possible to compute the summary statistics (posterior means and 90% probability intervals) that characterize the structural coefficients.

Now in order to compare alternative model specifications, we make use of the marginal likelihood function. This is the probability that the model assigns to having observed the data. It is defined as the integral of the likelihood function across the parameter space using the prior as the weighting function:

$$p(Y^T/H_i) = \int L(\theta/Y^T, H_i)p(\theta/H_i)d\theta \tag{16}$$

where $p(Y^T/H_i)$ is the probability of having observed the data under model specification H_i, whereas $L(\theta/Y^T, H_i)$ and $p(\theta/H_i)$ are, respectively, the likelihood function and the prior distribution under model specification H_i. A natural way of assessing which model is more plausible, is to construct the ratio of the marginal likelihood function under alternative model specifications. This ratio is known as the Bayes factor and takes the form:

$$B_{i,j} = \frac{p(Y^T/H_i)}{p(Y^T/H_j)} \tag{17}$$

where $B_{i,j}$ is the Bayes factor of model i over model j. As is clear, if $B_{i,j} > 1$, model i is more plausible than model j and vice versa. Since we are unable to obtain the marginal likelihood function in a closed-form we estimate it as in Geweke (1998) and Rabanal and Rubio-Ramírez (2005). In particular, we integrate over the draws used to construct the posterior distribution. These draws are generated through the Metropolis-Hasting algorithm.

3.1 Posterior Distribution

In order to derive the posterior distribution of the coefficients, we proceed in two steps. First, we find the posterior mode, which is the most likely point in the posterior distribution, and compute the Hessian at the mode. In doing so, we use a standard optimization routine[3]. In this case the likelihood function is computed by first solving the model and then using the Kalman filter. Second, we implement the Metropolis-Hasting algorithm to generate draws from the posterior. The algorithm generates a sequence of draws that is path dependence and it works as follows:

1. Start with an initial value of the parameters, say θ^0. Then, compute the product of the likelihood and the prior at this point:$L(\theta^0/Y^T)p(\theta^0)$

2. Now, from θ^0 generate a random draw, θ^1, such that $\theta^1 = \theta^0 + v^1$. Where v^1 follows a multivariate normal distribution and the variance-covariance matrix of v^1 is proportional to the inverse Hessian of the posterior mode. Then, for θ^1 compute $L(\theta^1/Y^T)p(\theta^1)$.

The new draw, θ^1, is accepted with probability R and is rejected with (1-R), where $R = \min\left\{1, \frac{L(\theta^1/Y^T)p(\theta^1)}{L(\theta^0/Y^T)p(\theta^0)}\right\}$

If the draw is accepted, it is possible to generate another draw, $\theta^2 = \theta^1 + v^2$ and assess whether this second draw is accepted or not. On the contrary, if the draw is rejected, we go back to the initial value, θ^0, and generate another draw. The idea of this algorithm is that, regardless of the starting value, more draws will be accepted from the regions of the parameter space where the posterior density is high. At the same time, areas of the posterior support with low density are less represented, but will eventually be visited.

4 Data Description and Choice of the Prior

Our analysis is based on quarterly data for the Chilean economy. The time series begin in 1990Q3 with the introduction of inflation targeting and span to 2005Q1. The output series is computed as the percentage deviation of real GDP from trend, the inflation variable is constructed as the deviation of core CPI inflation from the predetermined target. The nominal and real exchange rate enter as a first difference of the seasonally adjusted series. Finally, we consider data on the real indexed interest rate.

Choosing the priors it is not an easy task and requires an important degree of judgment. In any case, we choose them based on both evidence from previous research and national accounts data (Table 1). For the policy reaction to inflation, the ψ_1 coefficient, we choose a gamma distribution with mean of 1.0 and a standard deviation of 0.25. As a result, the prior distribution takes into account the range of values that have been reported for this coefficient in Schmidt-Hebbel and Tapia (2002) and Caputo (2005). The policy reaction coefficient to output and to exchange rate, ψ_2 and ψ_3, have also a gamma distribution with mean 0.25 and a standard deviation of 0.13, this gives a range for this coefficient that is roughly coherent with the available evidence (again see Schmidt-Hebbel and Tapia (2002) and Caputo (2005)). The persistence parameter, ρ_R, has

[3]The csmiwel command in GAUSS.

a mean of 0.5 and a standard deviation of 0.2. This is in line with previous research indicating that this coefficients is around 0.7.

On the other hand, the degree of openness, α, is set to 0.3 with a very low standard deviation. This is coherent with the average ratio of imports to GDP computed based on national accounts information. The steady state real interest rate, R has a prior confidence interval of that ranges from 1.1 to 4.4 percent. For the k and τ coefficients, we don't have information on previous research, hence we choose a wide confidence interval centered on 0.5. For the inflation persistence coefficient, k_1, we choose a wide range that is coherent with the results of Cespedes *et al* (2005). Finally, for the autorregressive parameters and the coefficients capturing the standard deviation, we follow Lubik and Schorfheide (2005). In particular, we choose prior distributions reflecting some degree of persistence in the shocks with a standard deviation that is allowed to vary substantially.

