How Does Monetary Policy Change?
Evidence on Inflation Targeting Countries

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Abstract

In this paper, 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 moment estimator at time-varying parameter model with endogenous regressors. This methodology has several important advantages for estimation of policy rules. In particular, unlike the Markov-switching methods, it models the policy as gradually evolving rather than imposing its sudden switches from (one regime to another). It also deals with the issue of endogeneity in policy rules and delivers superior statistical properties in small samples than the traditional Kalman filtering. Our main findings are threefold. First, monetary policy changes gradually, pointing to the importance of applying time-varying estimation framework. Second, the interest rate smoothing parameter is much lower that what previous time-invariant estimates of policy rules typically report. External factors matter for all countries, albeit the importance of the exchange rate diminishes after the adoption of inflation targeting. Third, the short-term response of interest rates on inflation is particularly high during the periods, when central bankers want to break the record of high inflation such as in the U.K. or in Australia at the beginning of 1980s. Contrary to common wisdom, the response becomes less aggressive after the adoption of inflation targeting suggesting the positive effect of this regime on anchoring inflation expectations. This hypothesis is supported by our finding that inflation persistence as well as policy neutral rate typically decreased after the adoption of inflation targeting.

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1 Introduction

The Taylor-type regressions have been applied extensively in order to describe the monetary policy setting for many countries. The research on the U.S. monetary policy usually assumed that the monetary policy was subject to structural breaks when the FED chairman changed. Clarida et al. (2000) claimed that the U.S. inflation was unleashed during the 1970's because the FED's interest rate response to inflation upsurge was too weak, while the increase of such response in the 1980's was behind the inflation moderation. Although there is ongoing discussion on the sources of this Great Moderation (Benati and Surico, 2009), the fact that monetary policy setting evolves over time is generally accepted. While evidence for the U.S. is vast, knowledge about the evolution of monetary policy for other countries often does not go beyond common wisdom (Nelson, 2004).

Both the evolution of monetary policy setting and the exogenous changes in the economic system over time represent a problem for empirical analysis. In particular, the coefficients of monetary policy rules estimated over longer periods are structurally unstable. The common solution used in the literature was a sub-sample analysis (Clarida et al. 1998, 2000). Such an approach is based on rather strong assumption that the moments of structural breaks are known, but also that the policy setting did not evolve within each sub-period. Consequently, this gives impetus to applying the empirical framework that allows for regime changes or in other words the time variance in the model parameters (Cogley and Sargent, 2001, 2005). Countries that 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 IT implementation. Moreover, there is ongoing debate whether the IT represents a rule-based policy given the substantial degree of discretionality the IT central banks have to achieve the inflation target. Bernanke et al. (1999) claim the IT is a framework or a constrained discretion rather than a mechanical rule. Consequently, the monetary policy rule of IT central bank is by nature time varying and its empirical analysis must reflect it.

Our study aims to investigate the evolution of monetary policy for the countries that have had a 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. Because we are interested in the monetary policy evolution over relatively longer periods, we are not considering countries where the IT was in place for a relatively short period of time (Finland, Spain), or was introduced recently (such as Armenia, Hungary, Korea, Norway, South Africa or Switzerland).

We apply the recently developed time-varying parameter model with endogenous regressors (Kim and Nelson, 2006), as this technique allows us to evaluate the changes in policy rules over time as well as to deal with endogeneity of policy rules. Unlike Kim and Nelson (2006) we do not rely on the Kalman filter that is conventionally employed to estimate time-varying models, but employ the moment-based estimator proposed by Schlicht and Ludsteck (2006) than has superior statistical properties in small samples.

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. This is so, as the regressors that are included in the corresponding equations to address the endogeneity are found to be statistically significant. Second, the interest rate smoothing parameter is much lower that what previous time-invariant estimates of policy rules typically report. This is in line with a recent critique of Rudebush (2006), who emphasizes that the degree of smoothing is rather low. External factors matter for all countries, albeit the importance of the exchange rate diminishes after the adoption of inflation targeting. Third, the response of interest rates on inflation is particularly high during the periods when central bankers desired to break the record of high inflation such as in the U.K. at the beginning of 1980s. Contrary to common wisdom, the response can become less aggressive after the adoption of inflation targeting suggesting the positive effect of this regime on anchoring inflation expectations or low inflation environment. This result is consistent with Kuttner and Posen (1999) who show that transition from discretion to inflation targeting should imply increased response of short-term interest rate to inflation shocks. However, if price stability has already been achieved and inflation shocks are not sizeable, there is no reason why the interest rate response should substantially increase.[1]

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; and Section 5 features the conclusion. An appendix with detailed description of the methodology and additional results follows.

[1] Of course, the price stability is sustainable only if the monetary policy is credible and the economic agents believe that the central bank would be determined to fight the inflation shock if it appeared.

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

New Zealand was the first country to adopt the IT (in 1990). A particular feature besides an announcement of the inflation target (the band of 1-3%, currently) is that the governor of the Reserve Bank has an explicit agreement with the government. Huang et al. (2001) study the monetary policy rule over the first decade of the IT. They find that the policy of the Reserve Bank (RBNZ) was clearly aimed at inflation target and did not explicitly respond to output fluctuations. The response to inflation was symmetric and backward-looking rule does as good job as forward-looking one for 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 since the IT was implemented and that 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) disregards asymmetries in the RBNZ policy rule.

The Reserve Bank of Australia (RBA) turned to the IT in 1991 (with the target of 2-3%) after decades of exchange rate pegs (till 1984) and consecutive monetary targeting. Using a sub-sample analysis, de Bouwer and Gilbert (2005) confirm that RBA's consideration of inflation was very low in pre-IT period, and that the concern for output stabilization was clearly predominant. The response to inflation (both actual and expected) increased substantially since the 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 policy (bad fit of time-invariant rule) in pre-IT period, a consistent response to inflation during the IT, and signs of asymmetry in both periods. Karedekikli and Lees (2007) document that the policy asymmetry is related to RBA's distaste for negative output gaps, which may be a bit surprising for IT central bank.

The Bank of Canada (BoC) introduced the IT in 1991 (the current target of 1-3%). Demers and Rodríguez (2002) find that an 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 monthly growth rate of real GDP drive the changes of policy rate, while unemployment rate and the exchange rate do not. On the contrary, Dong (2008) confirms the BoC considerations for real exchange rate movements.

Sweden adopted the IT in 1993 (2% target with a tolerance band of 1 percentage point) just after the korona was allowed to float. The independence of Sveriges Rissbank (SR) was legally increased in 1999. Jansson and Vredin (2003) study its policy rule, concluding that the inflation forecast (published by the bank) is the only relevant variable driving interest rate changes. Kuttner (2004) finds additionally role for the output gap but in terms of its growth forecast (rather than its observed value). Berg et al. (2004) provide rigorous analysis of the sources of deviations between the SR policy rate and the targets implied from diverse empirical rules. They claim that higher inflation forecasts at early stages of the IT regime (due to lack of credibility) generates higher implied target from the forward-looking rule and therefore induce spurious indication of policy shock. Their qualitative analysis of SR documents clarifies the rationale behind the actual policy shocks, such as more gradualism (stronger inertia) in periods of macroeconomic uncertainty.

Finally, there are a few multi-country studies. Meirelless-Aurelio (2005) analyzes the time-invariant rule of the same countries as us, finding a significant dependency of the results on the real-time versus historic measures of variables. Lubik and Shorfheide (2007) estimate an open economy structural model of 4 IT countries (AUS, CAN, NZ, UK) by Bayesian methods, with the aim to see whether the IT central bank responds to exchange rate movements. They confirm this claim for the BoE and BoC. Dong (2008) enrich their setting by some more realistic assumptions (the exchange rate endogeneity, incomplete exchange-rate pass-through), finding additionally that the RBA responded to the exchange rate.

