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Evidence on Inflation Targeting
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How Does Monetary Policy Change?
Evidence on Inflation Targeting Countries

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Abstract

We examine the evolution of monetary policy rules in a group of inflation targeting countries (Australia, Canada, New Zealand, Sweden and the United Kingdom), applying a moment-based estimator in a time-varying parameter model with endogenous regressors. Using this novel flexible framework, our main findings are threefold. First, monetary policy rules change gradually, pointing to the importance of applying a time-varying estimation framework. Second, the interest rate smoothing parameter is much lower than typically reported by previous time-invariant estimates of policy rules. External factors matter for all countries, although the importance of the exchange rate diminishes after the adoption of inflation targeting. Third, the response of interest rates to inflation is particularly strong during periods when central bankers want to break a record of high inflation, such as in the UK or Australia at the beginning of the 1980s. Contrary to common wisdom, the response becomes less aggressive after the adoption of inflation targeting, suggesting a positive anchoring effect of this regime on inflation expectations. This result is supported by our finding that inflation persistence as well as the policy neutral rate typically decreased after the adoption of inflation targeting.

JEL Classification: E43, E52, E58.
Keywords: Taylor rule, inflation targeting, monetary policy, time-varying parameter model, endogenous regressors.

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1 Introduction

Taylor-type regressions have been applied extensively in order to describe monetary policy setting for many countries. The research on U.S. monetary policy usually assumes that monetary policy was subject to structural breaks when the Fed chairman changed. Clarida et al. (2000) claims that the U.S. inflation during the 1970s was unleashed because the Fed's interest rate response to the inflation upsurge was too weak, while the increase of such response in the 1980s was behind the inflation moderation. Although there is ongoing discussion on the sources of this Great Moderation (Benati and Surico, 2009), it is generally accepted that monetary policy setting evolves over time.

The evolution of monetary policy setting as well as exogenous changes in the economic system over time raises several issues for empirical analysis. In particular, the coefficients of monetary policy rules estimated over longer periods are structurally unstable. The solution used in the literature is typically sub-sample analysis (Clarida et al., 1998, 2000). Such an approach is based on the rather strong assumption that the timing of structural breaks is known, but also that policy setting does not evolve within each sub-period. Consequently, this gives impetus to applying an empirical framework that allows for regime changes or, in other words, time variance in the model parameters (Cogley and Sargent, 2001, 2005). Countries that have implemented the inflation targeting (IT) regime are especially suitable for such analysis because it is likely that the monetary policy stance with respect to inflation and other macroeconomic variables changed as a consequence of the implementation of IT. Moreover, there is ongoing debate of to what extent IT represents a rule-based policy. Bernanke et al. (1999) claim that IT is a framework or constrained discretion rather than a mechanical rule. Consequently, the monetary policy rule of an IT central bank is likely to be time varying.

Our study aims to investigate the evolution of monetary policy for countries that have long experience with the IT regime. In particular, we analyze the time-varying monetary policy rules for Australia, Canada, New Zealand, Sweden and the United Kingdom. As we are interested in monetary policy evolution over a relatively long period, we do not consider countries where IT has been in place for a relatively short time (Finland, Spain), or was introduced relatively recently (such as Armenia, the Czech Republic, Hungary, Korea, Norway and South Africa). We apply the recently developed time-varying parameter model with endogenous regressors (Kim and Nelson, 2006), as this technique allows us to evaluate changes in policy rules over time, and, unlike Markov-switching methods, does not impose sudden policy switches between different regimes.

On top of that, it also deals with endogeneity of policy rules. Unlike Kim and Nelson (2006) we do not rely on the Kalman filter, which is conventionally employed to estimate time-varying models, but employ the moment-based estimator proposed by Schlicht and Ludsteck (2006)[1] for its mathematical and descriptive transparency and minimal requirements as regards initial conditions. In addition, Kim and Nelson (2006) apply their estimator to evaluate changes in U.S. monetary policy, while we focus on inflation targeting economies.

Anticipating our results, we find that monetary policy changes gradually, pointing to the importance of applying a time-varying estimation framework (see also Koop et al., 2009, on evidence that monetary policy changes gradually rather than abruptly). When the issue of endogeneity in time-varying monetary policy rules is neglected, the parameters are estimated inconsistently, even though the resulting errors are economically not large. Second, the interest rate smoothing parameter is much lower than typically reported by previous time-invariant estimates of policy rules. This is in line with a recent critique by Rudebusch (2006), who emphasizes that the degree of smoothing is rather low. External factors matter for understanding the interest rate setting process for all countries, although the importance of the exchange rate diminishes after the adoption of inflation targeting. Third, the response of interest rates to inflation is particularly strong during periods when central bankers want to break a record of high inflation, such as in the UK at the beginning of the 1980s. Contrary to common wisdom, the response can become less aggressive after the adoption of inflation targeting, suggesting a positive anchoring effect of this regime on inflation expectations or a low inflation environment. This result is consistent with Kuttner and Posen (1999) and Sekine and Teranishi (2008), who show that inflation targeting can be associated with a smaller response of the interest rate to inflation developments if the previous inflation record was favorable.

The paper is organized as follows. Section 2 discusses the related literature. Section 3 describes our data and empirical methodology. Section 4 presents the results. Section 5 concludes. An appendix with a detailed description of the methodology and additional results follows.

[1] The description of this estimator is also available in Schlicht (2005), but we refer to the more recent working paper version, where this estimator is described in a great detail. Several important parts of this framework were introduced already in Schlicht (1981).

2 Related Literature

2.1 Monetary policy rules and inflation targeting

Although the theoretical literature on optimal monetary policy usually distinguishes between instrument rules (the Taylor rule) and targeting rules (the inflation-targeting based rule), the forward-looking specification of the Taylor rule, sometimes augmented with other variables, has commonly been used for the analysis of decision making of IT central banks. The existing studies feature great diversity of empirical frameworks, which makes the comparison of their results sometimes complicated. In the following we provide a selective survey of empirical studies aimed at the countries that we focus on.

The United Kingdom adopted IT in 1992 (currently a 2% target and a ±1% tolerance band) and the policy of the Bank of England (BoE) is subject to the most extensive empirical research. Clarida et al. (1998) analyzed the monetary policy setting of the BoE in the pre-IT period, concluding that it was consistent with the Taylor rule, yet additionally constrained by foreign (German) interest rate setting. Adam et al. (2005) find by means of sub-sample analysis that the introduction of IT did not represent a major change in monetary policy conduct, unlike the granting of instrument independence in 1997. Davradakis and Taylor (2006) point to significant asymmetry of British monetary policy during the IT period; in particular the BoE was concerned with inflation only when it significantly exceeded its target. Assenmacher-Wesche (2006) concludes by means of a Markov-switching model that no attention was paid to inflation until IT was adopted. Conversely, Kishor (2008) finds that the response to inflation had already increased, especially after Margaret Thatcher became prime minister (in 1979). Finally, Trecroci and Vassalli (2009) use a model with time-varying coefficients and conclude that policy had been getting gradually more inflation averse since the early 1980s.

New Zealand was the first country to adopt IT (in 1990). A particular feature besides the announcement of the inflation target (currently a band of 1–3%) is that the governor of the Reserve Bank (RBNZ) has an explicit agreement with the government. Huang et al. (2001) study the monetary policy rule over the first decade of IT. He finds that the policy of the RBNZ was clearly aimed at the inflation target and did not respond to output fluctuations explicitly. The response to inflation was symmetric and a backward-looking rule does as good a job as a forward-looking one at tracking the interest rate dynamics. Plantier and Scrimgeour (2002) allow for the possibility that the neutral real interest rate (implicitly assumed in the Taylor rule to be constant) changes in time. In this framework they find that the response to inflation increased after IT was implemented and the policy neutral interest rate tailed away. Ftiti (2008) additionally confirms that the RBNZ did not explicitly respond to exchange rate fluctuations and Karedekikli and Lees (2007) disregard asymmetries in the RBNZ policy rule.

The Reserve Bank of Australia (RBA) turned to IT in 1993 (with a target of 2–3%) after decades of exchange rate pegs (till 1984) and consecutive monetary targeting.[2] De Brouwer and Gilbert (2005) using sub-sample analysis confirm that the RBA's consideration of inflation was very low in the pre-IT period and a concern for output stabilization was clearly predominant. The response to inflation (both actual and expected) increased substantially after IT adoption but the RBA seemed to consider exchange rate and foreign interest rate developments as well. Leu and Sheen (2006) find a lot of discretionality in the RBA's policy (a low fit of the time-invariant rule) in the pre-IT period, a consistent response to inflation during IT, and signs of asymmetry in both periods. Karedekikli and Lees (2007) document that the policy asymmetry is related to the RBA's distaste for negative output gaps.

The Bank of Canada (BoC) introduced IT in 1991 in the form of a series of targets for reducing inflation to the midpoint of the range of 1–3% by the end of 1995 (since then the target has remained unchanged). Demers and Rodríguez (2002) find that the implementation of this framework was distinguished by a higher inflation response, but the increase in the response to real economic activity was even more significant. Shih and Giles (2009) model the duration analysis of BoC interest rate changes with respect to different macroeconomic variables. They find that annual core inflation and the monthly growth rate of real GDP drive the changes of the policy rate, while the unemployment rate and the exchange rate do not. On the contrary, Dong (2008) confirms that the BoC considers real exchange rate movements.

Sweden adopted IT in 1993 (a 2% target with a tolerance band of 1 percentage point) just after the krona had been allowed to float. The independence of Sveriges Riksbank (SR) was legally increased in 1999. Jansson and Vredin (2003) studied its policy rule, concluding that the inflation forecast (published by the Riksbank) is the only relevant variable driving interest rate changes. Kuttner (2004) additionally finds a role for the output gap, but in terms of its growth forecast (rather than its observed value). Berg et al. (2004) provide a rigorous analysis of the sources of deviations between the SR policy rate and the targets implied by diverse empirical rules. They

[2] In Australia, the adoption of inflation targeting was a gradual process. As from January 1990, the RBA increased the frequency of its communications via speeches and the style of the Bank's Bulletin started to correspond to the inflation reports as introduced in New Zealand. The exact inflation target was defined explicitly later, in April 1993. Greenville (1997) describes the policy changes in Australia in great detail.

claim that higher inflation forecasts at the early stages of the IT regime (due to a lack of credibility) generate a higher implied target from the forward-looking rule and therefore induce spurious indications of policy shocks. Their qualitative analysis of SR documents clarifies the rationale behind actual policy shocks, such as more gradualism (stronger inertia) in periods of macroeconomic uncertainty.

Finally, there are a few multi-country studies. Meirelles Aurelio (2005) analyzes the time-invariant rules of the same countries as us, finding significant dependence of the results on real-time versus historical measures of variables. Lubik and Schorfheide (2007) estimate by Bayesian methods an open economy structural model of four IT countries (AUS, CAN, NZ, UK) with the aim of seeing whether IT central banks respond to exchange rate movements. They confirm this claim for the BoE and BoC. Dong (2008) enriches their setting by incorporating some more realistic assumptions (exchange rate endogeneity, incomplete exchange-rate pass-through), finding additionally a response to the exchange rate for the RBA.

2.2 Time variance in monetary policy rules

The original empirical research on monetary policy rules used a linear specification with time-invariant coefficients. Instrument variable estimators such as the GMM gained popularity in this context, because they are able to deal with the issue of endogeneity that arises in the forward-looking specification (Clarida et al., 1998).[3] While a time-invariant policy rule may be a reasonable approximation when the analyzed period is short, structural stability usually fails over longer periods.

The simplest empirical strategy for taking time variance into account is to use sub-sample analysis (Taylor, 1999; Clarida et al., 2000). The drawback of this approach is its rather subjective assumptions about points of structural change and structural stability within each sub-period. An alternative is to apply an econometric model that allows time variance for the coefficients. There are various methods dealing with time variance in the context of estimated monetary policy rules.

The most common option is the Markov-switching VAR method, originally used for business cycle analysis. Valente (2003) employs such a model with switches in the constant term

[3] One exception is when a researcher uses real-time central bank forecasts for Taylor-type rule estimation, i.e. the data available to the central bank before the monetary policy meeting. In such case, the endogeneity problem will not arise and least squares estimation may perform well (Orphanides, 2001). However, as we will discuss in more detail below, the use of real-time data may not solve the issue of endogeneity completely.

representing the evolution of the inflation target (the inflation target together with the real equilibrium interest rate makes the constant term in a simple Taylor rule). Assenmacher-Wesche (2006) uses the Markov-switching model with shifts both in the coefficients and in the residual variances. Such separation between the evolution of policy preferences (coefficients) and exogenous changes in the economic system (residuals) is important for the continuing discussion on the sources of the Great Moderation (Benati and Surico, 2008; Canova and Gambetti, 2008). Sims and Zha (2006) present a multivariate model with discrete breaks in both coefficients and disturbances. Unlike Assenmacher-Wesche they find that the variance of the shock rather than the time variance of the monetary policy rule coefficient has shaped macroeconomic developments in the U.S. in the last four decades.

The application of Markov-switching VAR techniques turns out to be complicated for IT countries, where the policy rules are usually characterized as forward-looking and some regressors become endogenous. The endogeneity bias can be avoided by means of a backward-looking specification (lagged explanatory variables), but this is very probably inappropriate for IT central banks, which are arguably forward-looking.[4] However, there is another distinct feature of the Markov-switching model that makes its use for the analysis of time variance in the monetary policy rule rather questionable. The model assumes sudden switches from one policy regime to another rather than a gradual evolution of monetary policy. Although at first sight one may consider the introduction of IT to be an abrupt change, there are some reasons to believe that a smooth monetary policy transition is a more appropriate description for IT countries (Koop et al., 2009). Firstly, the IT regime is typically based on predictability and transparency, which does not seem to be consistent with sudden switches. Secondly, it is likely that inflation played a role in interest rate setting even before the IT regime was introduced, because in many countries a major decrease of inflation rates occurred before IT was implemented. Thirdly, the coefficients of different variables (such as inflation, the output gap or the exchange rate) in the monetary policy rule may evolve independently rather than moving from one regime to another at the same time (see also Darvas, 2009). For instance, a central bank may assign more weight to the observed or expected inflation rate when it implements IT, but that does not mean that it immediately disregards information on real economic activity or foreign interest rates. Finally, there is relevant evidence, though mostly for the U.S., that monetary policy evolves rather smoothly over time (Boivin, 2006; Canova and Gambetti, 2008; Koop et al., 2009). Therefore, based on this research, a smooth transition seems to be a more appropriate description of reality. In a similar manner, it

[4] Psaradakis et al. (2006) proposed a solution to the endogeneity problem in the context of the Markov-switching model in the case of the term structure of interest rates.

is possible to estimate the policy rule using STAR-type models. Nevertheless, it should be noted that STAR-type models assume a specific type of smooth transition between regimes, which can be more restrictive than the flexible random walk specification that we employ in this paper. Therefore, we leave the empirical examination of Markov-switching as well as STAR-type models for further research.[5]

Besides simple recursive regression (e.g. Domenech et al., 2002), the Kalman filter has been employed in a few studies to estimate a coefficient vector that varies over time. Such a time-varying model is also suitable for reflection of possible asymmetry of the monetary policy rule (Dolado et al., 2004).[6] An example of such asymmetry is that the policy maker responds more strongly to the inflation rate when it is high than when it is low. Boivin (2006) uses such a time-varying model estimated via the Kalman filter for the U.S., Elkhoury (2006) does the same for Switzerland, and Trecroci and Vassalli (2010) do so for the U.S., the UK, Germany, France and Italy. However, none of these studies provides a specific econometric treatment to the endogeneity that arises in forward-looking specifications.