5 Results[4]

5.1 Inflation Persistence

In the first exercise we estimate a model that allows for both a response to exchange rate and inflation inertia. The results, presented in second column of Table 2, show that monetary policy has responded aggressively towards inflation. In particular, the ψ_1 coefficient is 1.57. On the other hand, the policy response to output and exchange rate, the ψ_2 and ψ_3 coefficients respectively, are positive but much more smaller than the policy response to inflation. Furthermore, the response to the exchange rate is less important than the reaction to output. On the other hand, the degree of interest rate inertia, the ρ_R coefficient is 0.53. Those result are, overall, in line with previous studies for Chile, although the policy response to inflation is in this case much more important.

The degree of openness, captured by the α coefficient is 0.20. The slope of the Phillips curve, the k coefficient, is above the prior mean whereas the τ coefficient suggest that τ^{-1}, the intertemporal substitution elasticity, is nearly 4 which is above the value found in for some developed countries in Lubik and Schorfheide (2005).the persistence coefficient, κ_1 is 0.86 which is somehow bigger than the value estimated by Cespedes *et al* (2005). The long-term real interest rate, R is nearly 2.6%. On the other hand, the degree of persistence in the real exchange rate is small; around 0.2 whereas the persistence in foreign output and foreign inflation shocks is 0.98 and 0.47 respectively. Finally, technology shocks are mor persistent than the value suggested by the prior.

In the second exercise, we remove the inflation persistence from the model. In particular, we set k_1 to zero. The results from estimating this specification are presented in the third column

[4]The results in this section hold, also, when the CPI inflation is used instead of the core inflation measure. Results using CPI inflation are not presented, but are available upon request.

of Table 2. As we can see, in general, the structural parameters do not change significantly. In this case there is still a strong policy response to inflation whereas the response to output and exchange rate continues to be much more smaller. The main consequence of having removed inflation persistence is that the overall fit of the model worsen. In particular, the log likelihood function has a value that is well below the one for a model in which $k_1 > 0$. In this case, the Bayes factor is 26, 80. Hence, we can conclude that inflation persistence is a feature of the Chilean inflation process.

5.2 Alternative Policy Rule Specifications

Having established that inflation inertia should be consider, we now investigate the plausibility of alternative specification for the policy reaction function. In the first exercise, we remove the assumption that the Chilean central bank reacts to exchange rate movements. In this case, the ψ_3 coefficient takes a value of zero. As we can see in Table 3 second column, in general, results do not change significantly from the case in which $\psi_3 > 0$ (third column,Table 3). In particular, the policy response to inflation and output is almost the same. By the same token, the Phillips curve slope, the coefficient κ, remains unchanged. The steady state real interest rate, R, increases to nearly 3%. Overall, the model that does not consider a policy response to exchange rate devaluations is supported by the data. In particular, the log likelihood function is in this case -539.86 that is slightly above the value found under the specification in which $\psi_3 > 0$. To see whether the central bank has reacted to real exchange movements, rather than to nominal devaluations, we specify an alternative model, fourth column in Table 3, in which we allow for a policy response to Δq_t. Again, the overall results do not change significantly. Moreover, there data do not support this alternative specification: the model´s fit worsen both compared to the case in which $\psi_3 = 0$ and to the case in which $\psi_3 > 0$ and the policy is reacting .

Based on the above results, we can say that inflation inertia and a mute policy response to exchange rate movements are supported by the Chilean data. Given this fact, we investigate whether this result is robust to the way in which the policy rule is specified. In particular, we consider a model in which the policy reaction function is characterized by a Taylor-type rule that responds to expected rather than to contemporaneous inflation. In doing so, we consider a specification for the policy rule like the one presented in equation (11). In the first exercise of this type we set the ψ_3 coefficient to zero. In this case, when monetary policy is forward-looking, the response to inflation seems more aggressive than in the case in which monetary policy reacts to contemporaneous inflation (see fifth column, Table 3) This can be explained by the fact that a given shock to contemporaneous inflation is not completely transmitted to future inflation. Hence, a given policy response to a contemporaneous shock, under a rule that reacts to current inflation, can be replicated, in a forward-looking Taylor rule, only if the policy response to future inflation is larger. On the other hand, under the forward-looking policy rule, the rest of the coefficients are broadly similar to those under contemporaneous targeting inflation. This forward-looking

Taylor-type rule is, however, not supported by the data as we can conclude from the fact that the log-likelihood of the model decreases to -541,37.

Finally, we consider two alternative forward-looking policy rules. The first one, sixth column of Table 3, allows for a response to nominal devaluations, whereas the second one, seventh column of Table 3, entails a response to real exchange rate movements. As we can see, in both cases the data reject the models.