2.2 Time variance in monetary policy rules

The original empirical research on monetary policy rules used linear specification with time-invariant coefficients. Instrument variable estimators like 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).[2] While the 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 to take the time variance into account is to use a sub-sample analysis (Taylor, 1999, Clarida et al., 2000). The drawback of this approach is a rather subjective assumption about the points of structural change and the structural stability within each sub-period. An alternative is to apply an econometric model that allows for time variance in the coefficients. There are various methods dealing with the 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) employes such model with the switches in the constant term 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

[2] One exception is when 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 a case, endogeneity problem shall not arise and least square 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.

(2006) uses the Markov-switching model with the shifts both in the coefficients and the residual variances. Such separation between the evolution of policy preferences (coefficients) and the exogenous changes in economic system (residuals) is important for the continuing discussion on 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 the macroeconomic development in the US in last four decades.

The application of Markov-switching VAR techniques turns out to be complicated for the 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 backward-looking specification (lagged explanatory variables) but this is very likely inappropriate for IT central banks that are arguably forward-looking.[3] However, there is another distinct feature of the Markov-switching model that makes its use for the analysis of time variance in monetary policy rule problematic. The model assumes sudden switches from one policy regime to another rather than a gradual evolution of the monetary policy. Although at first sight, one may consider the introduction of the IT being an abrupt change, there are some reasons to believe that the smoothed transition of monetary policy is more appropriate description for the 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 the sudden switches. Secondly, it is likely that inflation played already a certain role for the interest rate setting, even before the IT regime was introduced, because in many countries the major decrease of inflation rates occurred before the moment when the IT was implemented. Thirdly, the coefficient of different variables (such as inflation, the output gap or exchange rate) in the monetary policy rule may evolve independently, rather than moving from one regime to another at the same time. For instance, a central bank may assign more weight to an observed or expected inflation rate when it implements the IT but it does not mean that it immediately disregards the information on real economic activity or foreign interest rate setting. Finally, there is relevant evidence, though mostly for the U.S., that the monetary policy evolves rather smoothly over time (Boivin, 2006, Canova and Gambetti, 2008). Therefore, a smooth transition seems to be a more adequate description of the reality.

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

Besides simple recursive regression (e.g. Domenech et al., 2002), the Kalman filter was employed in a few studies to estimate a coefficient vector that varies over time. This approach assumes smooth evolution of the monetary policy rule since its parameters are assumed to follow an autoregressive process. Such time-varying model is also suitable for reflection of possible asymmetry of the monetary policy rule (Dolado et al., 2004). [4] 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 time varying model estimated via the Kalman filter for the US, Elkhoury (2006) for Switzerland, and Trecrocci and Vasalli (2007) for the US, the UK, Germany, France and Italy. However, none of these studies provide a specific econometric treatment to the endogeneity that arise in forward-looking specifications.

In this respect, Kim (2006) proposes a two-step procedure to deal with the endogeneity problem in the time-varying parameter model. Using this methodology, Kim and Nelson (2006) find that the US monetary policy evolved in a different manner than the previous research suggested. In particular, they document that FED's interest in stabilization of real economic activity significantly increased since early 1990's.[5] Kishor (2008) applies the same technique for the analysis of time-varying monetary policy rule of Japan, Germany, the UK, France and Italy. He detects time-varying response, not only with respect to the inflation rate and the output gap but also to foreign interest rate. The relevance of the endogeneity correction can be demonstrated by the difference between Kishor's results and those of Trecrocci and Vasalli (2007), given that both study the same sample of countries.[6] The time-varying parameter model with specific treatment of endogeneity can be relevant even when the 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 forecasting horizon, the endogeneity may be still present in the model (see Boivin, 2006). Moreover, this estimation procedure is also viable to reflect the measurement error and the heteroscedasticity in the model (Kim et al., 2006). However, the Kalman filter applied at state-space model may suffer one important drawback in small samples, as it is rather sensitive to the initial values of the parameters, which are unknown. The moment-based estimator proposed by Schlicht and Ludsteck (2006), which is employed in our paper and described below, avoids this problem. Moreover, it is flexible enough so to incorporate the endogeneity correction proposed by Kim (2006).

[4] Granger (2008) shows that any non-linear model can be approximated by a time-varying parameter linear model.
[5] Kim et al. (2006) confirms this finding with real-time data and additionally detected significant decrease in the response to expected inflation during the 1990's.
[6] Horváth (2009) employs the time-varying model with endogenous regressors for estimation of the neutral interest rate for the Czech Republic confirming the importance of endogeneity bias correction terms.

3 Empirical Methodology

3.1 The empirical model

Pursuant to Taylor (1993), most empirical studies assumes that the central bank adjust the nominal interest rate to the state of the economy (see Clarida *et al.*, 1998, 2000). Such policy rule, which is arguably forward-looking in the case of the IT central bank, can be written as follows:

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

where r_t^* denotes the targeted interest rate, $\bar{r}$ is the policy neutral rate, 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[7]. 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 when interest rates are set. Eq. (1) links the policy instrument (nominal interest rates) to a constant term (the neutral rate – that would prevail if the 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 the actual interest rate setting. First, it does not account for the interest rate smoothing of central banks. Most studies in line with Clarida *et al.* (1998) assume that the central bank adjusts the interest rate sluggishly to the targeted value, which can be tracked by simple partial-adjustment mechanism:

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

where $\rho \in [0,1]$ is the smoothing parameter.[8] Although this is line with common wisdom that central banks are averse to abrupt changes, most studies estimating time-invariant models find an 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

[7] The definition of inflation target varies slightly across the IT countries. However, mostly it aims at mid-term rather than short-term period. The target value can also vary in time, which was pronounced especially in emerging countries that implemented IT as a gradual disinflation strategy. On the contrary, for countries that are studied here, the target value has not significantly changed over time.

[8] We estimate also the monetary policy rules including higher lags of interest rates, but fail to find them significant.

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 the interest rate smoothing is logically reinforced in 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 inflation rate and the output gap. However, many central banks that implemented the IT are small open economies that may consider additional variables, in particular the exchange rate and foreign interest rate. Therefore, in our empirical model we substitute Eq. (2) into Eq. (1), eliminating unobserved forecast variables, and we include 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 the constant term, which in Eq. (1) coincide with the policy neutral rate $\overline{r}$. However, if the model is augmented by additional variables x_{t+k} that are not in a deviation from target value, the constant term does not have to keep this interpretation. We set *i* equal to 2.[9] Consequently, the disturbance term ε_t is a combination of forecast errors and is thus orthogonal to all information available at time t (Ω_t).

Pursuant our previous discussion, the interest rate rule described above will be estimated within a framework that allows for time variance in the coefficients. Kim (2006) shows that the conventional time-varying parameter model (Kalman filter applied at state-space representation) delivers inconsistent estimates when explanatory variables are correlated with the disturbance term. The 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 the estimation of monetary policy rules (Kim and Nelson, 2006, Kim et al. 2006, Kishor, 2008). [10] Following Kim (2006) we can rewrite Eq. (3) as follows:

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

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

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

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

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

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

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

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

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

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

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. The Eqs. (10) and (11) tracks the relationship between the endogenous regressors (π_{t+i} 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). We assume that the parameters in Eqs. (10) and (11) are time-invariant. Next, the correlation between the standardized residuals φ_t and ν_t and ε_t is $\kappa_{\varphi,\varepsilon}$ and $\kappa_{\nu,\varepsilon}$, respectively (note that σ_φ and σ_ν are standard errors of φ_t and ν_t, respectively). The consistent estimates of the coefficients in Eq. (4) are obtained in two steps. In the first step, we estimate the equations (10) and (11) and save the standardized residuals φ_t and ν_t. In the second step, we estimate Eq. (12) below along with Eq. (5) – (9). Note that (12) now includes bias correction terms[11], i.e. (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 ε_t is uncorrelated with the regressors.

$$r_t = (1-\rho_t)\left[\alpha_t + \beta_t\left(\pi_{t+i} - \pi^*_{t+i}\right) + \gamma_t y_t + \delta_t x_t\right] + \rho_t r_{t-1} + \kappa_{\nu,\varepsilon}\sigma_{\varepsilon,t}\nu_t + \kappa_{\varphi,\varepsilon}\sigma_{\varepsilon,t}\varphi_t + \varepsilon_t$$

$$,\ \zeta_t \sim N\left(0, (1-\kappa^2_{\nu,\varepsilon} - \kappa^2_{\varphi,\varepsilon})\sigma^2_{\varepsilon,t}\right) \tag{12}$$

[11] 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 significant at conventional significance levels (not reported, available upon request).