In this respect, Kim (2006) proposes a two-step procedure for dealing with the endogeneity problem in the time-varying parameter model. Kim and Nelson (2006) find with this methodology that U.S. monetary policy has evolved in a different manner than suggested by previous research. In particular, the Fed's interest in stabilizing real economic activity has significantly increased since the early 1990s.[7] Kishor (2008) applies the same technique for analysis of the time-varying monetary policy rules of Japan, Germany, the UK, France and Italy.

[5] We run a number of experiments on simulated data with the true coefficients containing large sudden shifts to see whether the estimated coefficients are gradually changing or shifting. The specification of the basic experiment was as follows: The intercept follows a slowly moving random walk (variance of innovations set to 0.3). For the independent variable we used expected inflation in the UK, with the beta coefficient set to 0.75 up to the 60th observation and 1.75 afterwards. Then we included a lag of the dependent variable with a constant coefficient equal to 0.5 and residuals with distribution N(0.15). The dependent variable was generated as the sum of these components. This example can be linked to a reaction function of a hypothetical central bank that smoothes the interest rate, that does not have the output gap in its reaction function, and that changed its aggressiveness abruptly in the middle of the sample. Then we estimated the model using the VC method and stored the value of the estimated change of the beta coefficient at the time of the switch in the data generating process. We repeated this small experiment 30 times for different sets of intercept and residuals and the estimated value of the switch was 0.98 on average, ranging from 0.9 to 1.02. The average sizes of the innovations in beta were below 0.09 in the remaining part of the samples. Clearly, in this simple setting an abrupt change in policy is detected by the model with respect to the size and timing of that change. Further experiments contained more variables in the reaction function and more switches going upwards and downwards as well. We found that, generally, sudden changes (larger than the average changes in the other time-varying coefficients) in the true coefficients resulted in switches in the estimated time-varying coefficients, too, and the varying coefficients did not imply gradual changes in these cases. Nevertheless, the ability of the time-varying parameter models to identify sudden structural breaks in parameters remains to be confirmed by careful Monte Carlo simulations, as pointed out by Sekine and Teranishi (2008).

[6] Granger (2008) shows that any non-linear model can be approximated by a time-varying parameter linear model.

[7] Kim et al. (2006) confirmed this finding with real-time data and additionally detected a significant decrease in the response to expected inflation during the 1990s.

He detects a time-varying response not only with respect to the inflation rate and the output gap, but also with respect to the foreign interest rate. The relevance of endogeneity correction can be demonstrated by the difference between Kishor's results and those of Trecroci and Vassalli (2010), who both study the same sample of countries.[8] The time-varying parameter model with specific treatment of endogeneity can be relevant even when real-time data are used instead of ex-post data (Orphanides, 2001). When the real-time forecast is not derived under the assumption that nominal interest rates will remain constant within the forecast horizon, endogeneity may still be present in the model (see Boivin, 2006). Moreover, this estimation procedure is also viable for reflecting measurement error and heteroscedasticity in the model (Kim et al., 2006). However, the Kalman filter applied to a state-space model may suffer one important drawback in small samples: it is rather sensitive to the initial values of parameters which are unknown. The moment-based estimator proposed by Schlicht (1981), Schlicht (2005) and Schlicht and Ludsteck (2006), which is employed in our paper and described below, allows this problem to be avoided. Moreover, it is flexible enough to incorporate the endogeneity correction proposed by Kim (2006).

3 Empirical Methodology

3.1 The empirical model

In line with Taylor (1993) most empirical studies' models assume that the central bank targets the nominal interest rate in line with the state of the economy (see Clarida et al., 1998, 2000). Such policy rule, which in the case of an IT central bank is arguably forward-looking, can be written as follows:

$$r_t^* = \bar{r} + \beta\left(E\left[\pi_{t+i} \left|\Omega_t\right.\right] - \pi_{t+i}^*\right) + \gamma E\left[y_{t+j} \left|\Omega_t\right.\right] \tag{1}$$

where r_t^* denotes the targeted interest rate, $\bar{r}$ is the policy neutral rate[9], π_{t+i} stands for the central bank forecast of the yearly inflation rate *i* periods ahead, and π_{t+i}^* is the central bank's inflation target.[10] y_{t+j} represents a measure of the output gap. $E[\ ]$ is the expectation operator and Ω_t is the information set available at the time *t* when interest rates are set. Eq. (1) links the

[8] Horváth (2009) employs the time-varying model with endogenous regressors for estimation of the neutral interest rate for the Czech Republic and confirms the importance of endogeneity bias correction terms.
[9] The policy neutral rate is typically defined as the sum of the real equilibrium rate and expected inflation.
[10] The definition of the inflation target varies slightly across IT countries. However, the target is usually mid-term rather than short-term. The target value can also vary over time. This variation has been especially pronounced in emerging countries that implemented IT as a gradual disinflation strategy. By contrast, for the countries studied here, the target value has not changed significantly over time.

policy instrument (nominal interest rates) to a constant term (the neutral rate that would prevail if expected inflation and output were at their targeted levels), the deviation of expected inflation from its target value and the output gap.

Nevertheless, Eq. (1) is often argued to be too restrictive to provide a reasonable description of actual interest rate setting. First, it does not account for interest rate smoothing by central banks. In line with Clarida et al. (1998) most studies assume that the central bank adjusts the interest rate sluggishly to the targeted value. This can be tracked by a simple partial-adjustment mechanism:

$$r_t = \rho r_{t-1} + (1-\rho) r_t^* \tag{2}$$

where $\rho \in [0,1]$ is the smoothing parameter. Although this is line with the common wisdom that central banks are averse to abrupt changes, most studies that estimate time-invariant models find unusually high policy inertia. For instance, using quarterly data ρ typically exceeds 0.8. Rudebusch (2006) points to an inconsistency between this finding and the practical impossibility to predict interest rate changes over a few quarters. Therefore, it is possible that the lagged dependent value takes over the impact of either autocorrelated shocks or omitted variables. The intensity of interest rate smoothing is logically reinforced in a linear time-invariant specification, as the response to some variables can be asymmetric and/or vary in time. Second, Eq. (1) assumes that the central bank aims only at the inflation rate and the output gap. However, many central banks that have implemented IT are in small open economies and may consider additional variables, in particular the exchange rate and the foreign interest rate, hence Eq. (1) can be extended for δx_t, where the coefficient δ measures the impact of additional variable x_t on interest rate setting. Therefore, in our empirical model we substitute Eq. (2) into Eq. (1), eliminating unobserved forecast variables and including additional variables, which results in Eq. (3):

$$r_t = (1-\rho)\left[\alpha + \beta\left(\pi_{t+i} - \pi_{t+i}^*\right) + \gamma y_t + \delta x_t\right] + \rho r_{t-1} + \varepsilon_t \tag{3}$$

Also, α is a constant term which in Eq. (1) coincides with the policy neutral rate $\overline{r}$. However, if the model is augmented by additional variables that are not in the form of deviations from the target value, the constant term need not keep this interpretation. We set i equal to 2.[11] Consequently, the disturbance term ε_t is a combination of forecast errors and is thus orthogonal to all information available at time t ($\mathbf{\Omega}_t$).

In line with our previous discussion, the interest rate rule described above will be estimated within a framework that allows time variance of the coefficients. Kim (2006) shows that the conventional time-varying parameter model (the Kalman filter applied to a state-space representation) delivers inconsistent estimates when the explanatory variables are correlated with the disturbance term. Endogeneity arises in forward-looking policy rules based on ex-post data, but it can appear even with real-time data, as discussed before. Kim (2006) proposes an estimator of the time-varying coefficient model with endogenous regressors. A few recent contributions use this framework for estimation of monetary policy rules (Kim and Nelson, 2006, Kim et al. 2006, Kishor, 2008).[12] Following Kim (2006) we can rewrite Eq. (3) as follows:

$$r_t = (1-\rho_t)\left[\alpha_t + \beta_t(\pi_{t+i}) + \gamma_t y_t + \delta_t x_t\right] + \rho_t r_{t-1} + \varepsilon_t \tag{4}$$

$$\alpha_t = \alpha_{t-1} + \vartheta_{1,t},\ \vartheta_{1,t} \sim i.i.d.N\left(0,\sigma^2_{\vartheta_1}\right) \tag{5}$$

$$\beta_t = \beta_{t-1} + \vartheta_{2,t},\ \vartheta_{2,t} \sim i.i.d.N\left(0,\sigma^2_{\vartheta_2}\right) \tag{6}$$

$$\gamma_t = \gamma_{t-1} + \vartheta_{3,t},\ \vartheta_{3,t} \sim i.i.d.N\left(0,\sigma^2_{\vartheta_3}\right) \tag{7}$$

$$\delta_t = \delta_{t-1} + \vartheta_{4,t},\ \vartheta_{4,t} \sim i.i.d.N\left(0,\sigma^2_{\vartheta_4}\right) \tag{8}$$

$$\rho_t = \rho_{t-1} + \vartheta_{5,t},\ \vartheta_{5,t} \sim i.i.d.N\left(0,\sigma^2_{\vartheta_5}\right) \tag{9}$$

$$\pi_{t+i} = Z'_{t-j}\xi + \sigma_\varphi \varphi_t,\ \varphi_t \sim i.i.d.N(0,1) \tag{10}$$

$$y_t = Z'_{t-j}\psi + \sigma_v v_t,\ v_t \sim i.i.d.N(0,1) \tag{11}$$

The measurement equation (4) of the state-space representation is the monetary policy rule. The transition equations (5)–(9) describe the time-varying coefficients as a random walk process without drift. Eqs. (10) and (11) track the relationship between the endogenous regressors (π_{t+i}

[11] Although the targeting horizon of central banks is usually longer (4–8 quarters), we prefer to proxy inflation expectations by inflation in *t+2* for the following reasons. First, the endogeneity correction requires a strong correlation between the endogenous regressor and its instruments. Second, the prediction error logically increases at longer horizons. Third, the countries we analyze did not apply inflation targeting during the whole estimation period. Consequently, it is preferable owing to data limitations to keep only two inflation leads rather than four or six.

[12] Note, however, that two of these contributions are, to our knowledge, unpublished as yet.

and y_t) and their instruments, Z_t. The list of instruments, Z_{t-j}, is as follows: π_{t-1}, π_{t-4}, y_{t-1}, y_{t-2}, r_{t-1} and r_t^+ (foreign interest rate). Unlike Kim (2006), we assume that the parameters in Eqs. (10) and (11) are time-invariant. Next, the correlation between the standardized residuals φ_t and v_t and ε_t is $\kappa_{\varphi,\varepsilon}$ and $\kappa_{v,\varepsilon}$, respectively (note that σ_φ and σ_v are standard errors of φ_t and v_t, respectively). The consistent estimates of the coefficients in Eq. (4) are obtained in two steps. In the first step, we estimate equations (10) and (11) and save the standardized residuals φ_t and v_t. In the second step, we estimate Eq. (12) below along with Eqs. (5)–(9). Note that (12) now includes bias correction terms[13], i.e. the (standardized) residuals from Eqs. (10) and (11), to address the aforementioned endogeneity of the regressors. Consequently, the estimated parameters in Eq. (12) are consistent, as ζ_t is uncorrelated with the regressors.

$$r_t = (1-\rho_t)\left[\alpha_t + \beta_t(\pi_{t+i}) + \gamma_t y_t + \delta_t x_t\right] + \rho_t r_{t-1} + \kappa_{v,\varepsilon}\sigma_{\varepsilon,t} v_t + \kappa_{\varphi,\varepsilon}\sigma_{\varepsilon,t}\varphi_t + \zeta_t$$

$$, \quad \zeta_t \sim N\left(0, (1-\kappa_{v,\varepsilon}^2 - \kappa_{\varphi,\varepsilon}^2)\sigma_{\varepsilon,t}^2\right) \tag{12}$$

The standard framework for second-step estimation is the maximum likelihood estimator via the Kalman filter (Kim, 2006). However, there are several difficulties with the estimation of the Kalman filter (and Kalman smoother) in applied work. First, if the variables are nonstationary, the results often depend on the proper choice of the initial values, but those values are not known in advance.[14] The problem with the initial conditions is larger if one-sided estimates are used, as illustrated in Leigh (2008) on estimates of the time-varying natural rate of interest in the U.S. Applying the Kalman smoother alleviates the issue and the differences in the estimates at the beginning of the sample decrease sharply. Second, the log likelihood function is highly nonlinear and in some cases the optimization algorithm fails to minimize the negative of the log likelihood for several reasons (either it can fail to calculate the Hessian matrix throughout the iteration process or, when the likelihood function is approximated to facilitate the computations, the covariance matrix of the observation vector can become singular for the provided starting values).

[13] Obviously, if the correction terms are statistically significant, it shows that endogeneity matters. Similarly to Kim and Nelson (2006) and Horváth (2009), we find that these terms are typically significant: in our sample the endogeneity correction for inflation is significant for all countries except the UK at least at the 10% level, and for the GDP gap it is significant for Canada (see table A.1 in the Appendix).

[14] Although there are a number of formal procedures for initialization of the Kalman filter in such cases (for example see Koopman et al., 1999), fundamental uncertainty about their values remains.

In this paper, we adopt the "varying coefficients" (VC) method (Schlicht and Ludsteck, 2006). The VC method generalizes the standard ordinary least squares approach. In fact, instead of minimizing the sum of the squares of the residuals $\sum_{t=1}^{T}\zeta_t^2$, it uses minimization of the weighted sum of the squares:

$$\sum_{t=1}^{T}\zeta_t^2+\theta_1\sum_{t=1}^{T}\vartheta_{1,t}^2+\theta_2\sum_{t=1}^{T}\vartheta_{2,t}^2+\ldots+\theta_n\sum_{t=1}^{T}\vartheta_{n,t}^2 \tag{13}$$

where the weights θ_i are the inverse variance ratios of the regression residuals ζ_t and the shocks in the time-varying coefficients ϑ_i, that is, $\theta_i=\sigma^2/\sigma_i^2$. Hence it balances the fit of the model and the parameter stability.[15] Additionally, the time averages of the regression coefficients estimated by such weighted least squares estimator are identical to the GLS estimates of the corresponding regression with fixed coefficients, that is, $\frac{1}{T}\sum_{t=1}^{T}\hat{a}_t=\hat{a}_{GLS}$.

The VC method has a number of advantages. First, it does not require initial conditions even for non-stationary variables prior to the estimation procedure. Instead, both the variance ratios and the coefficients are estimated simultaneously. Second, the property of the estimator that the time averages of the estimated time-varying coefficients are equal to their time-invariant counterparts permits easier interpretation of the results by comparison with time-invariant results. The features of the VC method make it feasible for our analysis: we deal with a time-varying model where the coefficients are assumed to follow a random walk, there is no *a priori* information about the initial values and the time series are rather short.[16]

Furthermore, Schlicht and Ludsteck (2006) compare the results from the VC method and from the Kalman filter, showing that both estimators give very similar results given the assumption

[15] It should be noted that throughout our computations we did not have to solve problems with convergence of the moment estimator, as it was almost always able to find equilibrium. Computational details of the VC method are described in the Appendix. Originally, Schlicht and Ludsteck (2006) start with a derivation of the maximum likelihood estimator of parameters *a* based on the idea of orthogonal parameterization, which is described in the Appendix. Then they prove that the weighted least squares estimator is identical to the maximum likelihood estimator and also that the likelihood estimator is identical to the moment estimator for very large samples.