5.3 Subsample Analysis

From the previous exercise,we conclude that a Taylor-type policy rule with no reaction to exchange rate is the one that better characterize the Chilean data. Moreover, a rule that targets contemporaneous inflation has a better fit than one that targets expected inflation. Now, given the fact that monetary and exchange rate policies have changed over the years in Chile it is of interest to assess whether the general findings are robust over time. In doing so, we estimate the model under alternative sample periods. Doing this type of exercise allows both to identify potential structural changes or specific policy reaction to some particular shocks.

In the first exercise the model is estimated from 1990 to 1997. During this period there was in place an exchange rate band along with an explicit inflation targeting regime, furthermore in this period there were no significant external shocks as the ones that hit the Chilean economy in 1998. The estimation results are presented in the second and third column of Table 4 and indicate that during this period the policy interest rate did not react to nominal devaluations. In fact a model with no reaction to exchange rate movements, $\psi_3 = 0$, is preferred to a model with a policy reaction to exchange rate movements. In terms of the policy coefficients, the response to inflation is less aggressive: a relatively smaller value for ψ_1, and in this period the policy interest rate turn out to be more persistent: the ρ_R coefficient increases, from 0.48 in the whole sample, to 0.68 in this subsample period. On the other hand, the slope of the Phillips curve, the κ coefficient, increases from 0.75 in the whole sample to 1.64 in the period 1990 to 1997.

For the subsample 1998 to 2005, the results, presented in the fourth and fifth columns of Table 4, are quite different. First, a model with an explicit response to nominal devaluations is preferred to a model with no reaction. Second, the response to inflation and output is more aggressive than in the previous subsample. Third, monetary policy is much less inertial: the ρ_R coefficient decreases, from 0.48 in the whole sample, to 0.26 in this subsample period. Finally, the slope of the Phillips curve, the κ coefficient, decreases from 0.75 in the whole sample to 0.19 in the period 1998 to 2005. Now, to what extent are the results in this subsample determined by the 1998 events?. In particular, in that year interest rates increases substantially in order to avoid the negative impacts of the Asian and Russian crisis on the Chilean currency. To address this question, we reestimate the model but now excluding the year 1998. The results, presented in the last two columns of Table 4, tend to confirm the findings for the 1998 to 2005 period: there is a more aggressive policy response to inflation, a less persistent policy reaction and a smaller Phillips curve slope. There are, however, two important difference. First, when 1998 is excluded, the model that fits better the data is one that does not consider a policy response to inflation, $\psi_3 = 0$. Hence, it is possible to conclude that monetary policy reacted to nominal devaluations only in 1998. Besides this period there was no systematic reaction to nominal devaluations. Second, in the period 1999 to 2005 inflation persistence, characterized by the κ_1 coefficient, decreases.

Based on the above results, it is possible to conclude that policy became more aggressive toward inflation and less persistent in the latter subsample. On the other hand, in this period inflation became less persistent and the probability of keeping prices fixed, which is inversely related to the κ coefficient[5], increases. Hence in the latter subsample prices became more sticky than in the early nineties when inflation was higher. These results are also found in Céspedes et al (2005) and Céspedes and Soto (2005).In particular,those studies show that price rigidity has increased while at the same time the degree of indexation based on past inflation has decreased over time. According to them, these changes are related to credibility gains by the monetary policy regime that has improved its tradeoff.

6 Monetary Policy Frontiers

The previous results show that the Chilean central bank has not reacted to exchange rate movements. Moreover, the rule that fits the data better is one that reacts to contemporaneous inflation misalignments and output deviations from its steady state level . What are the implication, in terms of welfare, of such a policy?. In particular, is there any advantage of allowing for a policy response to exchange rate movements?. In order to address these normative questions, we derive monetary policy frontiers that show how output and inflation volatility behave under alternative policy rules. In doing so, we follow Levin et al (1999) and assume the interest rate rule is chosen to solve the following optimization problem:

$$\underset{(\psi_1,\psi_2,\psi_3)}{\text{Min}} \lambda var(y_t) + (1-\lambda) var(\pi_t) \tag{18}$$

subject to (13) and to $var(R_t) \leq k^2$. On the other hand, y_t indicates the output, π_t indicates the four-quarter average inflation rate, and var(s) indicates the unconditional variance of variable s. The weight $\lambda \epsilon [0,1]$ reflects the policymaker's preference for minimizing output volatility relative to inflation volatility. As in Levin et al (1999), we constrain the level of interest rate volatility by imposing the upper bound k on the standard deviation of the policy interest rate. In doing so, the benchmark value of k is set equal to policy rate volatility under the estimated policy rule. Also, throughout our analysis, we only consider policy rules that generate a unique stationary rational expectations solution. To compute the policy frontier for a particular functional form of the interest rate rule, we determine the parameters of this rule which maximize the objective

[5] Following Gali and Monacelli (2005), it is possible to express the Phillips curve slope as $\kappa = \frac{(1-\beta\theta)(1-\theta)}{\theta} + \varphi(\tau + \alpha(2-\alpha)(1-\alpha))$ where $(1-\theta)$ is the probability of resetting prices and φ represents the labor supply elasticity.

function for each value of λ over the range zero to unity[6]. Thus, for a given form of the interest rate rule, the policy frontier traces out the best obtainable combinations of output and inflation volatility, subject to the upper bound interest rate volatility. As is noted by Levin et al (1999), this approach differs slightly from that commonly found in the literature, in which interest rate volatility is incorporated into the objective function and each policy frontier is drawn using a different weight on interest rate volatility. Instead, we maintain a strict distinction between the policymaker's preferences and the constraints implied by the model.