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 in applied work. First, if the variables are nonstationary, the results often depend on a proper choice of initial values, but those values are not known in advance.[12] The problem is even reinforced in small samples. Second, the Kalman filter uses only past information for estimation and thus the estimates are less efficient at the beginning of the sample.

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 squares of residuals $\sum_{t=1}^{T}\varepsilon^2$, it uses the minimization of the weighted sum of squares:

$$\sum_{t=1}^{T}\varepsilon^2+\theta_1\sum_{t=1}^{T}\vartheta_1^2+\theta_2\sum_{t=1}^{T}\vartheta_2^2+\ldots+\theta_{\mathrm{n}}\sum_{t=1}^{T}\vartheta_{\mathrm{n}}^2 \qquad (13)$$

where the weights θ_i are the inverse variance ratios of the regression residuals ε_t and the shocks in time-varying coefficients ϑ_t, that is $\theta_i=\sigma^2/\sigma_i^2$. Hence it balances the fit of the model and the parameter stability.[13] Additionally, the time averages of the regression coefficients, estimated by such weighted least squares estimator, are identical to their GLS estimates of the corresponding regression with fixed coefficients, that is $\frac{1}{T}\sum_{t=1}^{T}\hat{a}_t=\hat{a}_{GLS}$.

Schlicht and Ludsteck (2006) show that this methodology gives statistically superior results in a linear state-space model estimated from small sample. First, it is a moment-based method that uses orthogonal parameterization instead of parameterization by initial values. Therefore, the initial values do not have to be estimated. Second, as the method uses all observations instead of only past information, it is *de facto* two-sided filter, while the Kalman filter is one-sided. Third, while the VC method coincides with the Kalman filter in large samples, the former method is statistically superior in poorly conditioned or small samples. The features of VC method make it feasible for our analysis. We deal with time-varying model, where coefficients are assumed to

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

[13] Originally, Schlicht and Ludsteck (2006) start with a derivation of 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.

follow a random walk, there is no *a priori* information about the initial values and the time series are rather short.[14]

We expect β_t to be positive, as the central bank is likely to react to the increase of inflation expectations by increasing its policy rate. Particularly, 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 the monetary policy regime or institutional constraints (Adam et al., 2005). The effect of inflation targeting adoption on β_t is ambiguous. As put forward by Kuttner and Posen (1999), β_t can either increase or 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 (such as inflation targeting).[15] However, the strength of the response under inflation targeting as compared to discretion depends on the credibility of the regime. A credible monetary policy does not have to react so strongly to inflation surprises, as inflation expectations are likely to remain anchored.

Similarly, ρ, a measure of interest rate smoothing, is expected to be positive with the values between zero and one. Many time-invariant estimates of monetary policy rules find such a high value of this parameter (0.7-0.9) that the degree of interest rate smoothing must be substantial. Rudebush (2006) claims that this parameter is clearly overestimated, given that the interest rate forecastability is limited. On the contrary, the time-varying model shall reveal more appropriate size of inertia given that it enables that some variables affect the interest rate setting in some periods but not in others and allows for more .

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 predictor of future inflation. In the latter case, the insignificant coefficient may suggest that the central bank is primarily focused on inflation developments and does not consider the output gap important in delivering low inflation.

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[14] The estimation of the second step is carried out by Schlicht´s package VC that uses the moment estimator. For comparison, we estimated all the results with parameterization by initial values. These results are available upon request.

[15] See King (1997) on how inflation targeting allows coming close to optimal state-contingent rule.

There is a certain debate whether other variables should be included in the monetary policy rule. This is especially appealing for small open economies that may be concerned with exchange rate fluctuations, as well as the developments 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 are present implicitly. For example, the exchange rate influences the inflation forecast to which inflation targeting central banks are likely to react. On the other hand, empirical studies often favor to include these variables in the estimated policy rule. Having these considerations in mind, we decided to include the exchange rate and foreign interest rates too, to assess whether these two variables carry any additional information to understand the interest rate setting in our sample countries.

3.2 The dataset

We use quarterly data. The sample period varies by countries (the UK 1975:1Q -2007:4Q, Australia 1972:4Q – 2007:4Q, Canada 1975:1Q – 2007:4Q, New Zealand 1985:1Q – 2008:2Q, Sweden 1982:2Q – 2007:3Q).

The dependent variables capturing the policy rate is the discount rate (3-month treasury bills) for the UK, intebank 3-month interest rate for Australia, 3-month treasury bills rate for Canada, overnight interbank 90 days interest rate for NZ and intebank 3-month interest rate for Sweden. We chose the interest rate so as to be closely linked to monetary policy, but also to be available for sufficiently long period. The foreign interest rate is the German 3-month EUROLIBOR for the UK and Sweden and the US 3-month interbank interest rate for Australia, NZ, and Canada. The inflation is measured as year-on-year change of CPI, besides the UK where we use RPIX (retail price index excluding mortgage interest payments) and the NZ where we use CPIX (CPI without interest payments). The output gap is taken as reported in the OECD Economic Outlook (production function method based on NAWRU - non-accelerating wages rate of unemployment) besides for NZ where this series is short and where we use the output gap derived from the Hodrick-Prescott filter applied at GDP series (constant prices, seasonally adjusted). The exchange rate is measured by the chain-linked nominal effective exchange rate (NEER) besides Canada where we use the bilateral exchange rate USD/CAD. For the regressions we use the deviation of the index from the HP trend (deviation from the sample mean has been used for robustness check).