[16] The number of observations differs across the countries, ranging from 89 to 144. In the case of Kalman filter we can utilize the whole sample if we opt for initial conditions equal to the full sample OLS estimated values (recommended for stationary systems). Another approach derives the initial conditions related directly to the beginning of the sample from the first subset of available observations and the Kalman filter is performed on the latter part of the sample. Kim and Nelson (2006) adopted this approach and used the first 40 observations for the initialization. The estimation of the second step is carried out by Schlicht's VC package, which uses the moment estimator.

that the Kalman filter is initialized with the correct initial conditions. Yet in this case, the VC estimator has a slightly lower mean squared error and this difference is more pronounced for small samples.[17]

We assume that the variance of the disturbance term in Eq. (12) is not time-varying. Nevertheless, there is an ongoing discussion about to what extent changes in the macroeconomic environment are driven by changes in the variance of the disturbance term (i.e. exogenous changes in the economic system) vis-à-vis the variance in the coefficients of the monetary policy rule (see, for example, Benati and Surico, 2008, Canova and Gambetti, 2008, or Sims and Zha, 2006).

One can also think about Eqs. (4), (10) and (11) in terms of the New Keynesian model, with Eqs. (10) and (11) representing the Phillips and IS curves. It should be noted that our framework is in general less restrictive and imposes less structure than the full-blown New Keynesian model.

We expect β_t to be positive, as the central bank is likely to react to an increase in expected inflation by increasing its policy rate. In particular, β_t should be greater than one in the long-run solution of Eq. (4) if monetary policy is stabilizing. The development of β_t over time may be driven by a number of factors, such as changes in monetary policy regime or institutional constraints (Adam et al., 2005). The effect of the adoption of inflation targeting on β_t is ambiguous. As put forward by Kuttner and Posen (1999), β_t can both increase and decrease. They show that under a conservative central bank the response of short-term interest rates is greater than under discretion or the optimal state-contingent rule (inflation targeting),[18] while the strength of the response under inflation targeting as compared to discretion depends on the credibility of the regime. Credible monetary policy does not have to react so strongly to inflation surprises, as inflation expectations are likely to remain anchored. Sekine and Teranishi (2008) provide a new Keynesian model that reaches to the same conclusions. Siklos and Weymark

[17] For comparison, we estimated equation (12) using the conventional Kalman filter in the GROCER software using the function *tvp* (Dubois-Michaux, 2009). We parameterized the model with initial conditions taken from the OLS estimates of the parameters on the full sample and the initial forecast error covariance matrix set to 0. The matrix of the residuals of the time-varying coefficients is assumed to be diagonal as in the VC method. The results were very similar to those obtained from the VC method, with the estimated variances being the same in both methods. The only country where the estimated variance was different, was Sweden, with a lower variance in smoothing parameter ϱ and higher a variance in β. Still, the results were consistent with ours. These results are available upon request.

[18] See King (1997) on how inflation targeting allows one to come close to the optimal state-contingent rule.

(2009) estimate that inflation targeting in Australia, Canada and New Zealand reduced the magnitude of the interest rate changes to needed to maintain a low inflation environment.

Similarly, ρ , a measure of interest rate smoothing, is expected to be positive with values between zero and one. Many time-invariant estimates of monetary policy rules find the value of this parameter to be about 0.7–0.9, implying a substantial degree of interest rate smoothing. Rudebusch (2006) claims that such figures are clearly overestimated in the face of very low interest rate forecastability in the term structure of interest rates. On the contrary, the time-varying model in principle enables some variables to affect interest rate setting in one period but not in another, and is less prone to autocorrelated shocks.

Next, the effect of the output gap, γ_t , on interest rates is expected to be positive or insignificant. In the first case, the central bank may have an explicit concern for real activity or understand the output gap as a useful predictor of future inflation. In the latter case, the insignificant coefficient may suggest that the central bank is primarily focused on inflation and does not consider the output gap to be important in delivering low inflation.

There is a debate in literature about whether other variables should be included in the monetary policy rule. This is especially appealing for small open economies, which may be concerned with exchange rate fluctuations as well as the evolution of foreign interest rates. Taylor (2001) puts forward that even if the exchange rate or foreign interest rates are not explicitly included in the policy rule, they still remain present implicitly, as the exchange rate influences the inflation forecast, to which the inflation-targeting central bank is likely to react. It is also worth emphasizing that significance of the exchange rate or foreign interest rates does not necessarily mean that the central bank targets some particular values of these variables, but rather that the bank considers foreign developments to be important for its inflation forecast. On the other hand, empirical studies often favor the inclusion of these variables in the estimated policy rule. Having these considerations in mind, we decided to include the exchange rate and foreign interest rates, too, in order to assess whether these two variables carry any additional information for understanding the interest rate setting process in our sample countries.

3.2 The dataset

We use quarterly data. The sample period varies from country to country owing to data availability (the UK 1975:1Q–2007:4Q, Australia 1975:1Q–2007:4Q, Canada 1975:1Q–2007:4Q, New Zealand 1985:1Q–2007:3Q, Sweden 1982:2Q–2007:3Q), but on average the time coverage is about three decades.

Following Clarida, Galí and Gertler (1998), the dependent variable is the short-term interest rate, which is typically closely linked to the monetary policy rate. The reason for choosing the short-term interest rate rather than the monetary policy rate is the fact that the monetary policy rate and the change therein are censored (Podpiera, 2008). Therefore, the dependent variables capturing the policy rate are the discount rate (3-month Treasury bills) for the UK, the interbank 3-month interest rate for Australia, the 3-month Treasury bills rate for Canada, the overnight interbank 90-day interest rate for NZ and the interbank 3-month interest rate for Sweden. We choose the interest rate so as to be closely linked to monetary policy, but also to be available for a sufficiently long period. The foreign interest rate is the German 3-month Euribor for the UK and Sweden and the U.S. 3-month interbank interest rate for Australia, NZ, and Canada.

The inflation is measured as the year-on-year change in the CPI, except for the UK, where we use the RPIX (the retail price index excluding mortgage interest payments), and the NZ, where we use the CPIX (the CPI without interest payments).[19]

The output gap is taken as reported in the OECD Economic Outlook (the production function method based on the NAWRU – the non-accelerating wages rate of unemployment), except for NZ, where this series is short and where we use the output gap derived from the Hodrick-Prescott filter applied to the GDP series (constant prices, seasonally adjusted).[20] The exchange rate is measured by the chain-linked nominal effective exchange rate (NEER), except for Canada, where we use the bilateral USD/CAD exchange rate. For the regressions we use the deviation of the index from the HP trend (first differences were used for a robustness check) – see also Lubik and Schorfheide (2007).

[19] We use year-on-year data, as the inflation target is also defined on a year-on-year basis.
[20] There is no agreement on what is the best method for extraction of the unobserved output gap (Billmeier, 2009). We prefer to use the output gaps obtained by the OECD by means of the production function approach because they are based on a substantially richer information set than simple statistical detrending. Somewhat surprisingly, the OECD output gap and the simple HP gap evolve very closely.

4 Results

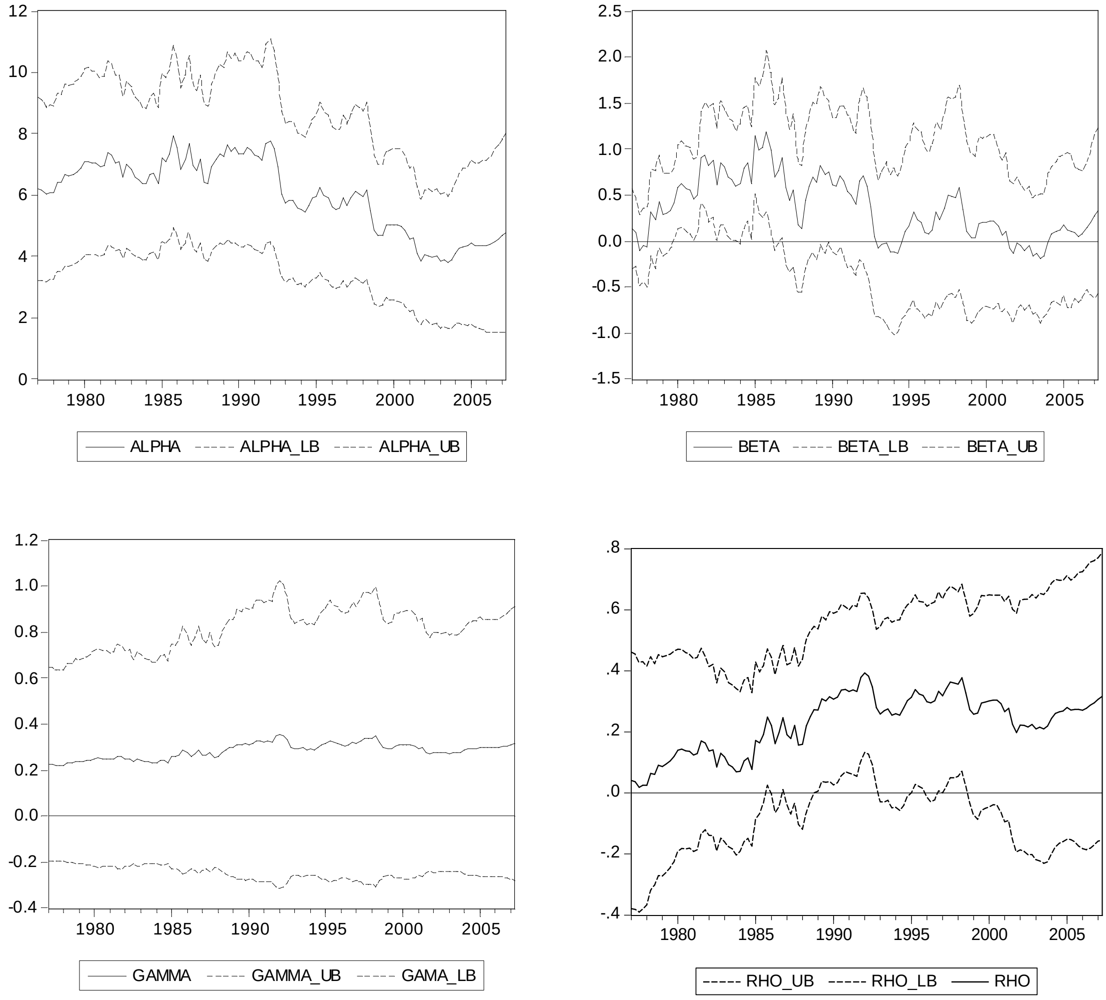
This section presents the country-specific estimates of the time-varying monetary policy rules in sub-sections 4.1–4.5. Sub-section 4.6 contains the estimates of time-varying inflation persistence and sub-section 4.7 provides a summary of the main policy-relevant findings.

4.1 United Kingdom

Our results show that the BoE significantly increased its response to inflation from the late 1970s till the mid-1980s. This overlaps with the Thatcher government and its major priority of inflation control. The overall decline of the response since 1985 can be related to the dismissal of the medium-term financial strategy (adopted in 1979). We find that the response of interest rates to inflation was gradually decreasing during the 1990s in spite of the introduction of IT. Although this finding may seem at first sight counterintuitive, it is important to keep in mind that, unlike in some emerging countries, IT was not implemented in the UK as a strong anti-inflation strategy. Inflation had already been contained in the 1980s and the very benign inflation environment was also supported by declining prices of raw materials on world markets. This corroborates with the findings of Kuttner and Posen (1999) and Sekine and Teranishi (2008), who show that inflation targeting can be associated with less aggressive monetary policy.

The effect of the output gap is estimated as positive (although the confidence intervals are rather large, probably reflecting the fact that the gap is an unobserved variable and calculated ex post) and does not vary substantially over time. The interest rate smoothing parameter is found to have values between 0.1 and 0.3, which is much lower than typically reported by time-invariant estimates of monetary policy rules (Clarida et al., 1998, 2000). Our estimates seem to be reasonable in the face of the recent critique by Rudebusch (2006). Finally, in this basic model the intercept can be interpreted as the policy neutral (nominal) interest rate. We can see that it has steadily declined over time, which is consistent with the low inflation environment that prevailed in the UK in the 1990s.

Figure 1 – Time-varying response coefficients in baseline (closed economy) policy rule, UK



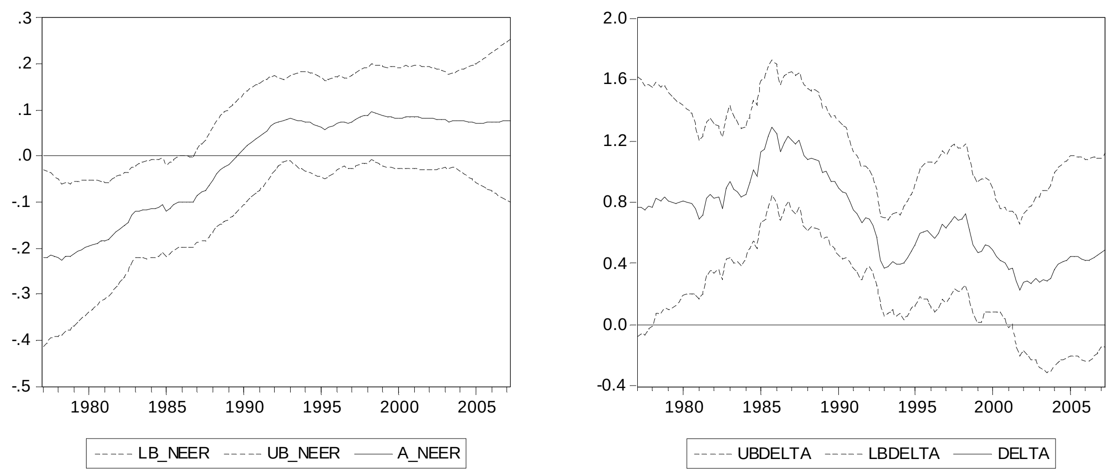
Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The upper-left graph depicts the evolution of the neutral rate. The upper-right graph depicts the evolution of the response of interest rates to inflation. The lower-left graph depicts the evolution of the response of interest rates to the output gap. The lower-right graph depicts the evolution of the interest rate smoothing parameter.

The results of our augmented model show that the monetary policy of the BoE was influenced by external factors, although their importance was greater in the 1980s than recently.[21] In particular, we find evidence that the BoE decreased its policy rate as the nominal effective exchange rate (NEER) strengthened during the 1980s even before the pound officially joined the ERM (1990). Yet, once the UK abandoned the ERM and introduced IT, the BoE no longer seemed to react explicitly to the exchange rate. Obviously, it considered the exchange rate indirectly, as exchange rate fluctuations influence the inflation forecast (for more on this see Taylor, 2001). The same reasoning applies to the response to the foreign interest rate (Euribor).

[21] In what follows, we present the evolution of the coefficients for the response to the exchange rate and foreign interest rates; the other coefficients remain largely unchanged and are not reported for the sake of brevity.

It was particularly strong during the 1980s and subsequently its importance somewhat declined. Our results show little support for the hypothesis that the monetary policy of the BoE follows that of the ECB, as the estimated response of the coefficient declined and the confidence intervals widened after the launch of the euro.

Figure 2 – Time-varying response coefficients in augmented (open economy) policy rule, UK



Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The left-hand graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (the deviation from the HP trend). The right-hand graph depicts the evolution of the response of the interest rate to the foreign interest rate.

There are two studies directly comparable with this paper. Kishor (2008) obtains results similar to ours in spite of using monthly data known to have slightly different dynamics. He finds that the anti-inflation stance peaked in the mid-1980s and tended to decline from then onwards in spite of the adoption of IT. Similarly, his finding that the response to the foreign interest rate significantly declined after the ERM crisis is complementary to our result that the BoE gave much less consideration to the evolution of the exchange rate (the NEER gap).