For a particular functional form of the interest rate rule, we determine the policy frontier by solving the optimization problem in equation (18) for a range of values of the objective function weight λ . In particular, for a given λ, we start with an initial value for the rule parameters. Those values are obtained, through a simple grid search, as the values that minimize the objective function[7]. Then an optimization routine, the csminwel algorithm, is applied which iteratively updates the parameter vector until an optimum is obtained.

In order to see how neglecting a policy response to exchange rate may affect the objective function, we derive the policy frontiers under two sets of potential policy coefficients. The first one, the restricted case, considers positive values for both ψ_1 and ψ_2 but imposes $\psi_3 = 0$. The second one, the unrestricted case, considers positive values for ψ_1, ψ_2 and ψ_3. Hence, by comparing the policy frontiers under the two cases, we can assess whether reacting to exchange rate movements is indeed welfare improving. To compute the policy frontier we stochastically simulate the model for all the structural shocks. However, as in Levin et al (1999), we exclude monetary policy innovations from this analysis. The results of this exercise, for alternative values of λ and for the case in which ψ_3 can be different from zero, are presented in Figure 1. As we can see, when policy is more concerned with reducing inflation volatility, a smaller λ, the inflation volatility is reduced at the cost of inducing a higher variation in output (figure 1). In terms of the optimal coefficients, we find that the optimal response to exchange rate movements turns out to be zero for all values of λ. Hence, the policy frontier when $\psi_3 = 0$ is the same as the policy frontier obtained when ψ_3 is allowed to take any positive value. Now, in terms of the optimal policy coefficients we see that when there is more concern about output volatility, a higher λ, the optimal response to output increase whereas the response to inflation goes down (figure 2). This result is quite standard in small open economies -see Levin et al (1999)-. For Chile, this results also holds and, more importantly, as mentioned before the optimal response to exchange rate movements turns out to be zero for any λ.

The results present so far suggest two things. First, in practice the Chilean central bank has not reacted to exchange rate movements and, simple optimal rules should not react to nominal devaluations, independently of the policymaker's preference for minimizing output volatility relative

[6] In all the exercises, we use the structural coefficients estimated under the more plausible model: full sample specification in Table 3 with no reaction to exchange rate movements. The only parameters that change are the policy rule coefficients, ψ_1, ψ_2 and ψ_3.

[7] In general, initial values are set in an arbitrary way. Determining them through a simple grid search ensures that the optimization routine is initialized in a region that is closer to the optimal coefficient values.

to inflation volatility. The reason why introducing a policy response to nominal devaluations does not reduce output or inflation volatility can be understood by analyzing the dynamic properties of alternative policy rules. The first policy rule is the one that imposes $\psi_3 = 0$ and correspond to the full sample specification. The second policy rule is imposes a value of $\psi_3 = 0.7$. The impulse response functions that each rule generates, in the face of structural shocks, are presented in Figure 3. In the face of a positive real exchange rate shock (second row in Figure 3), the nominal exchange rate depreciates. If the central bank reacts to this devaluation the policy interest rate will increase more than in the case in which $\psi_3 = 0$. As a result, output will experience a stronger contraction. In this case, inflation will not increase initially but will undershoot its target level after some quarters. Hence, in the face of real exchange rate shocks, a policy rule that respond to exchange rate devaluations will generate more output volatility without necessarily attenuating inflation fluctuations. Now, consider a shock to the rest of the world output level, $\varepsilon_t^{y^*}$, in this case this shock will generate a real exchange appreciation. This in turn, will imply a reduction in the nominal exchange rate. If the central bank reacts to the exchange rate, then the interest rate will increase but much less than in the case in which $\psi_3 = 0$. As a result, output will contract by less, but inflation is going to be above its target level for longer. As a result, in the face of a shock to the rest of the world output level, $\varepsilon_t^{y^*}$, inflation will be much more volatile increasing the welfare losses.

7 Conclusions

This paper estimates a DSGE model for Chile using Bayesian technics. This approach has the advantage of combining prior information about the structural coefficients with the likelihood function generated by the prior distribution of the coefficients. As a result, it is possible to obtain posterior probability distributions that are used to assess the adequacy of alternative policy and the validity of different specifications for the inflation equation. We do so in order to characterize the conduct of monetary policy in Chile and to see which policy specification is more likely to reflect the Chilean data. In particular, we test whether a policy rule that reacts to exchange rate movements is more plausible. Also, we investigate whether a forward-looking Taylor type rule is supported by the data. In this context, a second issue we address is the relevance of inflation inertia, a feature that has been excluded from standard New-Keynesian models. Finally, we assess the performance of alternative policy rules. In particular, we investigate what are the consequences, in terms of inflation and output volatility, of implementing a policy reaction function that respond, in addition to inflation and output, to exchange movements.