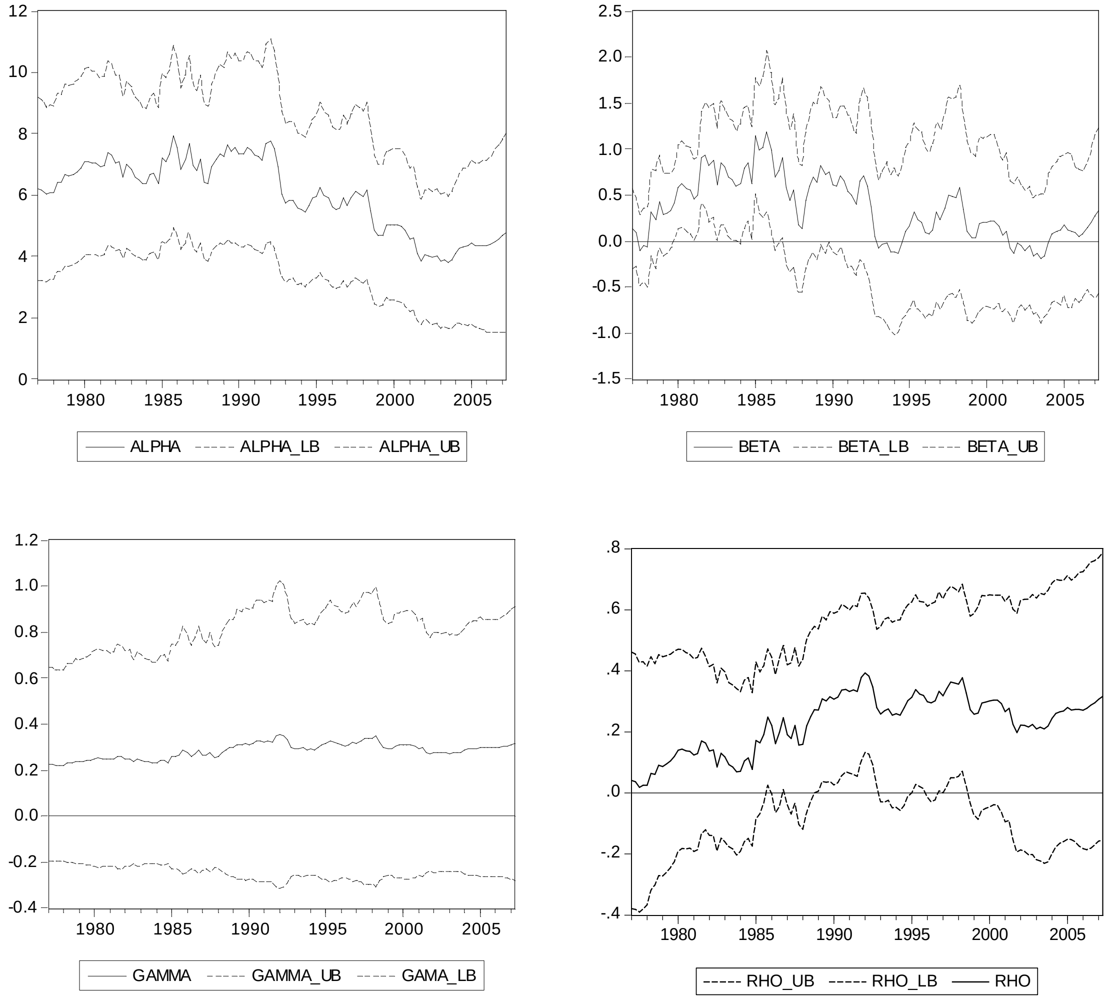
4 Results

4.1 United Kingdom

Our results show that the BoE significantly increased its response to inflation since late 1970's till mid-eighties. This overlaps with the Thatcher government and its major priority being the inflation control. The overall decline of the response since 1985 can be related to the dismissal of medium-term financial strategy (adopted just in 1979). We find that the response of interest rates on inflation was gradually decreasing during the 1990s in spite of the introduction of the IT. Although this finding can seem at the first sight contra-intuitive, it is important to keep in mind that, unlike in some emerging countries, the IT was not implemented in the UK as a strong anti-inflationist strategy. The inflation was already contained in the 1980's and very benign inflation enviroment was also supported by declining prices of raw materials on the world markets.

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

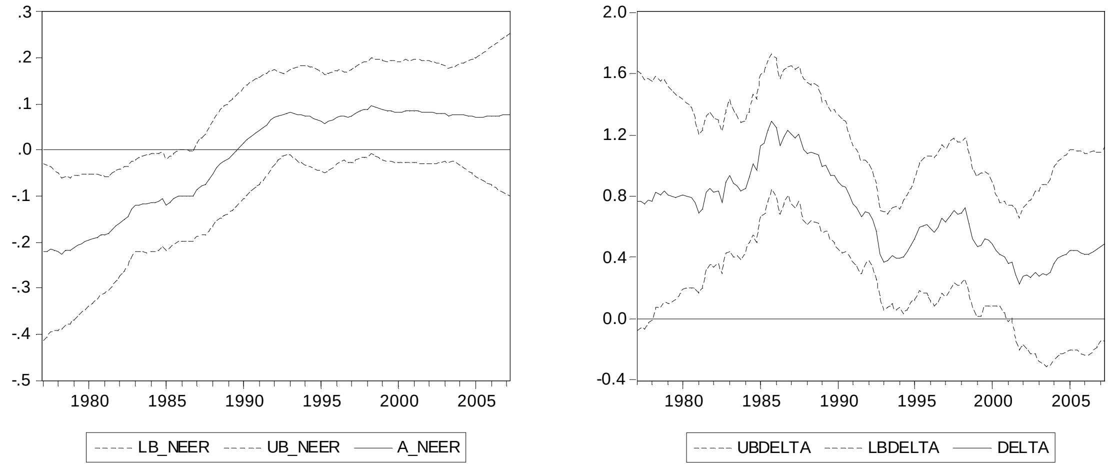
Figure 1 – Time-varying response coefficients in baseline (closed economy) policy rule, the U.K.



Note: 90% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. Upper left graph depicts the evolution of neutral rate. Upper right graph depicts the evolution of the response of interest rates to inflation. Lower left graph depicts the evolution of the response of interest rates to output gap. 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 albeit their importance was greater in the 1980s than recently. In particular, we find evidence that the BoE decreased its policy rate as the nominal effective exchange rate (NEER) strengthened during the 1980's even before the pound officially joined the ERM (1990). Yet, once the U.K. abandoned the ERM and introduced the IT, the BoE no longer seemed to explicitly react to exchange rate. Obviously, it has considered the exchange rate indirectly, as exchange rate fluctuations influence the inflation forecast (more on this see Taylor (2001)). The same applies to the response to foreign interest rate (Euribor). It was particularly strong during the 1980's and consequently its importance declined. Our results show little support for the hypothesis that the BoE follows the ECB as the estimated response of the coefficient declined and the confidence intervals widened after the launch of euro.

Figure 2 – Time-varying response coefficients in augmented (open economy) policy rule, the U.K.



Note: 90% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. Left graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (deviation from HP trend). Right graph depicts the evolution of the response of interest rate to foreign interest rate.

There are two directly comparable studies to our. Kishor (2008) obtains the results similar to ours in spite of using monthly data known to have slightly different dynamics. He finds that the anti-inflation stance was at peak in mid-eighties and since than was rather declining in spite of the IT adoption. Similarly, his finding that the response to foreign interest rate significantly declined since the ERM crises is complementary to our result that the BoE was giving much less consideration to the exchange rate evolution (NEER gap).

Trecroci and Vassalli (2009), who unlike Kishor (2008) and us do not correct for endogeneity in time-varying model, come to an opposite conclusion that the BoE response to inflation increased over time. Yet, other contra-intuitive results of their study point to possibility of endogeneity bias. First, the interest rate smoothing parameter takes on significantly negative values from 1980 till 1995. This would have implied not only that policy was not inertial but that there was actually a negative correlation between the present and past interest rate, which is inconsistent even in face of simple visual inspection of interest rate series. When we estimate our model without the endogeneity-correcting coefficients we obtain similar result (see Appendix, Figure A.1). Second, their coefficient of foreign (German) interest rate peaks in 1990 and is *de facto* invariant since than, which the authors interpret that as an implicit exchange-rate targeting. This finding is doubtful given the pound demise from the ERM and the IT implementation from 1992 onwards. In fact, British and German short-term that 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 the euro adoption.

4.2 New Zealand

The inflation targeting was introduced in New Zealand as the first country in the world by the Federal Bank Act signed in March 1990.[16] Our results indicate that the response of the RBNZ to the expected inflation was very close to unity during the whole sample period (1985-2007). In fact, it is clearly visible that the interest rate and inflation series commoves very closely. However, in Figure 3 we can also see that the official introduction of the IT does not seem to have represented a significant change in the interest rate setting (if anything there is very slight decrease of the response coefficient). Unlike in the UK, the response coefficient does not decrease substantially. This can be related to the fact that at the time the IT was introduced in New Zealand inflation rate was still relatively high (15%). Therefore, this policy was implemented in different context than e.g. in the U.K. where the inflation below double digit levels was already achieved during the 1980´s.[17] This result is together with the estimated insignificant response to the output gap consistent with the findings of time-invariant studies (Huang et al., 2001, Plantier and Scrimgeour, 2002) that the RBNZ applied a very strict version of inflation targeting. Finally, we find that the interest rate smoothing parameter is again rather modest.