Trecroci and Vassalli (2010), who, unlike Kishor (2008) and this paper, do not correct for endogeneity in the time-varying model, come to the opposite conclusion that the BoE's response to inflation increased over time. Yet, some counterintuitive results of their study point to the possibility of endogeneity bias. First, the interest rate smoothing parameter takes on significantly negative values from 1980 till 1995. This would imply not only that policy was not inertial, but also that there was actually a negative correlation between the present and past interest rate, which is inconsistent even in the face of a simple visual inspection of the interest rate series.

When we estimate our model without the endogeneity-correcting coefficients we obtain a similar result (see Appendix, Figure A.1).[22] Second, their coefficient for the foreign (German) interest rate peaks in 1990 and is *de facto* invariant since then, which the authors interpret as implicit exchange-rate targeting. This finding is doubtful given the pound's exit from the ERM and the implementation of IT from 1992 onwards. In fact, British and German short-term rates, which were almost at par in 1992, diverged and the interbank interest rate in the UK exceeded the German one by almost 4% on the eve of euro adoption.

4.2 New Zealand

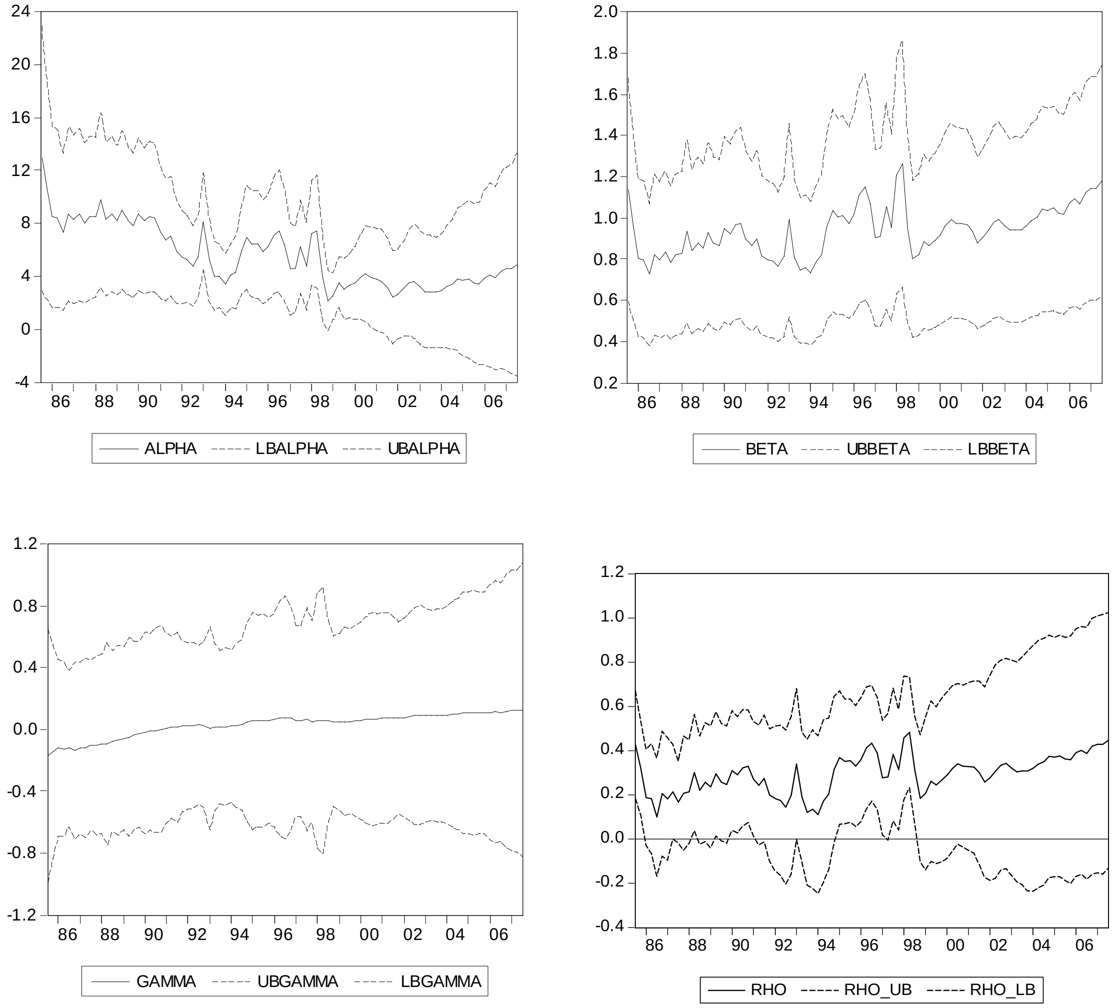
New Zealand was the first country in the world to introduce inflation targeting, doing so by means of the Federal Bank Act signed in March 1990.[23] Our results indicate that the response of the RBNZ to expected inflation was very close to unity throughout the sample period (1985–2007). In fact, it is clearly visible that the interest rate and inflation series move together very closely. However, in Figure 3 we can also see that the official introduction of IT does not seem to have engendered a significant change in interest rate setting (if anything there is very slight decrease of the response coefficient on inflation). Unlike in the UK, the response coefficient does not decrease substantially. This may be related to the fact that at the time IT was introduced in New Zealand the inflation rate was not far from double-digit values. Therefore, this policy was implemented in a different context than, say, in the UK, where single-digit inflation had already been achieved during the 1980s.[24] This result, together with the estimated insignificant response to the output gap, is consistent with the findings of time-invariant studies (Huang et al., 2001; Plantier and Scrimgeour, 2002) that the RBNZ applied a rather strict version of inflation targeting. Finally, we find that the interest rate smoothing parameter is again rather modest.

[22] The bias correction terms are significant when we estimate the model addressing endogeneity, even though the economic significance is in general not large – see the Appendix.

[23] Huang et al. (2001) argue that this policy was in effect since the end of 1988, when the RBNZ abandoned both monetary and exchange rate targeting. They also point to a specific feature of RBNZ monetary policy that could be referred to as "Open Mouth Operations". Between 1989 and 1999 the RBNZ specified a 90-day bank bill rate consistent with price stability and threatened to use quantitative controls to achieve the desired market rate if it deviated from the target. Therefore, the RBNZ did not control this interest rate permanently and directly.

[24] At the end of 1986 New Zealand introduced VAT, which had a direct impact on the inflation rate in the following two quarters. Consequently, we include a time dummy in Q1 and Q2 1987, whose coefficient is estimated as positive and significant.

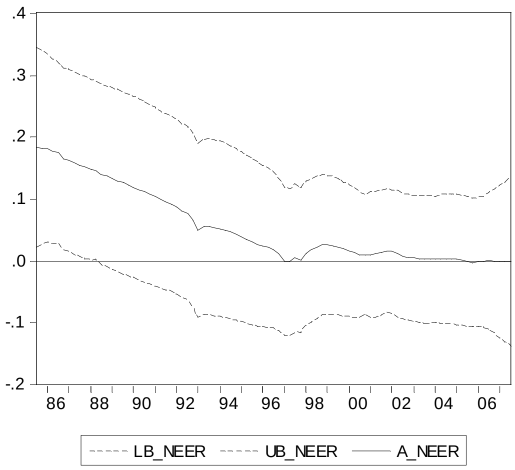
Figure 3 – Time-varying response coefficients in baseline (closed economy) policy rule, New Zealand

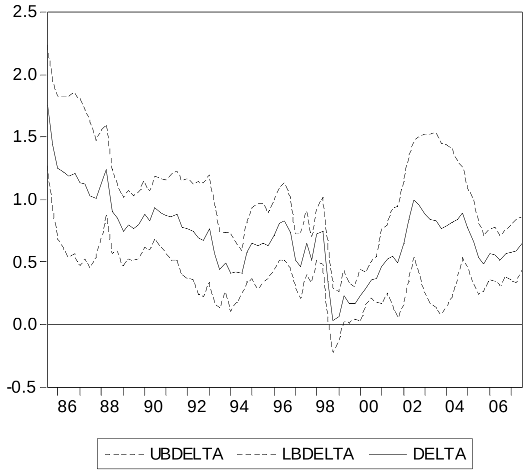


Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The upper-left graph depicts the evolution of the neutral rate. The upper-right graph depicts the evolution of the response of interest rates to inflation. The lower-left graph depicts the evolution of the response of interest rates to the output gap. The lower-right graph depicts the evolution of the interest rate smoothing parameter.

When we estimate the augmented model for New Zealand, the evidence for the exchange rate is not conclusive. We find a positive response to the NEER, which is rather counterintuitive in terms of the Taylor rule. However, the coefficient is significant only before the introduction of IT and its positive sign is probably related to currency appreciation following the interest rate increase. In that period the RBNZ aimed to keep the exchange rate within a predefined range and the interest rate was probably set so as to influence the exchange rate ex ante rather than ex post. Consistent with this finding, Ftiti (2008) in a time-invariant model rejects the hypothesis that the RBNZ responded to the exchange rate. On the other hand, we find some evidence in favor of consideration of the foreign interest rate, although its response coefficient generally decreased after the launch of IT.

Figure 4 – Time-varying response coefficients in augmented (open economy) policy rule, New Zealand





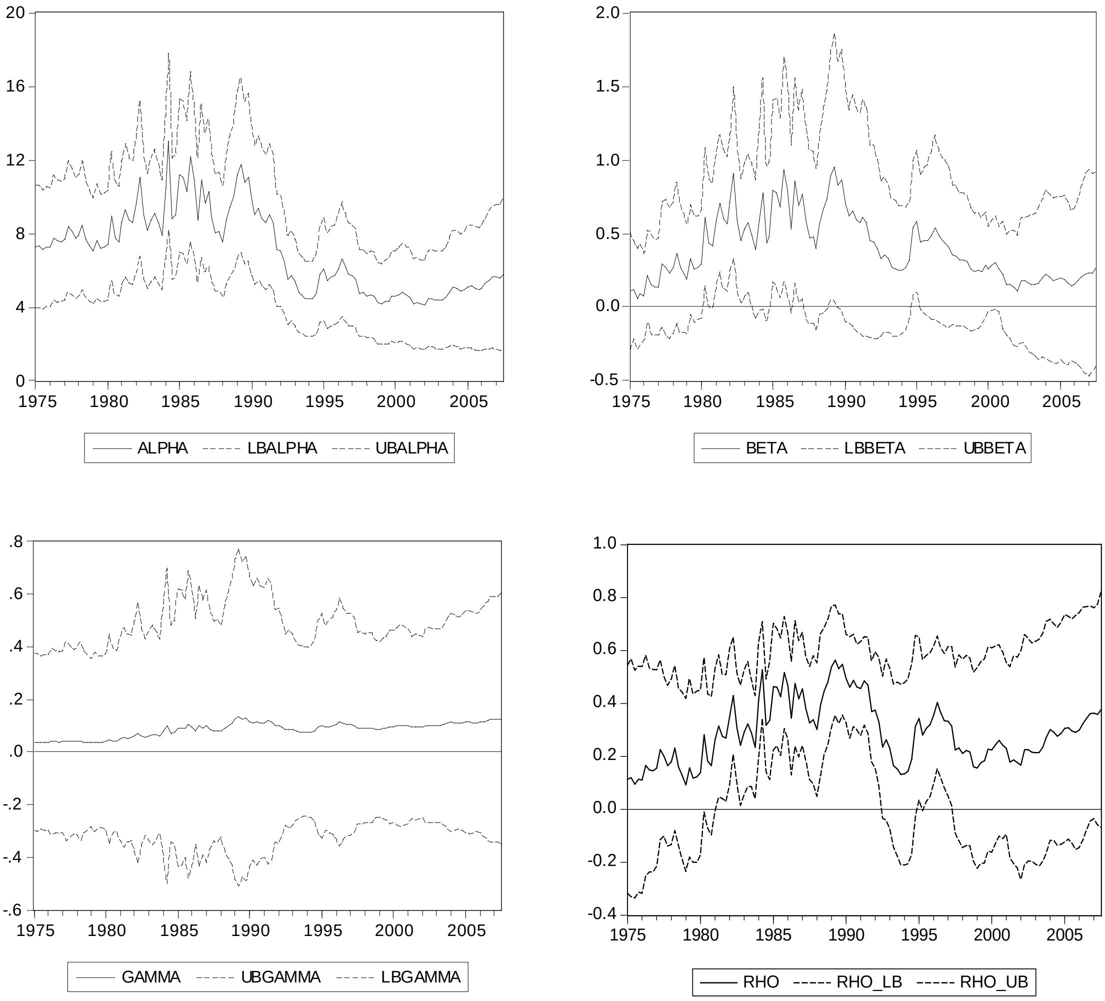
Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The left-hand graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (the deviation from the HP trend). The right-hand graph depicts the evolution of the response of the interest rate to the foreign interest rate.

4.3 Australia

Our results for Australia are available in Figures 5 and 6. The response of the interest rate to inflation is strongest in the 1980s, which is very similar to the UK experience. This period was characterized by inflation rates of around 10% and central bankers had to be quite aggressive in interest rate setting in order to break the record of high inflation deeply ingrained in public expectations. Neither monetary targeting (employed until 1984) nor the checklist approach (1985–1990) seemed to be successful in this regard. The fluctuation of the inflation response coefficient points to the discretionary nature of policy decisions (making this finding consistent with Leu and Sheen, 2006). The response coefficient peaks in 1990 on the eve of IT but declines after the adoption of this regime. It is again arguable whether it was the credibility of this regime that anchored inflation expectations and allowed the RBA to behave less aggressively. The original inflation decline may also have been related to the world recession in the early 1990s. Our results dispute the finding of De Brouwer and Gordon (2005) that the inflation response of the RBA increased as a result of the launch of inflation targeting.

As for other countries, the neutral rate declines in the 1990s, reflecting the global low inflation environment. The output gap is not found to be significant and the estimated interest rate smoothing is again rather low.

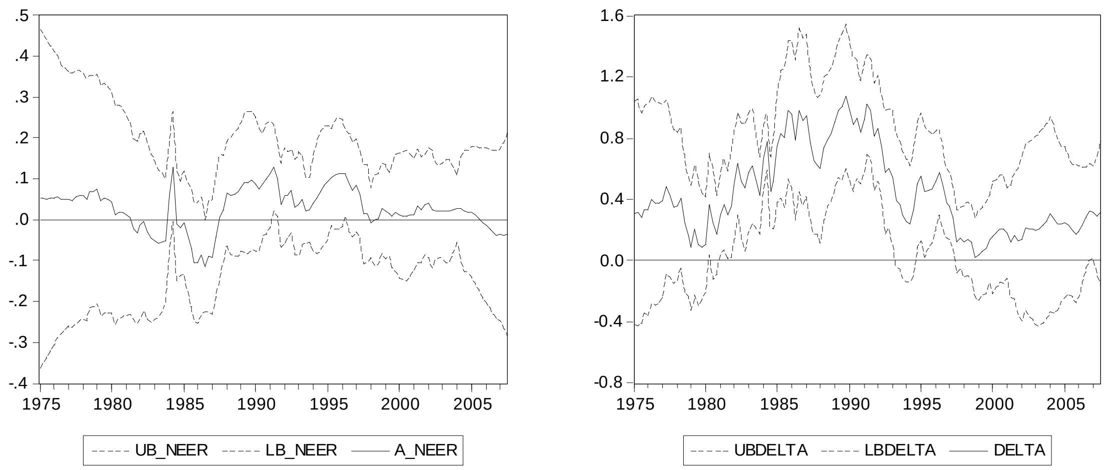
Figure 5 – Time-varying response coefficients in baseline (closed economy) policy rule, Australia



Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The upper-left graph depicts the evolution of the neutral rate. The upper-right graph depicts the evolution of the response of interest rates to inflation. The lower-left graph depicts the evolution of the response of interest rates to the output gap. The lower-right graph depicts the evolution of the interest rate smoothing parameter.