The main results indicate that the conduct of monetary policy in Chile during the last fifteen years can be characterized by a feedback rule in which interest rate reacts to contemporaneous inflation misalignment from the target and output deviations from trend. This policy presents an important degree of persistence in the early nineties. Furthermore, fluctuations in the nominal exchange rate are not offset by the monetary authority: the interest rate reaction is not different from zero (the exception being the 1998 year). On the other hand, a forward-looking specification is rejected by the data. A second sets of results indicates that: i) inflation persistence, which is usually absent from the standard New-Keynesian models, is a feature of the Chilean economy. In particular, it is not possible to reject the null hypothesis that the Phillips equation contains a lagged inflation term. Inflation persistence, on the other hand, has become less important in the 1999 to 2005 In this period price stickiness has increased.

Finally, we conclude that an optimal simple rule considers a policy reaction to inflation and output that is different from zero, however, the optimal response to asset price movements - exchange rate fluctuations- is zero. This result holds independently of the relative importance that the central bank gives to output and inflation volatility. In other words, reacting to exchange rate, in addition to the policy response to inflation and output, does not attenuate inflation and output volatility. This is particularly true in the face of real exchange rate shocks and innovations to the rest of the world output level.

Table 1. Priors

Name	Density	Mean / Mode	St.Dev / df	90% Interval	
ψ_1	Gamma	1.00	0.25	0.63	3.50
ψ_2	Gamma	0.25	0.13	0.09	0.48
ψ_3	Gamma	0.25	0.13	0.09	0.48
ρ_R	Beta	0.50	0.20	0.17	0.83
α	Beta	0.30	0.03	0.26	0.34
R	Gamma	2.50	1.00	1.11	4.43
k	Gamma	0.50	0.25	0.17	0.97
τ	Gamma	0.50	0.20	0.22	0.87
k_1	Beta	0.50	0.20	0.00	0.99
ρ_q	Beta	0.40	0.35	0.04	0.96
ρ_A	Beta	0.20	0.10	0.06	0.39
ρ_{y^*}	Beta	0.90	0.05	0.81	0.97
ρ_{π^*}	Beta	0.70	0.15	0.43	0.92
σ_R	Inverse Gamma	1.00	4.00	0.64	3.66
σ_q	Inverse Gamma	2.00	4.00	1.29	7.32
σ_A	Inverse Gamma	1.50	4.00	0.97	5.48
σ_R	Inverse Gamma	1.50	4.00	0.967	5.488
σ_{π^*}	Inverse Gamma	1.50	4.00	0.967	5.488

Table presents mean, standard deviation and 90% probability intervals
For inverse gamma distributions, mode and degrees of freedom are presented
For uniform distributions, upper and lower bounds are presented.

Table 2. Models with and without inflation inertia

Coefficient	With inertia $\kappa_1 > 0$	Without inertia $\kappa_1 = 0$
ψ_1	1.57	1.79
	(0.18)	(0.29)
ψ_2	0.44	0.37
	(0.17)	(0.18)
ψ_3	0.11	0.09
	(0.04)	(0.03)
ρ_R	0.53	0.52
	(0.08)	(0.06)
α	0.20	0.21
	(0.02)	(0.02)
R	2.58	2.02
	(0.69)	(0.71)
κ	0.88	0.86
	(0.21)	(0.36)
τ	0.25	0.28
	(0.06)	(0.05)
κ_1	0.86	0.00
	(0.08)	(0.00)
ρ_q	0.22	0.23
	(0.09)	(0.10)
ρ_A	0.57	0.59
	(0.04)	(0.06)
ρ_{y*}	0.98	0.97
	(0.01)	(0.01)
$\rho_{\pi*}$	0.47	0.55
	(0.10)	(0.07)
σ_R	0.60	0.61
	(0.08)	(0.07)
σ_q	2.81	2.89
	(0.20)	(0.20)
σ_A	0.90	0.90
	(0.10)	(0.18)
σ_{y*}	1.63	1.69
	(0.62)	(0.45)
$\sigma_{\pi*}$	2.11	2.11
	(0.16)	(0.16)
$log\ \widehat{L}$	-540.63	-550.83
Bayes factor		26,804