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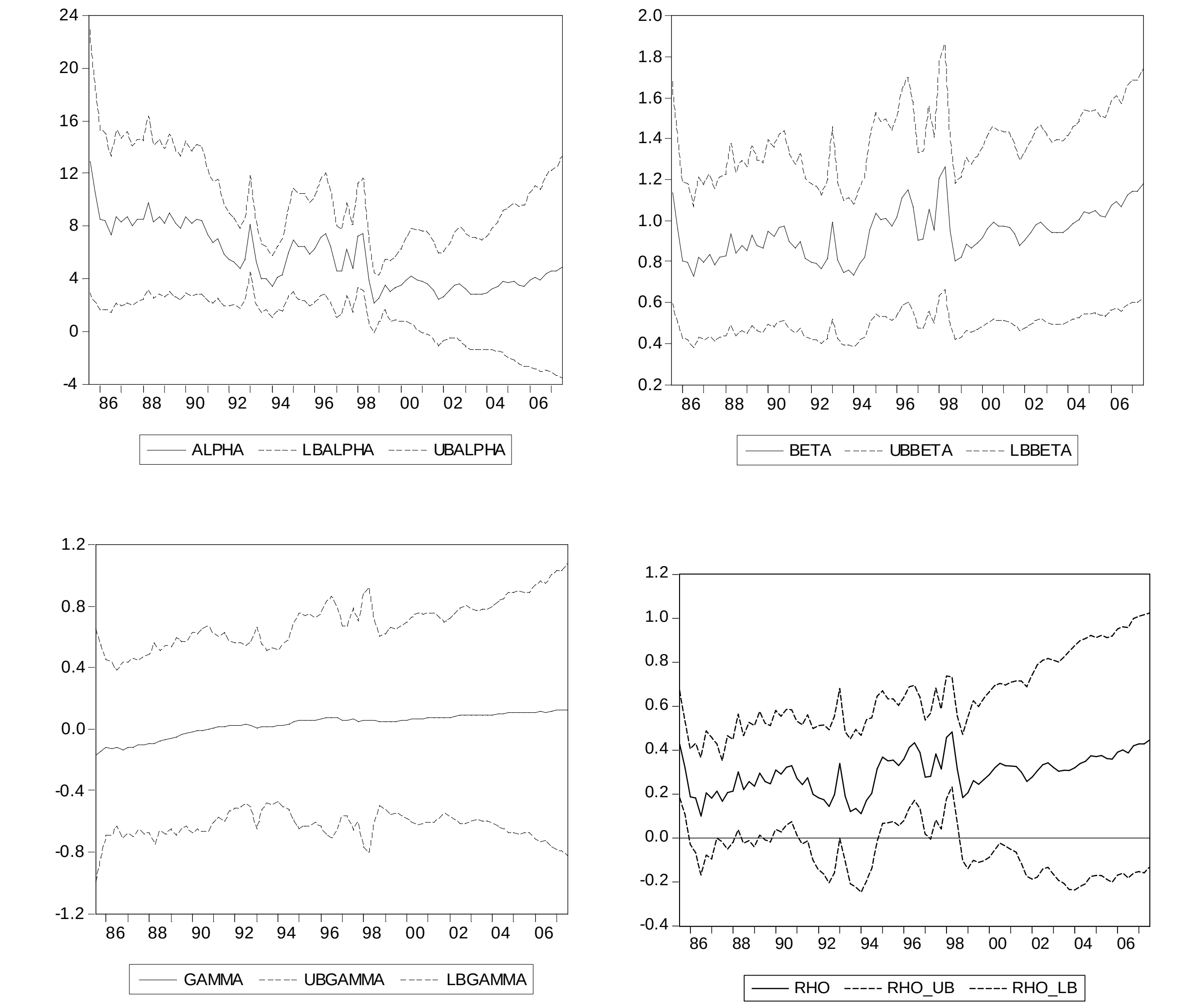
[16] Huang et al. (2001) argues that this policy was in effect since the end of 1988 when the RBNZ abandoned both monetary and exchange rate targeting. He also points to a specific feature of the monetary policy of the RBNZ that could be referred to as 'Open Mounth Operations'. Between 1989 and 1999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

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

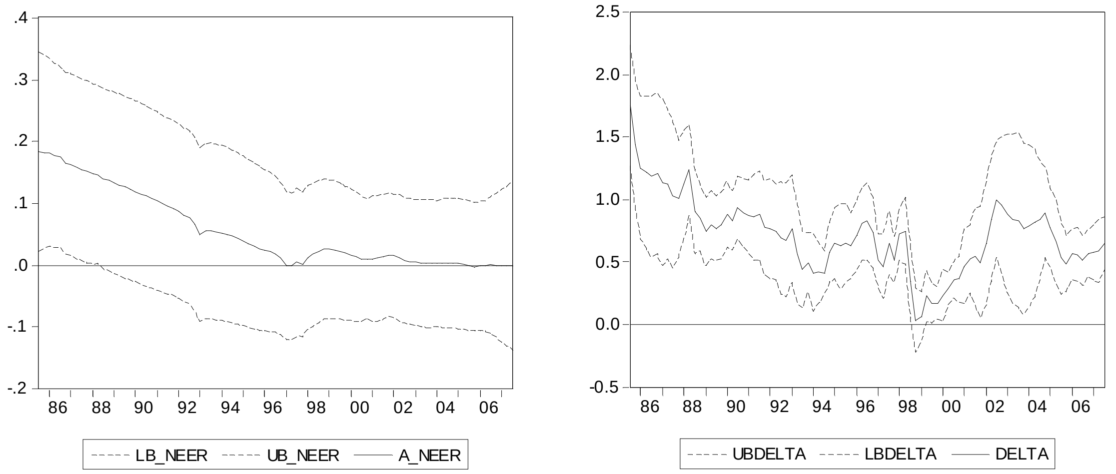
Figure 3 – Time-varying response coefficients in baseline (closed economy) policy rule, New Zealand



Note: 90% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. Upper left graph depicts the evolution of neutral rate. Upper right graph depicts the evolution of the response of interest rates to inflation. Lower left graph depicts the evolution of the response of interest rates to output gap. Lower right graph depicts the evolution of the interest rate smoothing parameter.

In the augmented model we test whether external variables have any direct effect on interest rate setting in New Zealand. The evidence for the exchange rate is not conclusive. We find positive response to NEER, which is rather contra-intuitive in terms of the Taylor rule. However, the coefficient is significant only before the introduction of the IT and its positive sign is likely related to currency appreciation following the interest rate increase. In that period the RBNZ aimed to keep the exchange rate in defined range and the interest rate was likely set so as to influence the exchange rate *ex-ante* (rather than *ex-post* respond to its movements). Consistently with this finding, Ftiti (2008) rejects in a time-invariant model that the RBNZ responded to exchange rate. On the other hand, we find some evidence in favor of consideration of foreign interest rate, yet its response coefficient generally decreased since the IT launch.