We find that the exchange rate does not have a significant effect on the short-term interest rate (besides the NEER we use also the trade-weighted index – TWI, which is an exchange rate measure reported and often referred to by the RBA) except in the period of 1985–87, when the currency depreciation (the Australian dollar was allowed to float in 1983 after a period of a moving peg vis-à-vis the TWI) was offset by an interest rate increase so as to curb the inflation pressures (see Greenville, 1997). The foreign interest rate parameter is estimated as being always positive, although it is significant only in the 1990s and its importance fades after IT was introduced. After 2001, Australian and U.S. interest rates diverge and the response coefficient approaches zero. This may be related to idiosyncratic developments in the U.S. when the Fed lowered the interest rate so as to face the fear of recession following the September 2001 terrorist attacks.

Figure 6 – Time-varying response coefficients in augmented (open economy) policy rule, Australia



Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The left-hand graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (the deviation from the HP trend). The right-hand graph depicts the evolution of the response of the interest rate to the foreign interest rate.

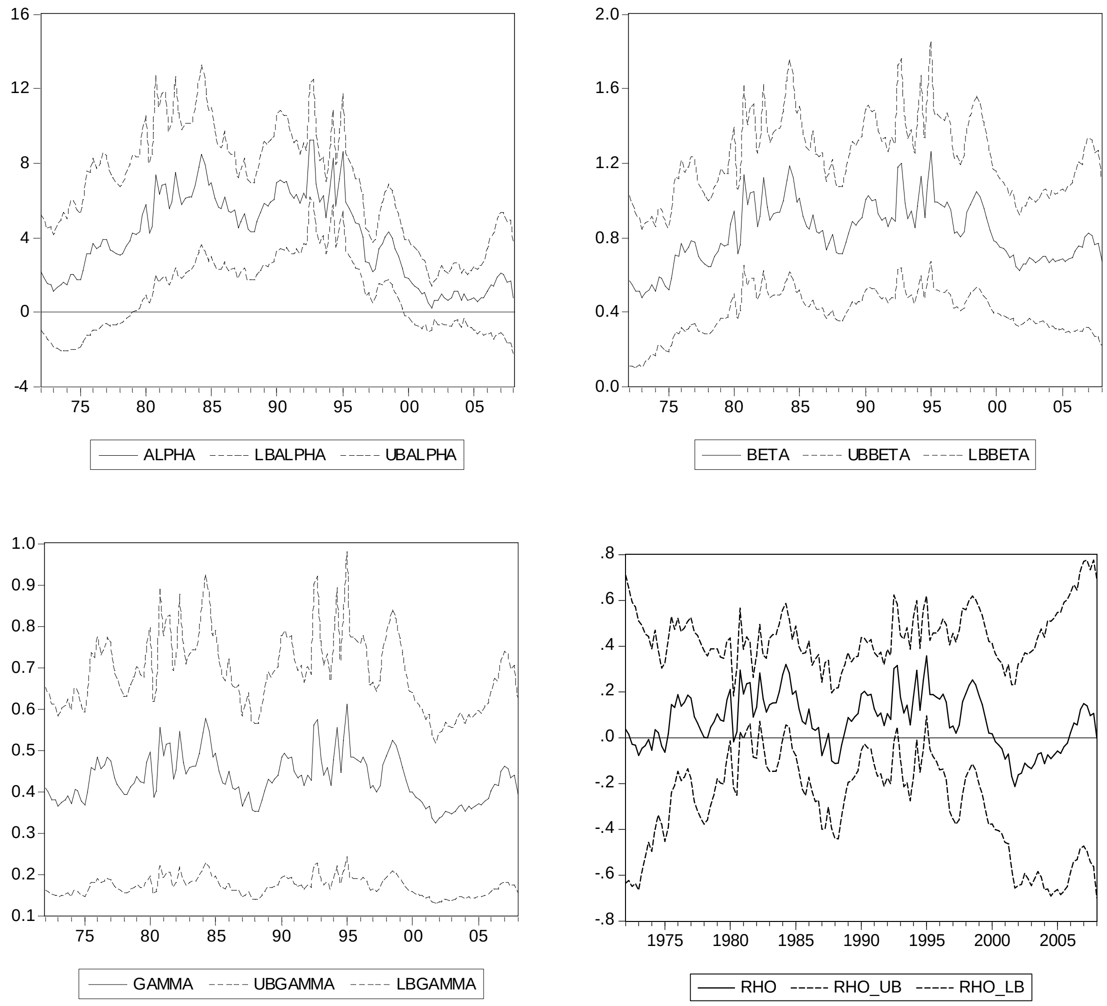
4.4 Canada

The monetary policy rule estimates for Canada are presented in Figures 7 and 8. The response of the interest rate to inflation peaks in the first half of the 1980s and in the mid-1990s. The former period was characterized by relatively high inflation rates, which unquestionably drove the rather aggressive policy of the BoC similarly as in the UK and Australia. It is arguable whether the original inflation rate was a consequence of the accommodative policy of monetary targeting applied between 1978 and 1982 (see Figure 7). Unlike in the UK and Australia, the inflation response coefficient increased temporarily after the announcement of IT, which reflects the fact that IT was adopted as a part of a joint disinflation strategy of the BoC and the federal government. Then it decreases only in the last decade (due to almost negligible inflation rates).[25]

[25] The BoC also reported the monetary condition index (MCI), a compound of the policy instrument (the interest rate) and the exchange rate. The MCI accompanies the proposal of Ball (1999) to target long-term inflation, i.e. the

The response to the output gap is significant and almost invariant in time (as in the UK), confirming the long-term preference of the BoC for smoothing economic fluctuations. The intensity of the response is unique among the IT countries in our sample. The interest rate smoothing is almost negligible.

Figure 7 – Time-varying response coefficients in baseline (closed economy) policy rule, Canada



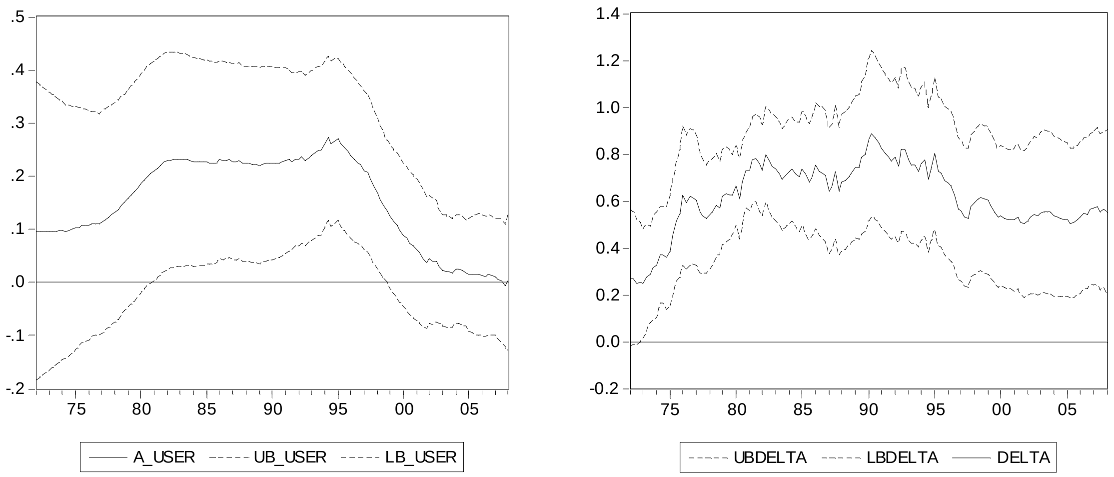
Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The upper-left graph depicts the evolution of the neutral rate. The upper-right graph depicts the evolution of the response of interest rates to inflation. The lower-left graph depicts the evolution of the response of interest rates to the output gap. The lower-right graph depicts the evolution of the interest rate smoothing parameter.

The dependence of Canadian monetary policy on external factors, in particular developments in the U.S., is confirmed in the model augmented by the exchange rate and foreign interest rate. The

inflation rate adjusted for the transitory effect of the exchange rate on import prices. However, there is no indication that the BoC actually ever used the MCI for practical policy making, and it ceased to publish it in 2006.

response to the exchange rate is positive but almost insignificant and dissipates in the last decade. On the other hand, the response to the U.S. interest rate dynamics is substantial for the whole period of analysis until the end of the sample.

Figure 8 – Time-varying response coefficients in augmented (open economy) policy rule, Canada



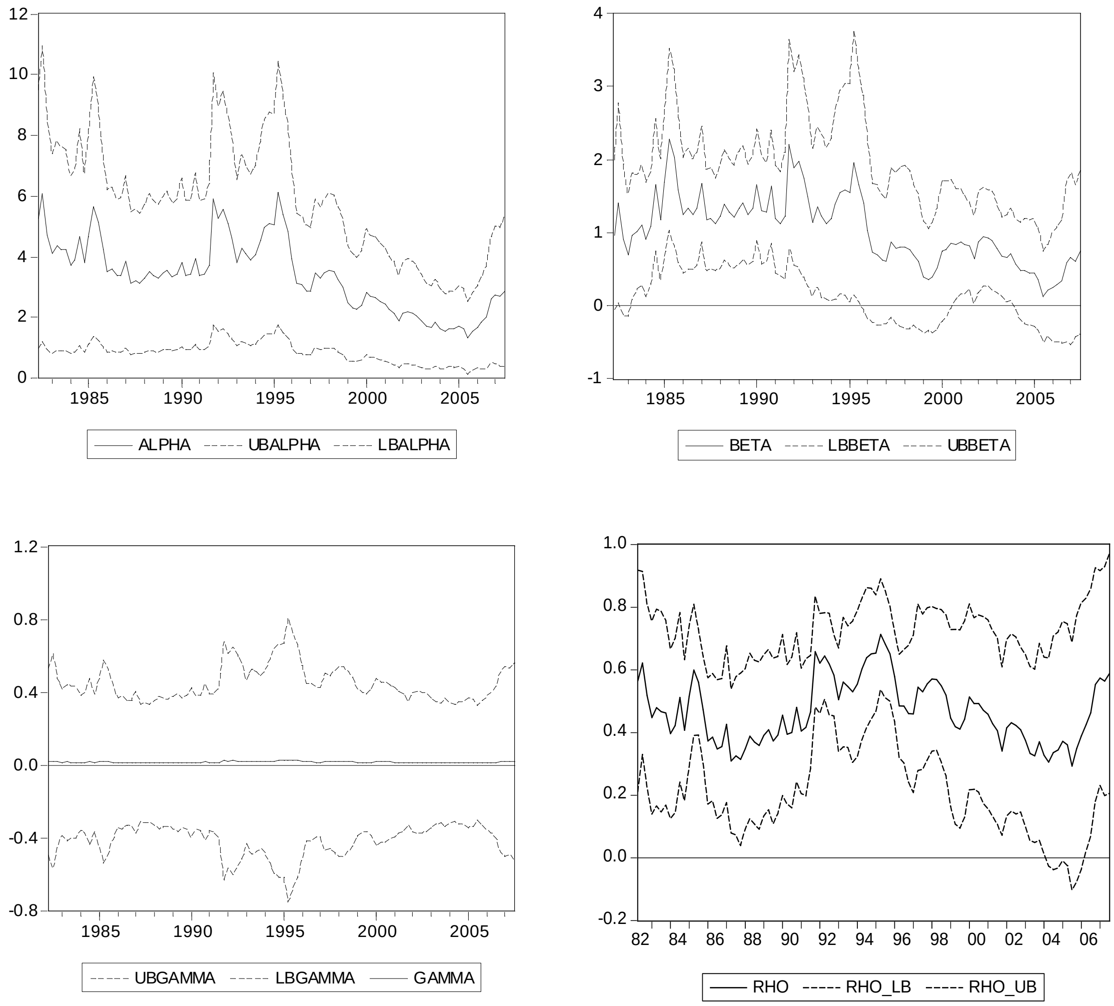
Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The left-hand graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (the deviation from the HP trend). The right-hand graph depicts the evolution of the response of the interest rate to the foreign interest rate.

4.5 Sweden

Our results suggest that the response of interest rates to inflation was stronger before and at the beginning of IT. This is in line with Berg et al. (2004), who argue that the introductory phase of IT in Sweden was characterized by building the credibility of the new regime. The decline of the neutral rate reflects the low inflation environment prevailing in Sweden from the mid-1990s onwards. The time-varying coefficient on interest rate smoothing is estimated to be somewhat larger in Sweden than in the other countries. This suggests that Sveriges Riksbank is likely to smooth its interest rates to a greater degree.[26]

[26] At the time of the ERM crisis (September 1992), the Swedish krona started to depreciate. SR tried (unsuccessfully) to maintain the previous exchange rate and massively increased the short-term interest rate. Consequently, we have included a time dummy in Q3 1992.

Figure 9 – Time-varying response coefficients in baseline (closed economy) policy rule, Sweden

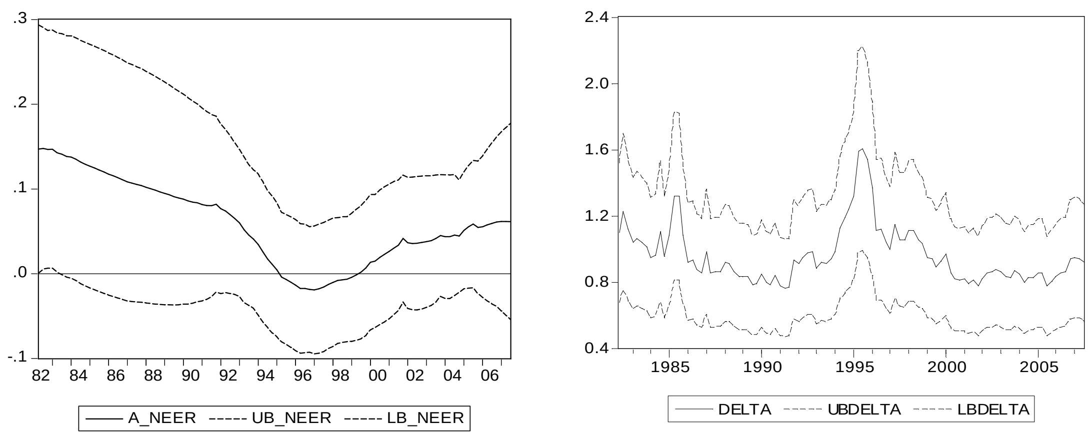


Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The upper-left graph depicts the evolution of the neutral rate. The upper-right graph depicts the evolution of the response of interest rates to inflation. The lower-left graph depicts the evolution of the response of interest rates to the output gap. The lower-right graph depicts the evolution of the interest rate smoothing parameter.

Our results imply a prominent role for external factors rather than the output gap for the determination of Swedish monetary policy. In particular, the coefficient on the foreign interest rate (Eurolibor) is sizeable throughout the sample period, which is rather interesting given that Swedish monetary policy has not officially been subject to any external constraint (at least since the krona's exit from the ERM in 1992). On the other hand, the role of the exchange rate is unclear. The NEER response coefficient is mostly positive, but with very wide confidence intervals, pointing to its insignificance.

Our results are consistent overall with the surveyed time-invariant studies emphasizing the predominant role of the inflation forecast (Jansson and Vredin, 2003) as well as more cautious policy decisions leading to more policy inertia during periods of macroeconomic instability such as the ERM crisis (Berg at al., 2004).

Figure 10 – Time-varying response coefficients in augmented (open economy) policy rule, Sweden



Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The left-hand graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (the deviation from the HP trend). The right-hand graph depicts the evolution of the response of the interest rate to the foreign interest rate.