Table presents posterior mean and standard deviations in parenthesis

Table 3. Models with alternative policy rules

	Contemporaneous Inflation			Expected Inflation		
Coefficient	$\psi_3 = 0$	$\psi_3 > 0$	$\psi_3 > 0$ and Δq_t	$\psi_3 = 0$	$\psi_3 > 0$	$\psi_3 > 0$ and Δq_t
ψ_1	1.62	1.57	1.67	1.73	1.70	1.79
	(0.24)	(0.18)	(0.23)	(0.29)	(0.26)	(0.31)
ψ_2	0.34	0.44	0.43	0.24	0.36	0.47
	(0.15)	(0.17)	(0.15)	(0.10)	(0.16)	(0.18)
ψ_3	-	0.11	0.09	-	0.14	0.12
		(0.04)	(0.03)		(0.04)	(0.05)
ρ_R	0.48	0.53	0.50	0.51	0.50	0.52
	(0.08)	(0.08)	(0.09)	(0.07)	(0.08)	(0.08)
α	0.21	0.20	0.22	0.20	0.21	0.21
	(0.02)	(0.02)	(0.02)	(0.02)	(0.01)	(0.02)
R	2.87	2.58	2.48	2.62	2.06	2.17
	(0.98)	(0.69)	(1.11)	(0.94)	(0.66)	(0.78)
κ	0.75	0.88	0.84	0.70	0.69	0.55
	(0.16)	(0.21)	(0.21)	(0.23)	(0.21)	(0.12)
τ	0.25	0.25	0.27	0.25	0.26	0.22
	(0.04)	(0.06)	(0.04)	(0.04)	(0.07)	(0.03)
κ_1	0.86	0.86	0.78	0.88	0.87	0.85
	(0.06)	(0.08)	(0.10)	(0.05)	(0.05)	(0.09)
ρ_q	0.21	0.22	0.24	0.22	0.30	0.32
	(0.07)	(0.09)	(0.10)	(0.09)	(0.10)	(0.10)
ρ_A	0.55	0.57	0.60	0.61	0.64	0.61
	(0.04)	(0.04)	(0.04)	(0.04)	(0.05)	(0.05)
ρ_{y*}	0.97	0.98	0.98	0.97	0.97	0.98
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
$\rho_{\pi*}$	0.58	0.47	0.61	0.58	0.55	0.63
	(0.08)	(0.10)	(0.11)	(0.08)	(0.11)	(0.09)
σ_R	0.62	0.60	0.67	0.56	0.55	0.59
	(0.07)	(0.08)	(0.13)	(0.04)	(0.03)	(0.06)
σ_q	2.92	2.81	2.78	2.96	2.96	2.81
	(0.23)	(0.20)	(0.26)	(0.33)	(0.26)	(0.18)
σ_A	1.01	0.90	0.92	0.89	0.85	0.88
	(0.15)	(0.10)	(0.11)	(0.10)	(0.12)	(0.11)
σ_{y*}	1.50	1.63	1.60	1.49	1.62	1.25
	(0.28)	(0.62)	(0.33)	(0.25)	(0.44)	(0.26)
$\sigma_{\pi*}$	2.15	2.11	2.12	2.10	2.18	1.97
	(0.19)	(0.16)	(0.22)	(0.19)	(0.19)	(0.15)
$log\ L$	-539.86	-540.63	-541.25	-541.37	-540.84	-545.43

Table presents posterior mean and standard deviations in parenthesis

Table 4. Subsample Estimations

	1990 - 1997		1998 - 2005		1999 - 2005	
Coefficient	$\psi_3 = 0$	$\psi_3 > 0$	$\psi_3 = 0$	$\psi_3 > 0$	$\psi_3 = 0$	$\psi_3 > 0$
ψ_1	1.10	1.27	1.84	1.38	1.67	1.56
	(0.21)	(0.23)	(0.32)	(0.38)	(0.29)	(0.27)
ψ_2	0.21	0.26	0.29	0.43	0.23	0.25
	(0.08)	(0.09)	(0.11)	(0.19)	(0.11)	(0.09)
ψ_3	-	0.08	-	0.20	-	0.14
		(0.04)		(0.07)		(0.06)
ρ_R	0.65	0.70	0.32	0.26	0.41	0.41
	(0.08)	(0.08)	(0.09)	(0.09)	(0.12)	(0.11)
α	0.24	0.25	0.21	0.21	0.21	0.21
	(0.02)	(0.02)	(0.01)	(0.01)	(0.02)	(0.02)
R	2.08	3.05	2.50	2.86	2.14	2.76
	(0.69)	(1.11)	(1.08)	(1.04)	(0.94)	(0.93)
κ	1.64	1.48	0.51	0.19	0.73	0.50
	(0.39)	(0.28)	(0.15)	(0.05)	(0.22)	(0.14)
τ	0.40	0.38	0.38	0.33	0.36	0.35
	(0.06)	(0.08)	(0.06)	(0.05)	(0.06)	(0.05)
κ_1	0.68	0.57	0.85	0.89	0.33	0.47
	(0.15)	(0.20)	(0.08)	(0.05)	(0.13)	(0.16)
ρ_q	0.08	0.07	0.38	0.80	0.47	0.49
	(0.04)	(0.04)	(0.14)	(0.08)	(0.04)	(0.04)
ρ_A	0.36	0.37	0.59	0.60	0.93	0.93
	(0.04)	(0.04)	(0.04)	(0.06)	(0.03)	(0.03)
ρ_{y*}	0.88	0.89	0.93	0.91	0.55	0.54
	(0.05)	(0.05)	(0.03)	(0.04)	(0.12)	(0.10)
$\rho_{\pi*}$	0.50	0.59	0.49	0.47	0.56	0.56
	(0.10)	(0.10)	(0.12)	(0.10)	(0.13)	(0.07)
σ_R	0.40	0.39	0.84	0.86	3.16	2.89
	(0.04)	(0.04)	(0.10)	(0.13)	(0.48)	(0.32)
σ_q	2.49	2.67	2.78	3.14	1.06	1.07
	(0.20)	(0.31)	(0.37)	(0.40)	(0.14)	(0.14)
σ_A	1.04	1.13	1.22	1.23	1.35	1.36
	(0.18)	(0.16)	(0.14)	(0.18)	(0.31)	(0.24)
σ_{y*}	2.03	1.92	1.74	1.34	1.95	2.05
	(0.41)	(0.67)	(0.41)	(0.23)	(0.22)	(0.40)
$\sigma_{\pi*}$	1.96	2.07	2.05	1.87	0.86	0.83
	(0.17)	(0.33)	(0.17)	(0.17)	(0.09)	(0.07)
log L	-256.60	-263.86	-260.85	-258.03	-209.37	-210.08