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



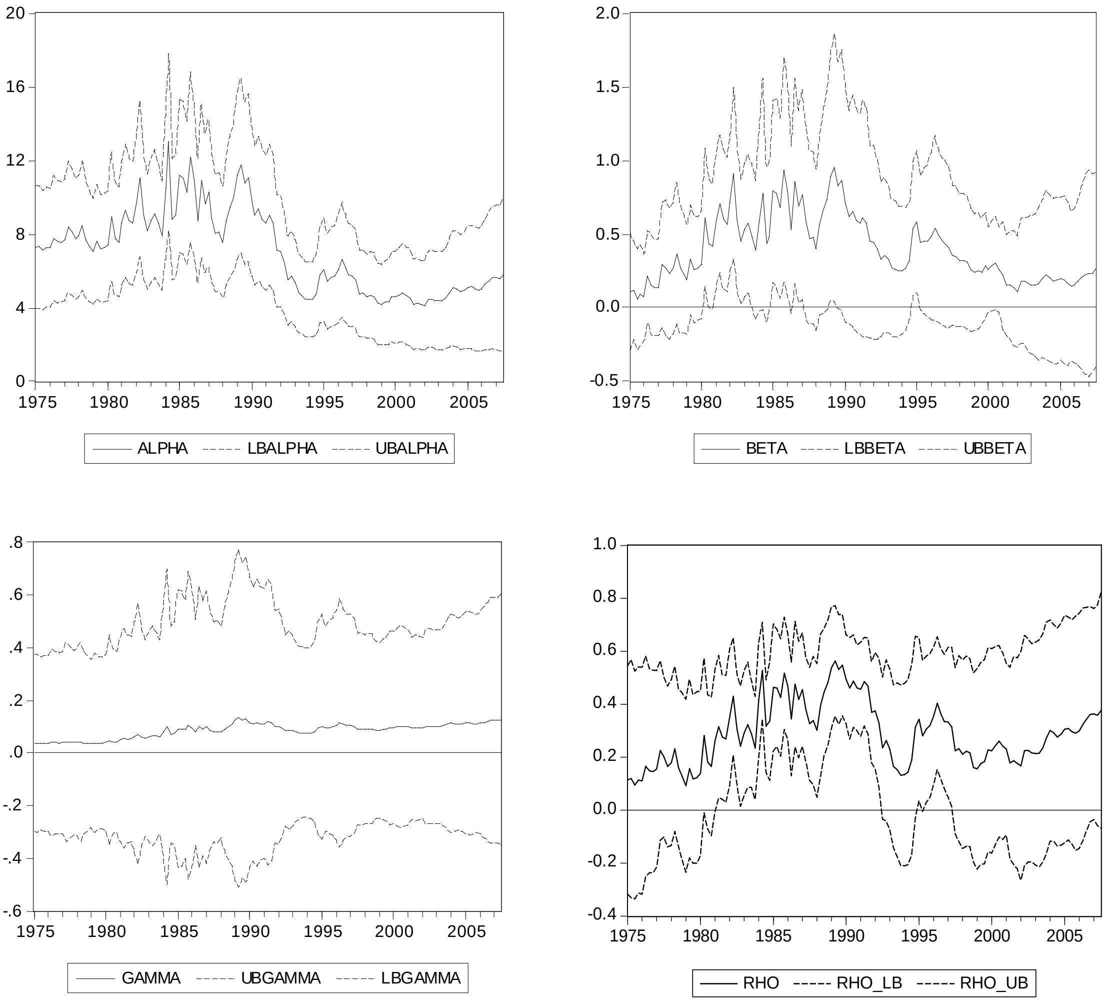
Note: 90% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. Left graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (deviation from HP trend). Right graph depicts the evolution of the response of interest rate to foreign interest rate.

4.3 Australia

Our results on Australia are available in Figure 5 and 6. The response of interest rate to inflation is strongest in the 1980s, which is very similar to the U.K. experience This period was characterized by inflation rates around 10% and the central bankers had to be quite aggressive in interest rate setting in order to break the record of high inflation built also deeply in the public expectations. Neither the monetary targeting (employed until 1984) nor checklist approach (1985-1990) seemed to be successful in this regard. The fluctuation of the inflation response coefficient point to discretionary nature of policy decisions (this finding is consistent with Leu and Sheen, 2006). The response coefficient peaks in 1990 on the eve of the IT but declines after adoption of this regime. It is again arguable whether it was the credibility of this regime what anchored the inflation expectations and allowed the RBA to behave less aggressively. The original inflation decline could be also related to the world recession in early 1990s. Our results dispute the finding of De Brouwer and Gordon (2005) arguing that the inflation response of the RBA increased as a result of inflation targeting launch. This result seems to be driven by the transmission of increased smoothing parameter (rho) to inflation response coefficient (beta) that is problem inherent to time-invariant model.

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

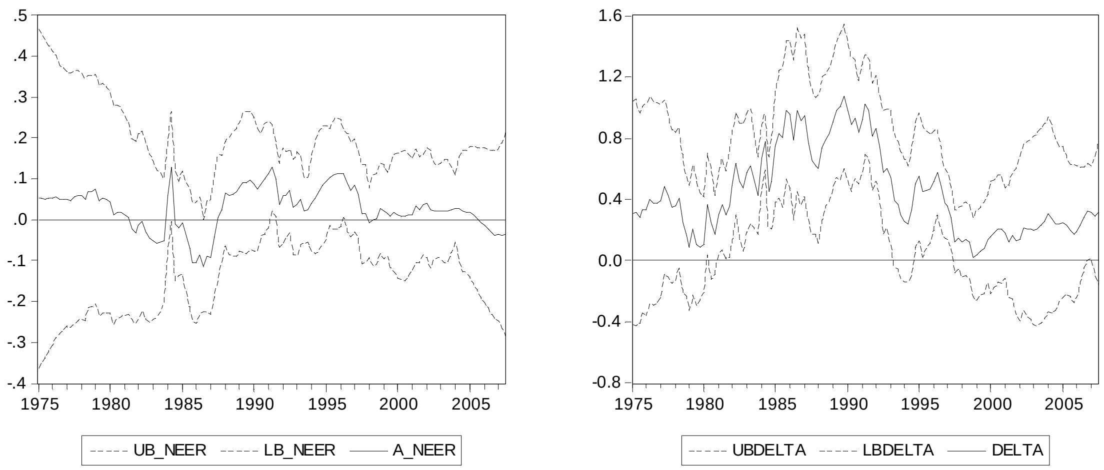
Figure 5 – Time-varying response coefficients in baseline (closed economy) policy rule, Australia



Note: 90% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. Upper left graph depicts the evolution of neutral rate. Upper right graph depicts the evolution of the response of interest rates to inflation. Lower left graph depicts the evolution of the response of interest rates to output gap. 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 irrespective which definition of the exchange rate we use (NEER or TWI) besides the period of 1985-87 when the currency depreciation (Australian dollar was allowed to float in 1983 after period of moving peg vis-à-vis TWI) was compensated by interest rate increase so as to curb the inflation pressures (see Greenville, 1997). Foreign interest rate parameter is estimated always positive, albeit it is significant only in the 1990s and its consideration fades back as the IT was introduced. Since 2001 the Australian and the US interest rates diverge and the response coefficient approach zero. This may be related to idiosyncratic developments in the U.S. when the FED lowered interest rate so as to face the fear of recession following September 2001 terrorist attack.

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



Note: 90% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. Left graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (deviation from HP trend). Right graph depicts the evolution of the response of interest rate to foreign interest rate.