4.6 Inflation Targeting and Inflation Persistence

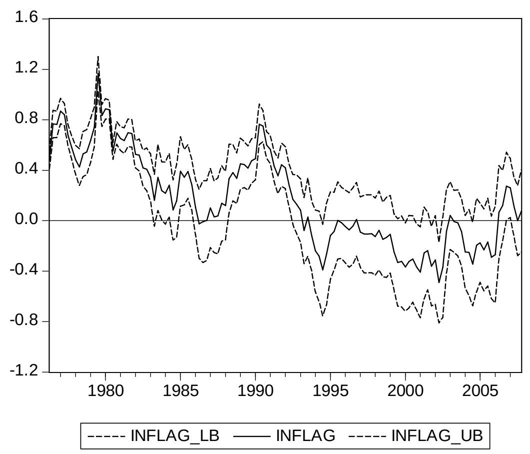
We have related our finding that the inflation response coefficient often falls after the adoption of IT to the hypothesis that this monetary framework has a positive effect on the inflation expectations of economic agents. If expected inflation is low, monetary policy need not be as aggressive as under a discretionary regime in order to achieve price stability. This argument is in line with recent studies on inflation dynamics (Benati, 2008; Zhang et al., 2008) claiming that under a credible policy regime (such as IT), inflation persistence (the dependence of current inflation on past values) fades away.

To shed some light on this issue, we used our estimation framework and fitted the AR(1) model with drift to inflation series, allowing the coefficient on lagged inflation as well as the constant to be time-varying. Our results (as reported in Figure 11) indicate that inflation persistence decreased over time for all the countries. Moreover, it is notable that the persistence fell especially during the 1990s as IT was introduced. This finding is confirmed when, in the spirit of the backward-looking Phillips curve, we include the lagged output gap as a forcing variable.[27] In addition, the results reported in Figure 11 clearly indicate that the moment estimator applied to the time-varying coefficient approach is suitable even if the estimated coefficient is subject to sudden switches rather than smooth transition (see the UK after the adoption of inflation targeting).

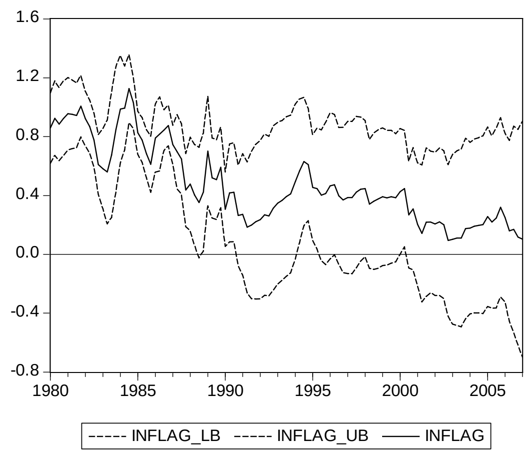
In general, our results are broadly consistent with those of Benati (2008), who performs sub-sample analysis under different policy regimes. Unlike Benati (2008), our approach does not need to impose breaks in the inflation process at any particular date, but simply observes whether and when such breaks occur. Our findings do not exclude the possibility that inflation persistence decreased because of other factors (the "good luck" hypothesis), but the temporal coincidence between the introduction of IT and the significant decrease of inflation persistence in several countries make a case for the "good policy" hypothesis. Taking the example of the UK, we can see that the inflation (rate) moderation goes back to the 1980s, when we still observe rather high inflation persistence in spite of very aggressive anti-inflationary (yet discretionary) policy. Unfortunately, as we do not have a full structural model we cannot tell much about the nature of shocks in the pre- and post-IT period as done in VAR studies on the Great Moderation. Using standard tests of structural stability, we can clearly reject structural stability of the inflation process defined by AR(1) in the pre- and post-IT period.

Figure 11 – Time-varying response coefficients in AR(1) model for inflation

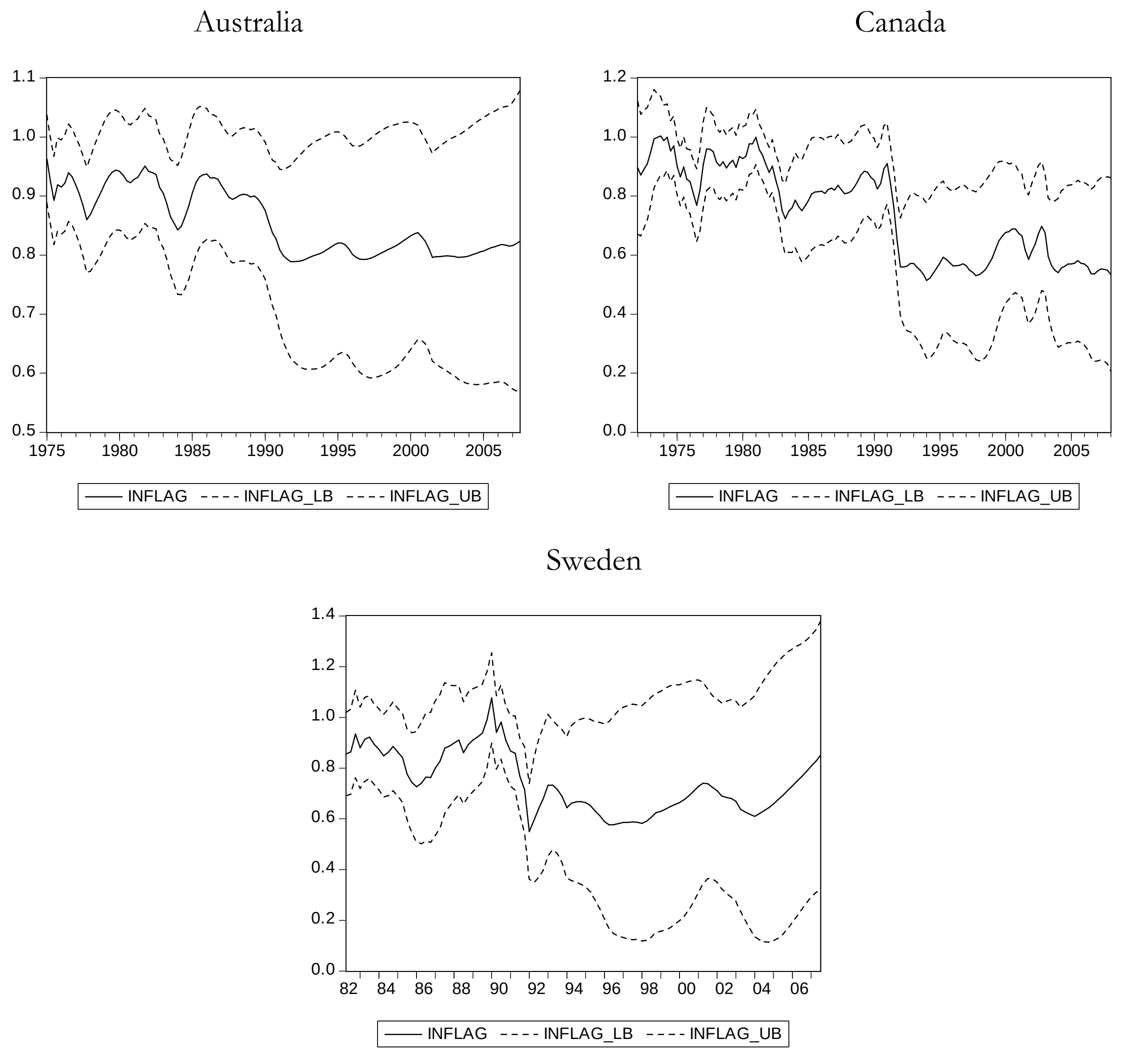
UK



New Zealand



[27] See Hondroyiannis et al. (2009) on the estimation of the Phillips curve within a somewhat different time-varying framework.

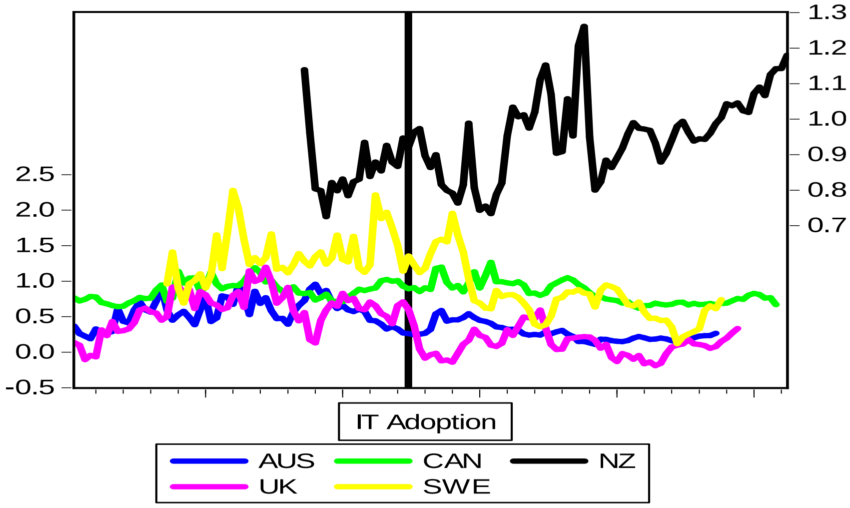


Note: 95% confidence bands; the coefficient of the AR(1) term (i.e. model $\pi_t = \alpha_t + \rho_t \pi_{t-1} + \varepsilon_t$) for the inflation series employed in the previous analysis for each country.

4.7 Monetary Policy Rules – Wrap-Up of Main Policy Findings

This sub-section summarizes the main policy-relevant findings of this paper. We focus on the following four issues: 1) monetary policy aggressiveness and the inflation rate, 2) monetary policy aggressiveness and inflation targeting, 3) interest rate smoothing and 4) inflation persistence and inflation targeting.

Figure 12 – Monetary Policy Aggressiveness and Inflation Targeting

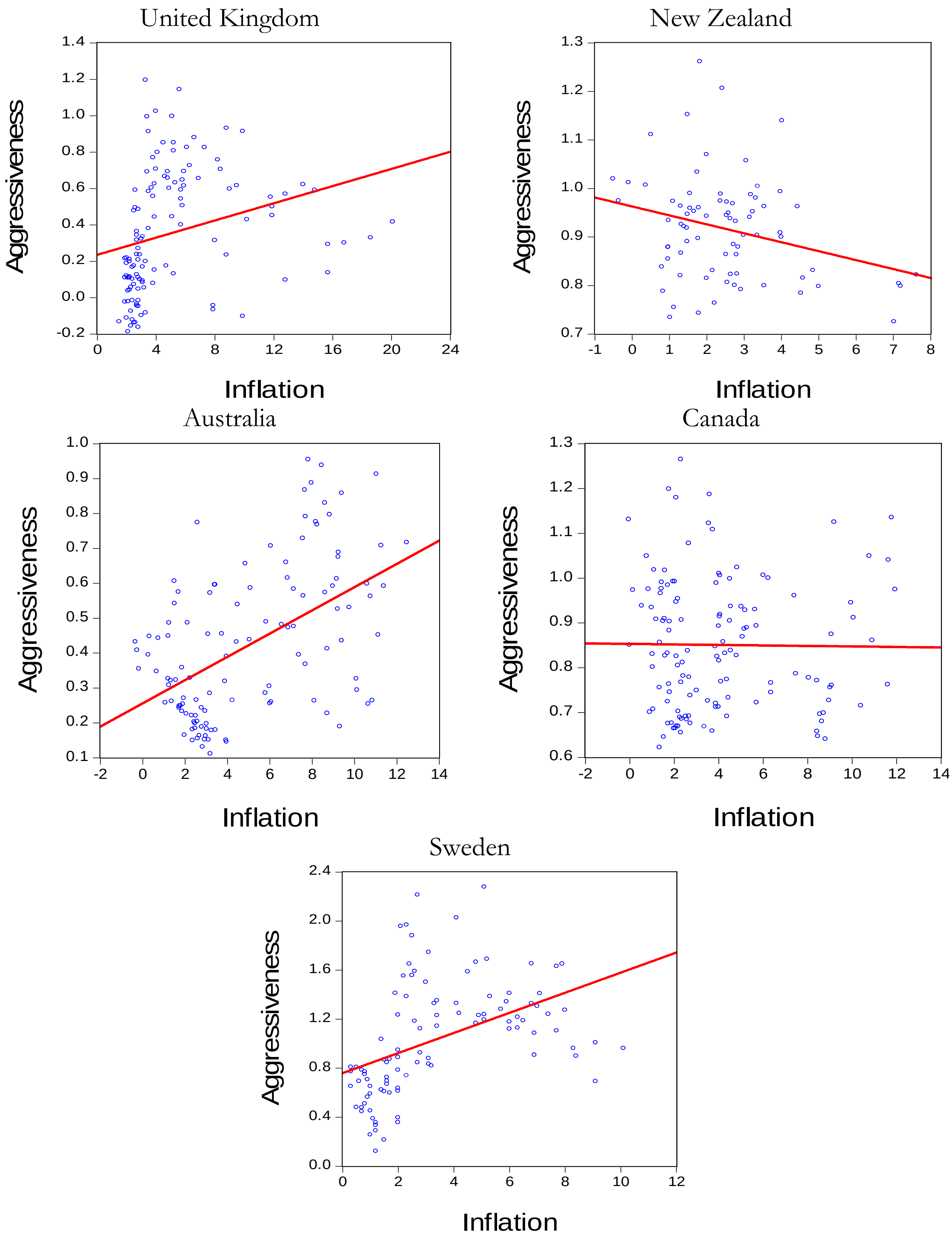


Note: The y axis depicts the evolution of the estimated parameter (β) (the response of interest rates to expected inflation) and the x axis represents time, with the year of inflation targeting adoption denoted by a black vertical line. The value of β for New Zealand is plotted on the right-hand axis.

Figure 12 presents the link between monetary policy aggressiveness (defined as the estimate of the response of interest rates to expected inflation) and inflation targeting. It can be seen that in no country did the aggressiveness parameter increase after the adoption of inflation targeting. In fact, this aggressiveness substantially decreased in the UK, Australia and Sweden. If we look at the link between aggressiveness and expected inflation within the IT period, the results suggest that in most countries the aggressiveness is higher the more expected inflation deviates from its target. This broadly corresponds to the findings of Davradakis and Taylor (2006), who document a non-linear policy rule for the UK of a similar pattern.

Figure 13 documents a relatively strong link between the level of inflation and monetary policy aggressiveness for the UK, Australia and Sweden. This is because 1) the response of interest rates to inflation has been particularly strong during periods when central bankers want to break a record of high inflation and 2) the response has been less aggressive after the adoption of inflation targeting with well-anchored inflation expectations. In contrast, New Zealand and Canada do not seem to exhibit a link between aggressiveness and inflation.

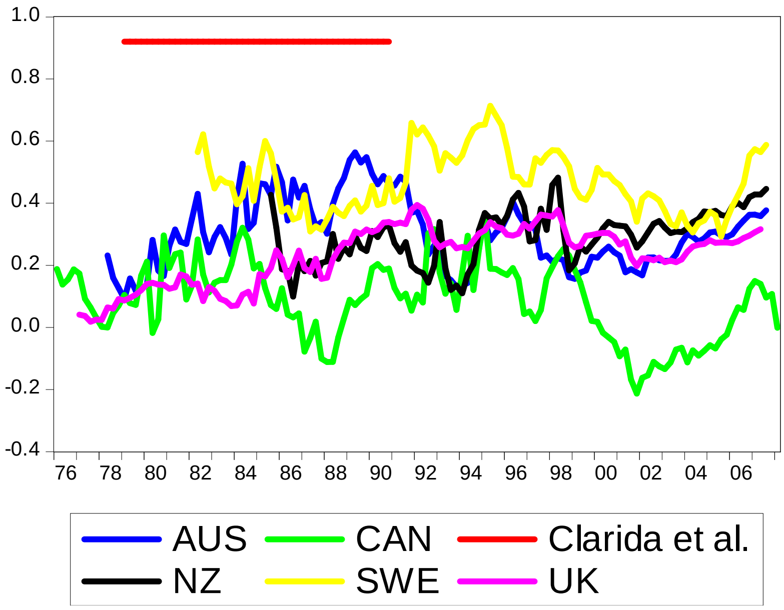
Figure 13 – Monetary Policy Aggressiveness and Inflation Rate



Note: The figure presents the scatter plots between the inflation rate and the response of interest rates to expected inflation (β), labeled as aggressiveness, for each country individually.