Table presents posterior mean and standard deviations in parenthesis

Figure 1. Policy frontiers

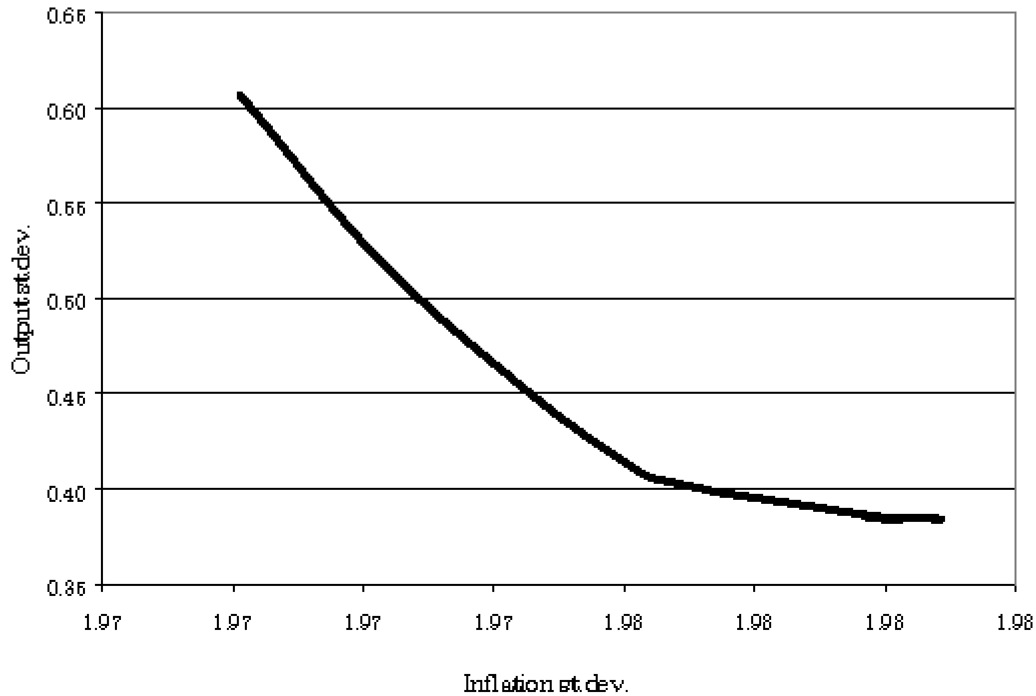


Figure 2. Policy coefficients

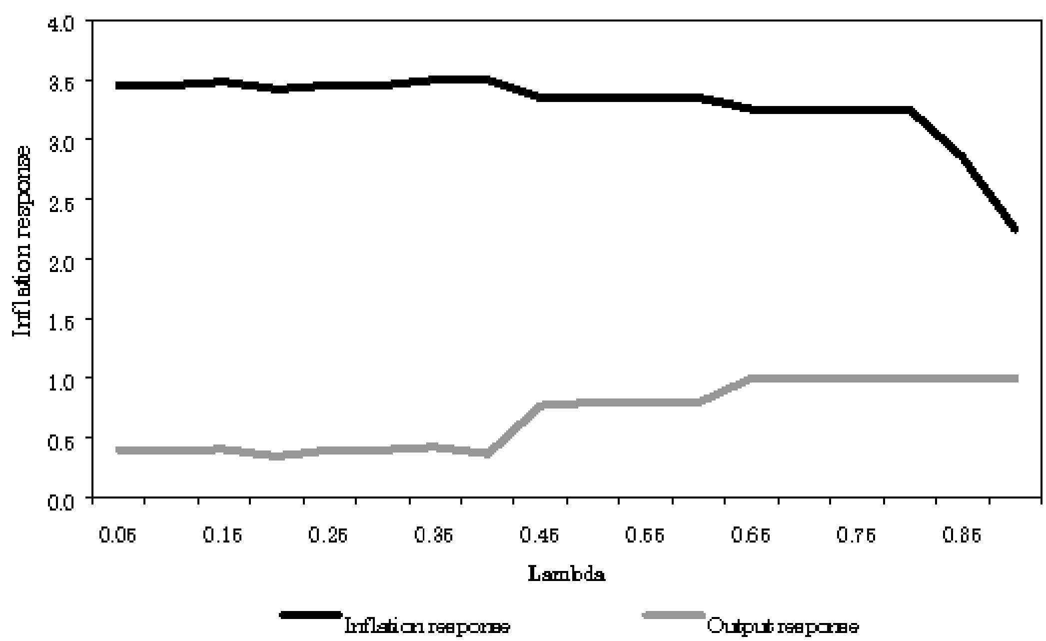