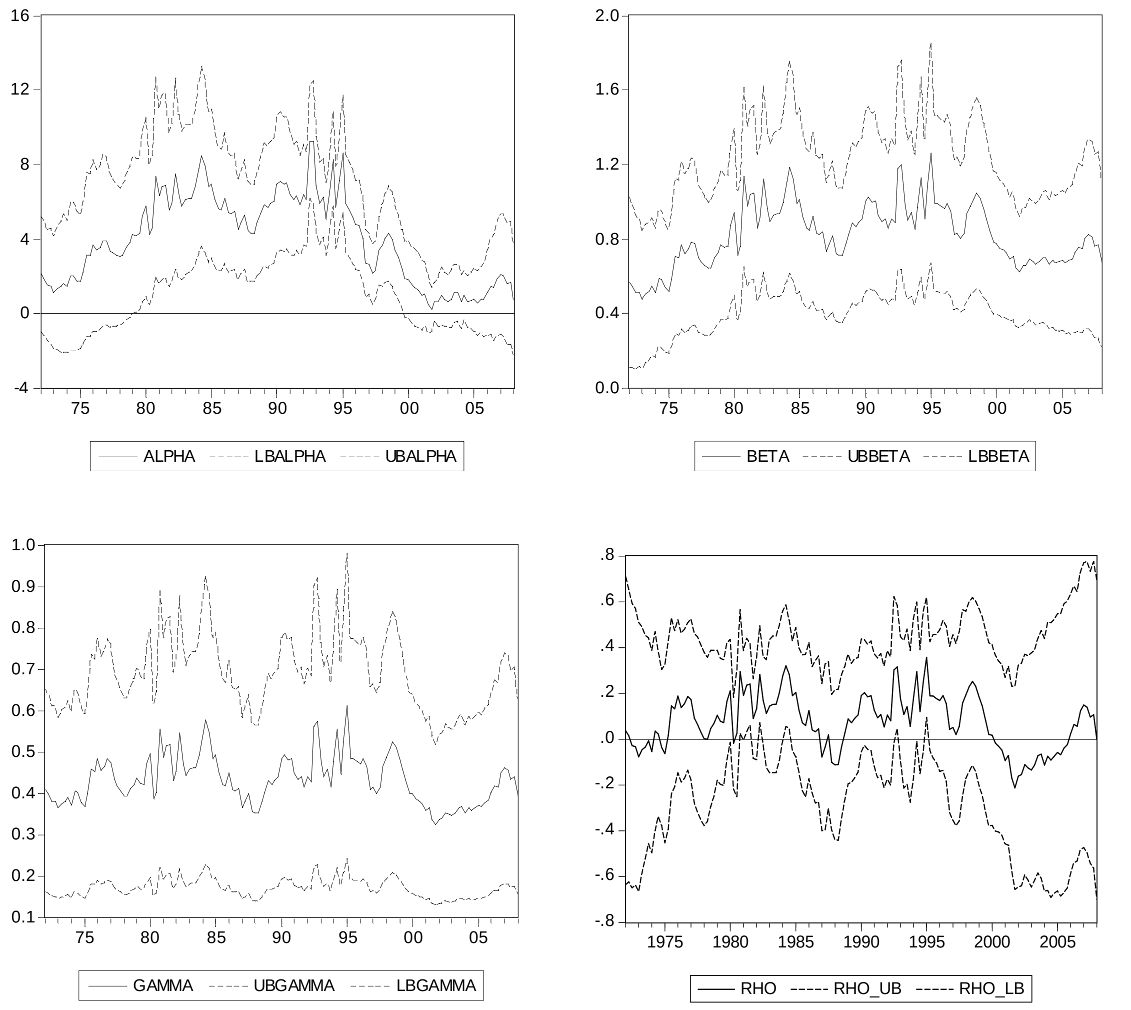
4.4 Canada

The monetary policy rules estimates for Canada are presented in Figure 7 and 8. The response of interest rate to inflation peaks in the first half of 1980s and mid-1990s. The former period was characterized by relatively high inflation rates that unquestionably drove rather aggressive policy of the BoC similarly as in the U.K. or Australia. It is arguable whether the original inflation rate was a consequence of accommodative policy of monetary targeting applied between 1978 and 1982 (see Figure 7). Unlike in the U.K. or Australia, the inflation response coefficient retains its magnitude after the IT adoption and decreases only in the last decade (due to almost negligible inflation rates).[18]

[18] The BoC also reported the monetary condition index (MCI) as a compound of policy instrument (interest rate) and exchange rate. MCI accompanies the proposed of Ball (1999) to target long-term inflation, i.g. inflation rate adjusted for the transitory effect of exchange rate on import prices. However, there is no indication that BoC actually ever used MCI for practical policy making and it ceased to publish it in 2006.

The response to the output gap is significant and almost invariant in time (as in the UK) confirming long-term preference of the BoC to smooth the 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

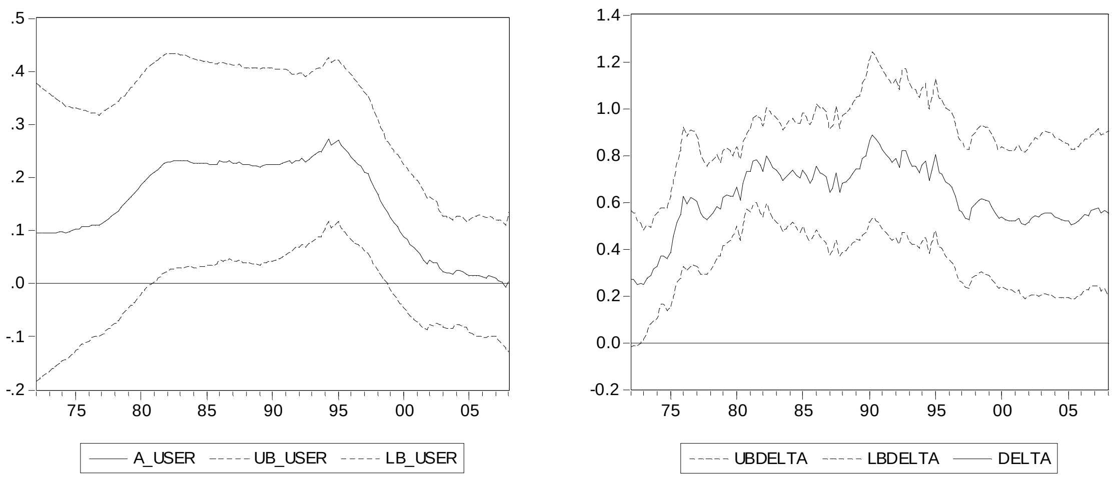


Note: 90% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. Upper left graph depicts the evolution of neutral rate. Upper right graph depicts the evolution of the response of interest rates to inflation. Lower left graph depicts the evolution of the response of interest rates to output gap. Lower right graph depicts the evolution of the interest rate smoothing parameter.

The dependence of Canadian monetary policy on external factors, in particular the developments in the U.S., is confirmed in model augmented by the exchange rate and foreign interest rate. The response to the exchange rate is positive but almost insignificant and dissipates in the last decade.

On the other hand, a response to the US interest rate dynamics has been substantial for the whole period of analysis until the present days.