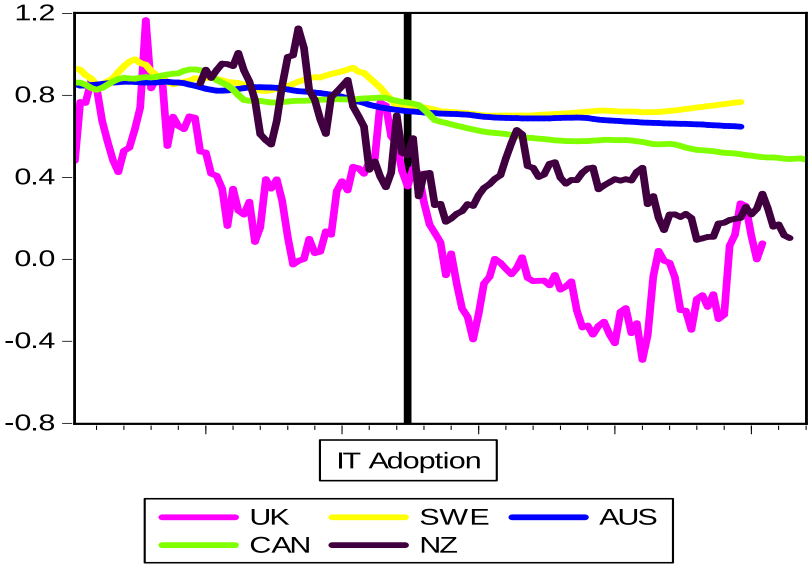
The evolution of the estimated interest rate smoothing parameter in comparison to the time-invariant estimates for the UK in 1979–1990 by Clarida et al. (2000) is available in Figure 14. Our time-varying estimates of interest rate smoothing are well below the time-invariant one, which seems reasonable in the light of the recent critique by Rudebusch (2006), who puts forward that the degree of interest rate smoothing is actually low. While omitted variables or persistent shocks were deemed to be behind the implausible degree of policy inertia, our empirical results suggest that omission of the time-varying nature of the response coefficient may be another reason for the overestimation of smoothing coefficient ρ in time-invariant policy rules. Moreover, given that the policy rule is modeled as a partial adjustment (see Eq. (3)), the overestimation of smoothing coefficient ρ may drive the upper bias in all the other coefficients. This seems to explain why our long-term inflation multiplier β is substantially lower than found in time-invariant studies. Though a high β is consistent with common wisdom, we suspect that it is a by-product of the upper bias in ρ.

Figure 14 – Interest Rate Smoothing



Note: The figure presents the evolution of the estimated interest rate smoothing parameter ρ over time in comparison to the interest rate smoothing parameter estimated in the time-invariant model of Clarida et al. (2000) for the UK.

Figure 15 – Inflation Targeting and Inflation Persistence



Note: The y axis depicts the evolution of the estimated inflation persistence parameter and the y axis represents time, with the year of inflation targeting adoption denoted by a black vertical line.

Finally, the results in Figure 15 plot the estimates of inflation persistence over time for all countries with respect to the inflation targeting adoption date. The results suggest that inflation persistence decreased after the adoption of inflation targeting, with a very distinct fall in the UK and New Zealand.

5 Concluding Remarks

In this paper, we shed light on the evolution of monetary policy in the main inflation targeting central banks during the last three decades. The evolution of monetary policy is evaluated within a novel framework of a time-varying parameter model with endogenous regressors (Kim and Nelson, 2006), further addressing small sample issues (Schlicht, 1981; Schlicht, 2005; Schlicht and Ludsteck, 2006).

In our view, the results point to the usefulness of this econometric framework for analysis of the evolution of monetary policy setting. The estimation of standard monetary policy rules reveals that policy changes gradually and the changes coincide with several important institutional reforms as well as with the periods when the central banks successfully decreased double-digit inflation rates to rates consistent with their definitions of price stability.

In this respect, our results suggest that the response of interest rates to inflation is particularly high during periods when central bankers want to break a record of high inflation, such as in the

UK in the early 1980s. Contrary to common wisdom, the response is often found to be less aggressive after the adoption of inflation targeting, suggesting a positive anchoring effect of this regime on inflation expectations. In other words, monetary policy need not be as aggressive as under a discretionary regime in order to achieve price stability (Kuttner and Posen, 1999). This result is supported by our finding that inflation becomes less inertial and the policy neutral rate decreases after the adoption of inflation targeting.

We find that external factors matter for interest rate setting in all our sample countries. To be more precise, the foreign interest rate is found to enter the monetary policy rule significantly. The importance of the exchange rate varies, being apparently more important before the countries adopted inflation targeting than afterwards.

Our results also indicate that interest rate smoothing is much lower than typically reported by time-invariant estimates of monetary policy rules (see, for example, Clarida et al., 1998, 2000). Our estimates support the recent critique by Rudebusch (2006), who argues that the degree of interest rate smoothing is rather low. We suggest that neglect of changes in monetary policy setting over time is the reason for the implausible degree of policy inertia previously found. Moreover, the fact that upper bias in the smoothing parameter affects all the estimates may explain some fundamental differences between our findings and those established in the literature, such as the size of the inflation response coefficient.

In terms of future research we believe it would be worthwhile to apply this framework to better understand whether and how monetary policy reacts to periods of financial instability and which types of financial instability are the most worrying for central banks. In consequence, this would improve the understanding of both the interest rate setting process and the reaction of monetary policy makers to the current global financial crisis in a more systematic manner.

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Appendix

1 The VC method (Schlicht and Ludsteck, 2006)

In this section, we closely follow the Schlicht and Ludsteck (2006) paper. Consider a standard linear model:

$$y_t = a'x_t + u_t, \quad a, x_t \in \mathbb{R}^n, u_t \sim N\left(0, \sigma^2\right), \quad t = 1,2,\ldots T \tag{A.1}$$

It can be extended for the case in which the coefficients *a* are allowed to follow a random walk. Then equation (A.1) is replaced by a system

$$y_t = a_t'x_t + u_t, \quad u_t \sim N\left(0, \sigma^2\right) \tag{A.2}$$

$$a_{t+1} = a_t + v_t, \quad v_t \sim N\left(0, \Sigma\right) \tag{A.3}$$

with one signal equation (A.2) and *n* state equations (A.3) for each time-varying parameter. The variance-covariance matrix Σ is assumed to be diagonal, that is

$$\Sigma = \begin{pmatrix} \sigma_1^2 & 0 & \ldots & 0 \\ 0 & \sigma_2^2 & \ldots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \ldots & \sigma_n^2 \end{pmatrix}$$

Define the following matrices:

$$X = \begin{pmatrix} x'_1 & 0 & \ldots & 0 \\ 0 & x'_2 & \ldots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \ldots & x'_T \end{pmatrix} \qquad P = \begin{pmatrix} -I_n & I_n & 0 & \ldots & 0 \\ 0 & -I_n & I_n & \ldots & 0 \\ \vdots & \vdots & \ddots & \ddots & \vdots \\ 0 & 0 & \ldots & -I_n & I_n \end{pmatrix}$$

of order T × Tn (T-1)n × Tn

$$y = \begin{pmatrix} y_1 \\ y_2 \\ \vdots \\ y_T \end{pmatrix} \qquad u = \begin{pmatrix} u_1 \\ u_2 \\ \vdots \\ u_T \end{pmatrix} \qquad a = \begin{pmatrix} a_1 \\ a_2 \\ \vdots \\ a_T \end{pmatrix} \qquad v = \begin{pmatrix} v_2 \\ v_3 \\ \vdots \\ v_T \end{pmatrix}$$

of order T × 1 T ×1 Tn × 1 (T-1)n × 1

The system (A.2) and (A.3) can be rewritten as

$$y = Xa + u, \quad u \sim N\left(0, \sigma^2 I_T\right) \tag{A.4}$$

$$Pa = v, \quad v \sim N\left(0, V\right), V = I_{T-1} \otimes \Sigma \tag{A.5}$$

Estimation of the model based on equations (A.4) and (A.5) requires derivation of a distribution function that maps the random variables u_t and v_t to a set of observations X_t. However, such inference is not possible because the matrix P in (A.5) is of rank $(T-1)n$ rather than Tn and thus it cannot be inverted. Furthermore, any v does not determine the path of a_t uniquely.

1.1 Orthogonal parameterization

The VC method used in this paper starts with an explicit definition of a set of possible values of *a* conditioned by matrix *P* and random variable *v*. Following equation (A.5) any solution *a* can be written as

$$a = P'\left(PP'\right)^{-1} v + Z\lambda \tag{A.6}$$

with $\lambda \in \mathbb{R}$ and $Z = \frac{1}{\sqrt{T}}\begin{pmatrix} I_n \\ I_n \\ \vdots \\ I_n \end{pmatrix}$.

Hence equation (A.5) becomes

$$y = u + XP'\left(PP'\right)^{-1} v + XZ\lambda \tag{A.7}$$

Equations (A.6) and (A.7) build an *orthogonal parameterization* of the true model (A.4) and (A5). The orthogonally parametrized model implies that a_t follows a random walk and that its path depends on all realizations of a random variable v_t.[28]

The equation (A.7) can be written as

$$y = XZ\lambda + w \tag{A.8}$$

where

$$w = XP'\left(PP'\right)^{-1} v + u \tag{A.9}$$

Variable w is normally distributed:

$$w \sim N\left(0, W\right), \quad W = XBX' + \sigma^2 I_T \tag{A.10}$$

with

$$B = P'\left(PP'\right)^{-1} V \left(PP'\right)^{-1} P \tag{A.11}$$

Let the matrix of the observations follow a conventional format:

$$X^* = \sqrt{T}\, XZ = X\begin{pmatrix} I_n \\ I_n \\ \vdots \\ I_n \end{pmatrix} = \begin{pmatrix} x_1' \\ x_2' \\ \vdots \\ x_T' \end{pmatrix} \tag{A.12}$$

Inserting (A.12) into (A.8) implies a generalized linear regression model

[28] To avoid excessive number o indexes, we skipped the time index *t* in the latter part of the text.

$$y = \frac{1}{\sqrt{T}} X^* \lambda + w = X\beta + w \tag{A.13}$$

with

$$\hat{\beta} = \frac{1}{\sqrt{T}} \hat{\lambda} \tag{A.14}$$

The estimate of $\hat{\lambda}$ satisfies

$$\hat{\lambda} = \left(Z'X'W^{-1}XZ\right)^{-1} Z'X'W^{-1}y \tag{A.15}$$

which is a standard GLS estimator of the classical regression problem with a covariance matrix of residuals W and observations ZX. Taking expectations of a from (A.6) and substituting $\hat{\lambda}$ for λ implies $Z'a = \lambda$ and hence $\frac{1}{T}\sum_{t=1}^{T} a_t = \beta$ in the GLS regression (A.13).

1.2 Estimation of coefficients

The orthogonal parameterization derived in the previous section might be used for direct ML estimation of the time-varying parameters *a*. However, the derivation of the ML estimate of the vector of parameters *a* leads to a formulation that is equivalent to the minimization of the weighted sum of squares

$$\sum_{t=1}^{T} u^2 + \theta_1 \sum_{t=1}^{T} v_1^2 + \theta_2 \sum_{t=1}^{T} v_2^2 + \ldots + \theta_n \sum_{t=1}^{T} v_n^2 \tag{A.16}$$

where the weights θ_i are the inverse variance ratios of the regression residuals u_t and the shocks in time-varying coefficients v_t, that is $\theta_i = \sigma^2 / \sigma_i^2$. The proof can be found in Schlicht and Ludsteck, 2006, section 5. Hence the estimator balances the fit of the model and the parameter stability.[29]

Now we derive the formula used for estimation of the coefficients. For given X and y the estimated disturbances are

$$\begin{aligned} \hat{u} &= y - X\hat{a} \\ \hat{v} &= P\hat{a} \end{aligned} \tag{A.17}$$

Using the expressions for the estimated disturbances (16), minimization of the weighted sum of squares (15) implies

$$\left(X'X + \sigma^2 P'V^{-1}P\right)\hat{a} = X'y \tag{A.18}$$

which is used for the estimation of coefficients $\hat{a}_t$.

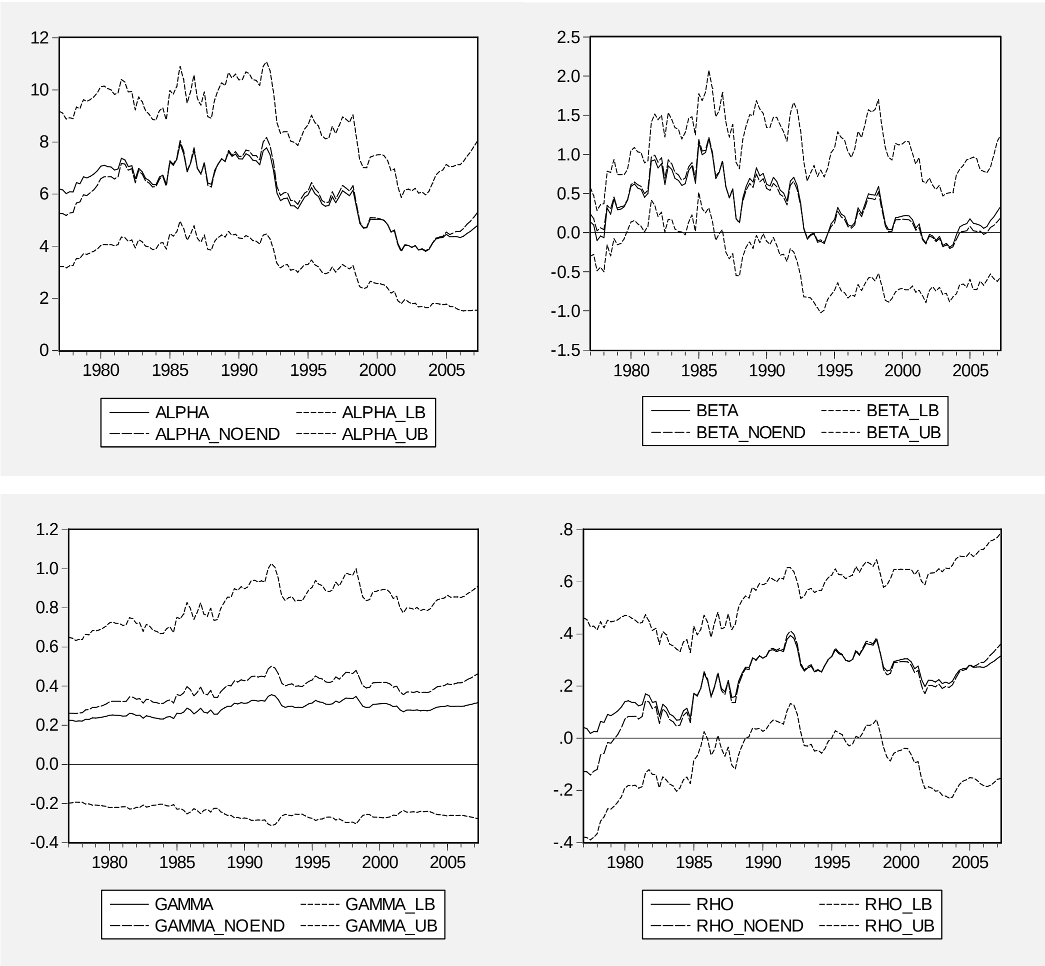
[29] Originally, Schlicht and Ludsteck (2006) start with a derivation of the maximum likelihood estimator of parameters *a* based on the idea of orthogonal parameterization, which is described in the Appendix. Then they prove that the weighted least squares estimator is identical to the maximum likelihood estimator.