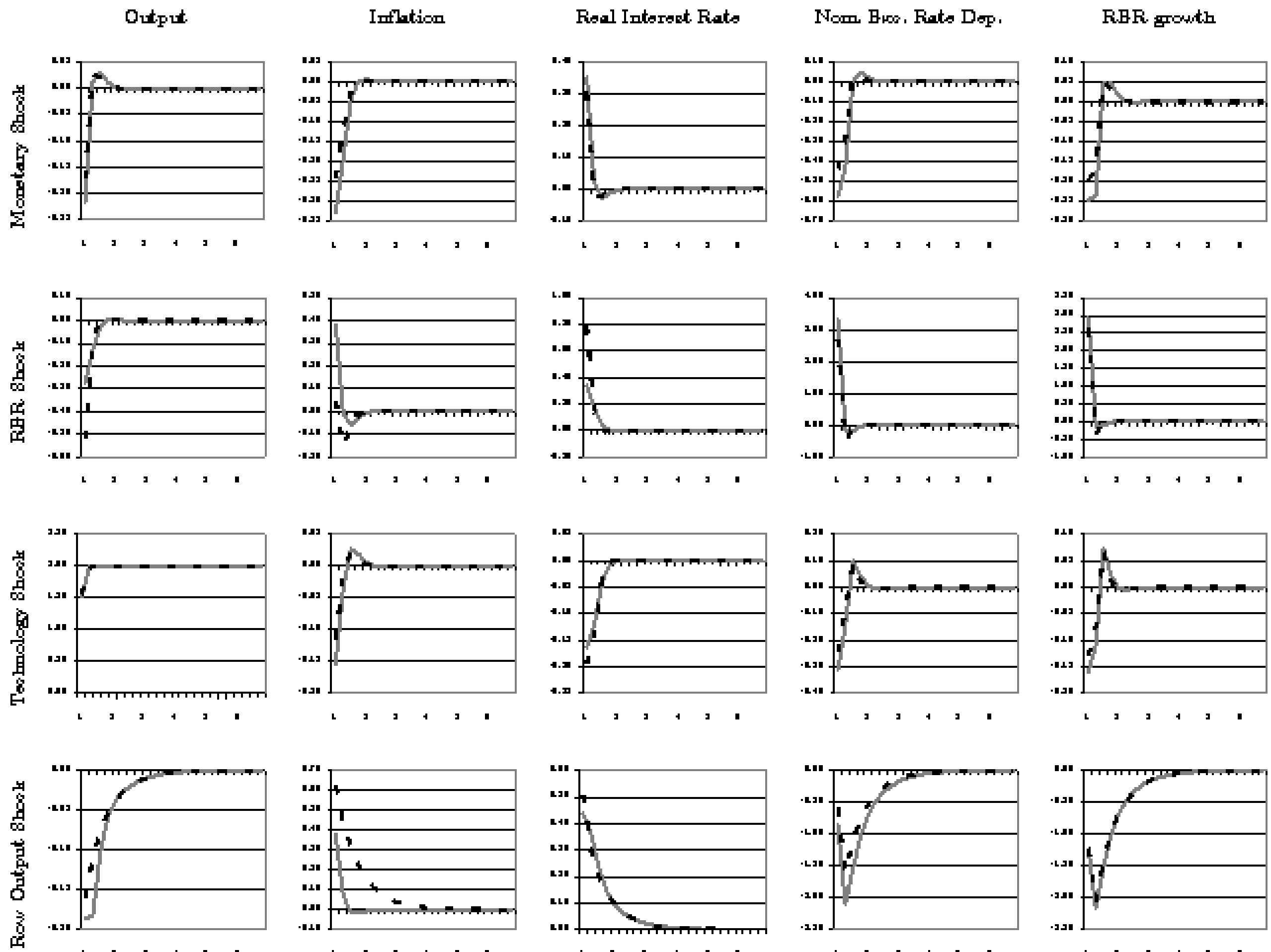


Figure 3. Impulse Response Function with and without reaction to exchange rate



Solid line $\psi_3 = 0$. Dashed line $\psi_3 > 0$

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