The coefficients estimated using the VC method have a straightforward interpretation: they have a time-invariant part, determined by a regression with fixed coefficients, and a random part reflecting the idea that some proportion of the variance of the dependent variable is caused by a change in the coefficients.

The estimation procedure proceeds as follows. The iterative procedure has two steps. First, given the variances of the residuals in both equations in (15), σ^2 and σ_i^2, the coefficients are estimated using (17). Second, the estimated residuals are calculated using (16) and their estimated second moments $\hat{u}'\hat{u}$ and $\hat{v}'_i\hat{v}_i$ are compared to their expected moments $E\left[\hat{u}'\hat{u}\right]$ and $E\left[\hat{v}'_i\hat{v}_i\right]$. These steps are repeated until the estimated moments are identical to their expected counterparts (for a precise derivation of the moment estimator as well as computational details see Schlicht and Ludsteck, 2006, sections 6–9).

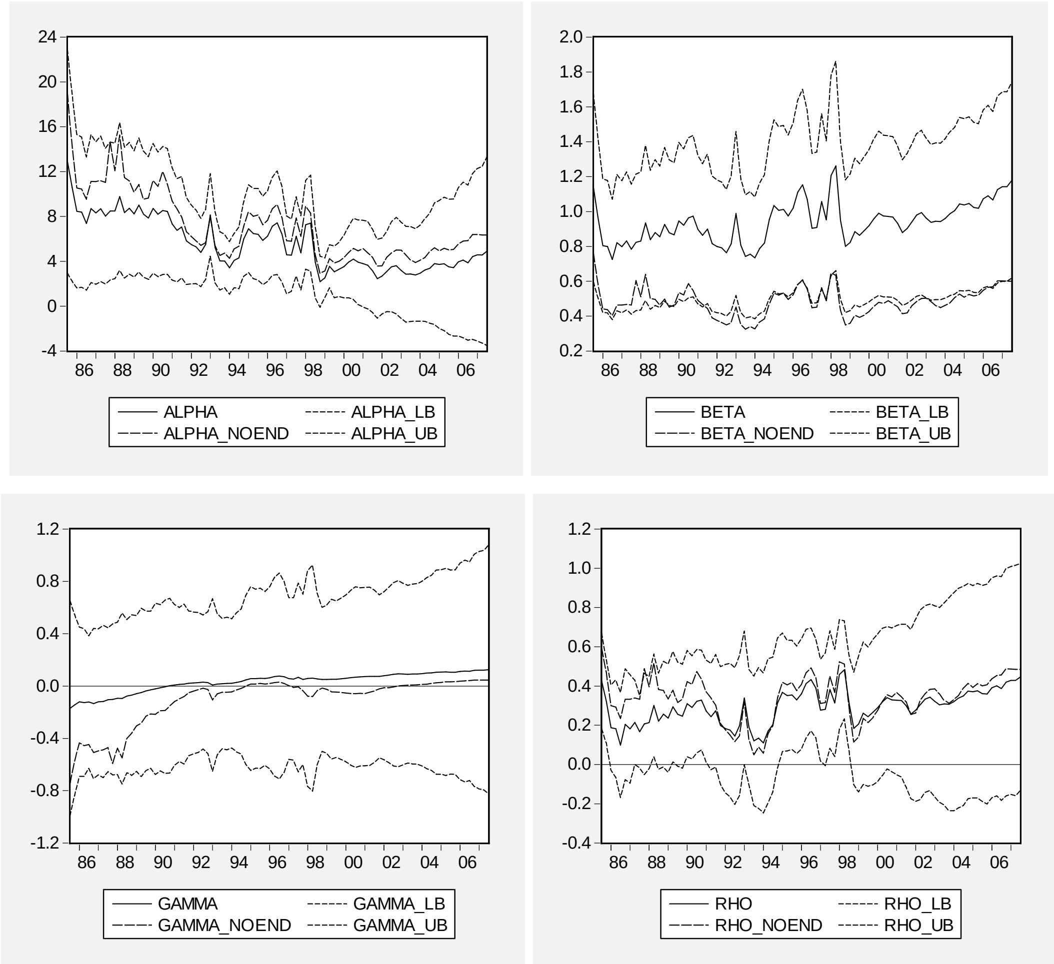
2 The effect of endogeneity in the policy rule

Figure A.1 – Time-varying response coefficients in baseline (closed economy) policy rule, UK



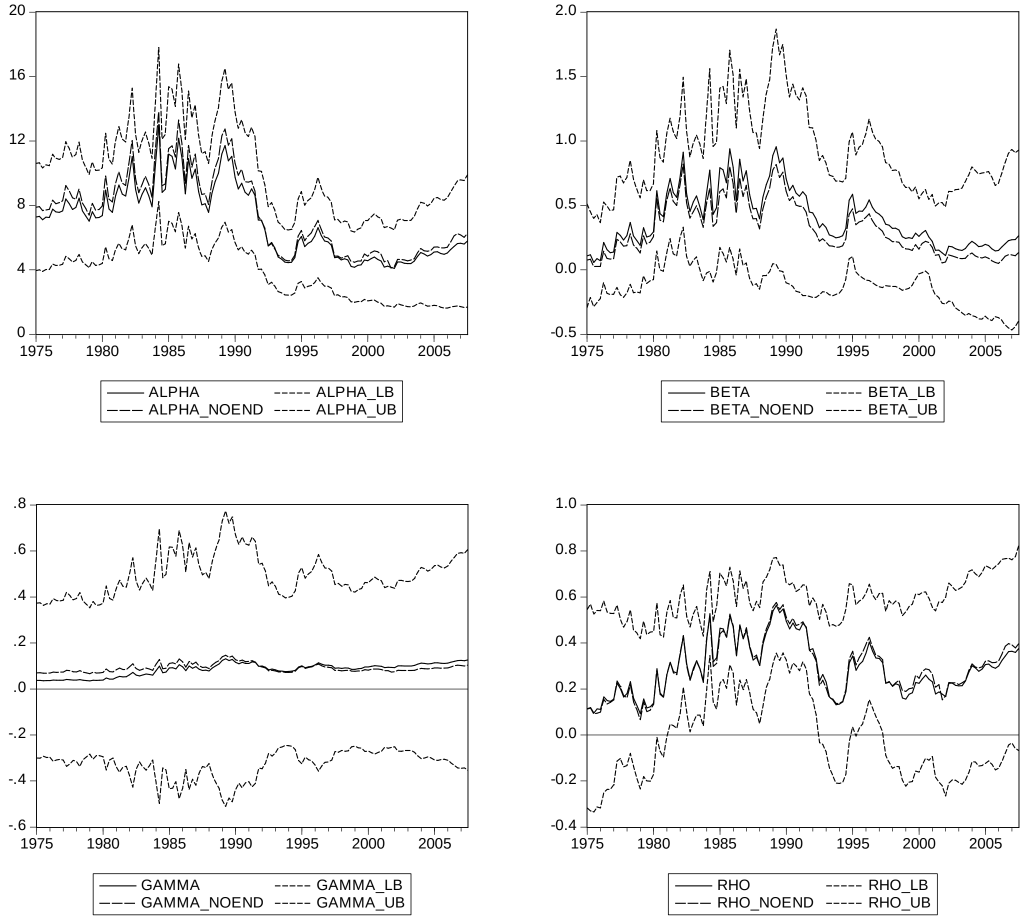
Note: Model with bias correction terms (solid line); model not dealing with endogeneity in monetary policy rules (dashed line). The 95% confidence bands correspond to the model with bias correction terms.

Figure A.2 – Time-varying response coefficients in baseline (closed economy) policy rule, New Zealand



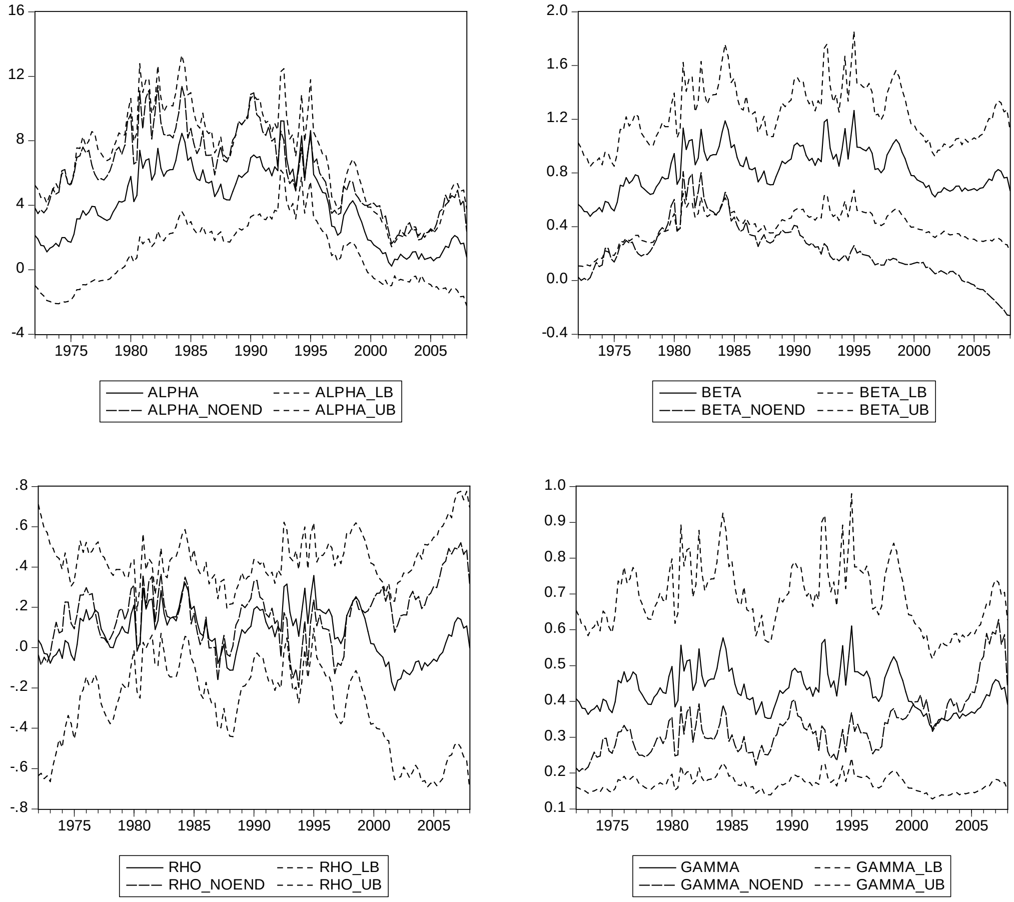
Note: Model with bias correction terms (solid line); model not dealing with endogeneity in monetary policy rules (dashed line). The 95% confidence bands correspond to the model with bias correction terms.

Figure A.3 – Time-varying response coefficients in baseline (closed economy) policy rule, Australia



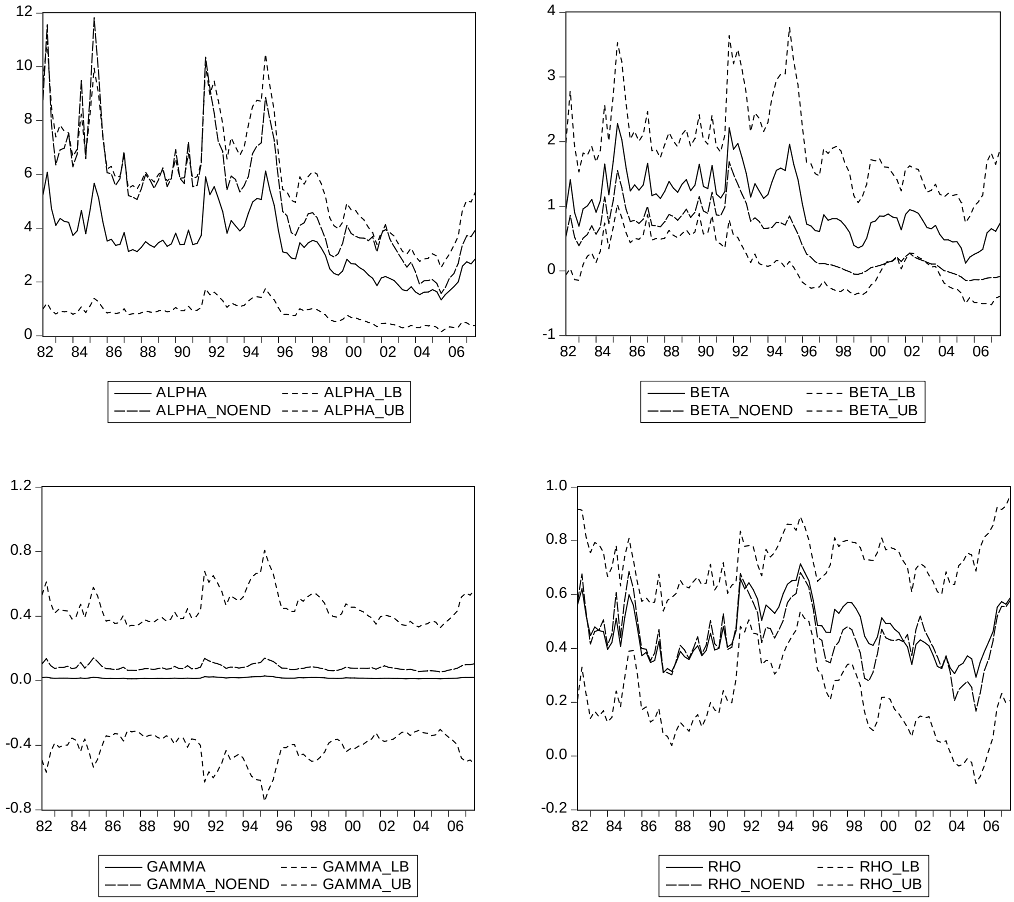
Note: Model with bias correction terms (solid line); model not dealing with endogeneity in monetary policy rules (dashed line). The 95% confidence bands correspond to the model with bias correction terms.

Figure A.4 – Time-varying response coefficients in baseline (closed economy) policy rule, Canada



Note: Model with bias correction terms (solid line); model not dealing with endogeneity in monetary policy rules (dashed line). The 95% confidence bands correspond to the model with bias correction terms.

Figure A.5 – Time-varying response coefficients in baseline (closed economy) policy rule, Sweden



Note: Model with bias correction terms (solid line); model not dealing with endogeneity in monetary policy rules (dashed line). The 95% confidence bands correspond to the model with bias correction terms.

Table A.1 Estimated Coefficients of Endogeneity Correction Terms

		UK	NZ	Aus	Can	Swe
Inflation	mean	-0.0183	**-1.06**	*-0.17*	**-0.873**	**-0.329**
	s.e.	0.269	0.337	0.0912	0.146	0.119
GDP gap	mean	0.0745	-0.031	0.014	**-0.127**	0.0781
	s.e.	0.125	0.27	0.092	0.0559	0.0808

Note: Bold: sign. at 5%, italic: sign. at 10%.

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