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



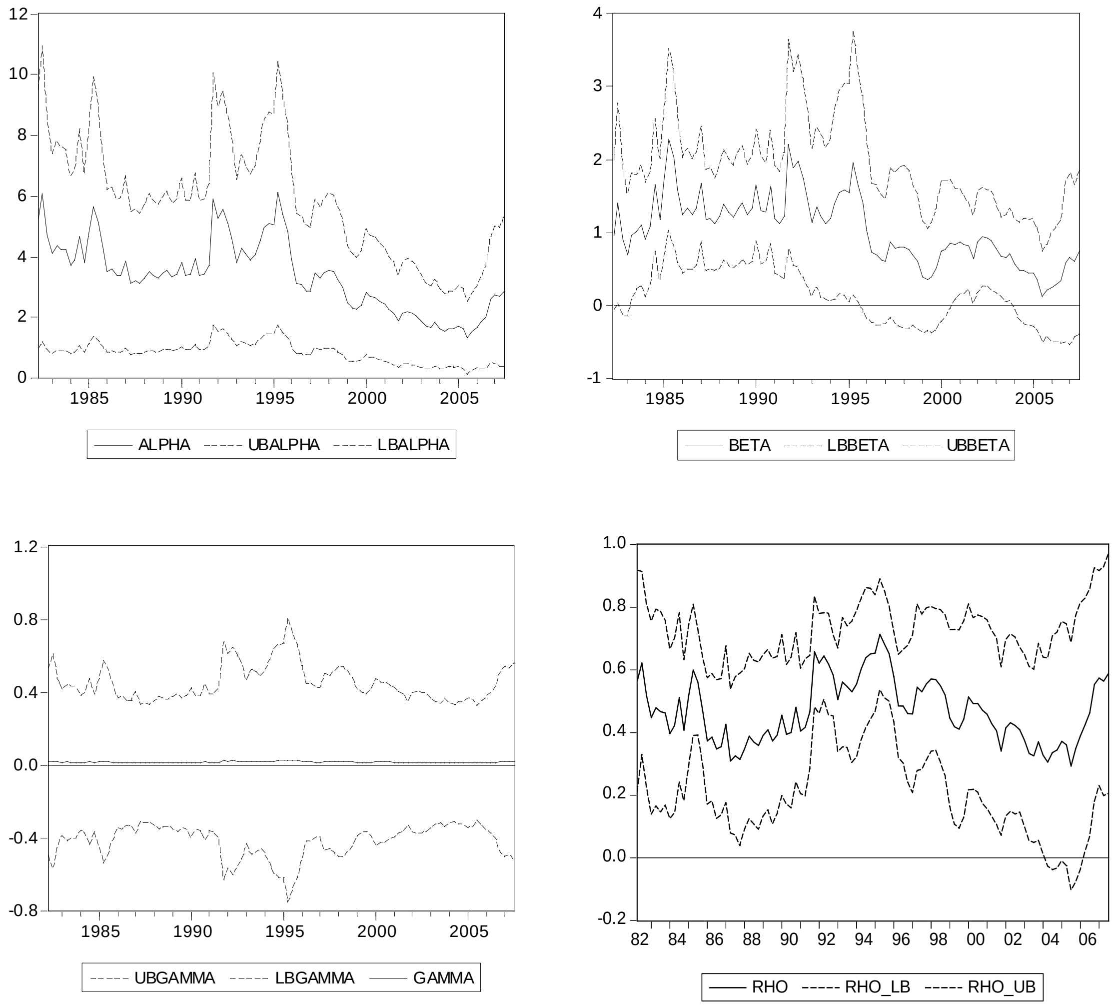
Note: 90% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. Left graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (deviation from HP trend). Right graph depicts the evolution of the response of interest rate to 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 the IT. This complies with Berg et al. (2004), who argue that the introductory phase of the IT in Sweden was characterized by building credibility of the new regime. The decline of neutral rate reflects the prevailing low inflation environment in Sweden from mid-1990s onwards. The time-varying coefficient on interest rate smoothing is estimated to be larger in Sweden than in case of the U.K. (and other countries). We hypothesize that this may reflect the differences in the governance of these two central banks. While the Sveriges Riksbank is known to achieve its monetary policy decision in rather collegial way, the Bank of England monetary policy committee reaches decisions rather individualistically (Blinder, 2007). As a result, Sveriges Riksbank is likely to smooth its interest rates to a greater degree.[19]

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

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

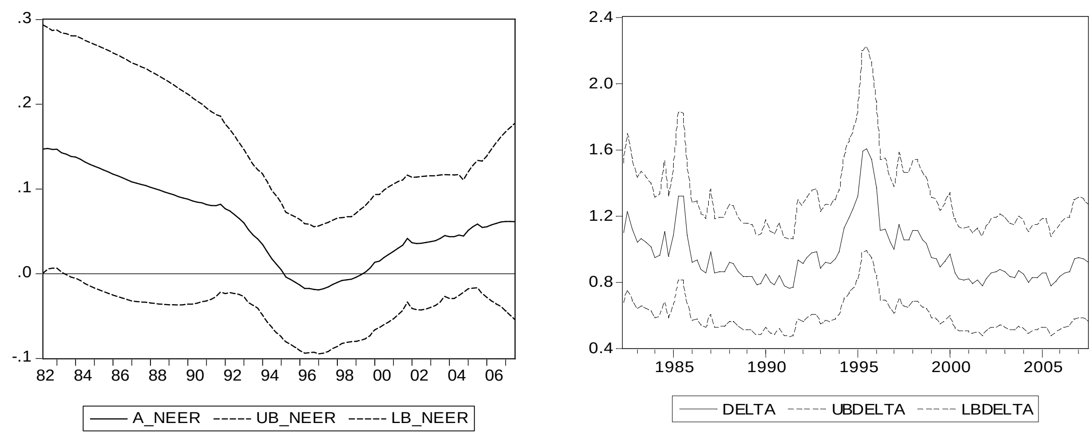


Note: 90% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. Upper left graph depicts the evolution of neutral rate. Upper right graph depicts the evolution of the response of interest rates to inflation. Lower left graph depicts the evolution of the response of interest rates to output gap. Lower right graph depicts the evolution of the interest rate smoothing parameter.

Our results imply a prominent role for the external factors rather than the output gap for determination of Swedish monetary policy. In particular the coefficient on foreign interest rate (Eurolibor) is sizeable during the whole sample period, which is rather interesting given that Swedish monetary policy has not officially been subject to any external constrain (at least since krone's departure 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 in overall consistent with the surveyed time-invariant studies pointing to predominant role of inflation forecast (Jansson and Vredin, 2003) but also 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: 90% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. Left graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (deviation from HP trend). Right graph depicts the evolution of the response of interest rate to 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 the IT to the hypothesis that this monetary framework has a positive effect on the inflation expectations of economic agents. If the expected inflation is low, the monetary policy does not need to 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 inflation persistence (the dependence of current inflation on its first lagged value) fades away under credible policy regime (such as the IT).

In order to shed some light on this issue, we have used our estimation framework and fitted the AR(1) model with drift to inflation series, allowing the coefficient on the lagged inflation to be time-varying. Our results indicate (as reported in Appendix, Figure A.2) that inflation persistence decreased over time for all the countries. Moreover, it is very notable that the persistence fell down especially during the 1990's as the IT was introduced. This finding is confirmed when we, in the spirit of the backward-looking Phillips curve, include the lagged output gap as a forcing variable.[20]

In general, our results are consistent with Benati (2008) who performs sub-sample analysis under different policy regimes. Unlike him, we do not need to impose breaks in inflation process at any particular date but simply observe whether and when such breaks occur. Our findings does not exclude the possibility that the inflation persistence decreased because of other factors (the "good luck" hypothesis), but the temporal coincidence between the introduction of the IT and the significant decrease of inflation persistence in several countries make an important case for the "good policy" hypothesis. Taking the example of the U.K., we can see that the inflation (rate) moderation goes back to 1980's when we still observe rather high inflation persistence, in spite of a very aggressive anti-inflationary (yet discretionally) policy.

5 Concluding Remarks

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

In our view, results point to the usefulness of this framework. Estimating standard monetary policy rules for our sample countries shows that the policy changes gradually and that these changes coincide with several important institutional reforms, as well as with the periods when central banks successfully decreased double-digit inflation rates into a low inflation environment.

Our results indicate that the interest rate smoothing is much lower than what time-invariant estimates of monetary policy rules typically report (see for example, Clarida et al., 1998, 2000). Our estimates seem to be reasonable, given the recent critique of Rudebush (2006), who argues that the degree of interest rate smoothing is rather low.

We find that external factors matter for interest rate setting in all our sample countries. To be more precise, 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.

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

In this respect, our results suggest that the response of interest rates to inflation is particularly high during the periods when central bankers want to break the record of high inflation such as in the U.K. in early 1980s. Contrary to common wisdom, the response is often found to be less aggressive after the adoption of inflation targeting, suggesting the positive effect of this regime on anchoring inflation expectations. In other words, the monetary policy needs not be as aggressive as under discretionary regime in order to achieve price stability. This result is supported by our finding that inflation becomes less inertial and policy neutral rate decreases after the adoption of inflation targeting.

In terms of future research, we believe that it would be worthwhile to apply this framework to understand better whether and how monetary policy reacts to the 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 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 the equation (A.1) is replaced by a system

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

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

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 following matrices:

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

of order T × Tn (T-1)n × Tn

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

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 the 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 explicit definition of a set of possible values of a conditioned by matrix *P* and random variable *v*. Following the equation (A.5) any solution a can be written as

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

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

Hence equation (A.5) can 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.

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

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

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 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 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[21].

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 $\hat{a}_t$.

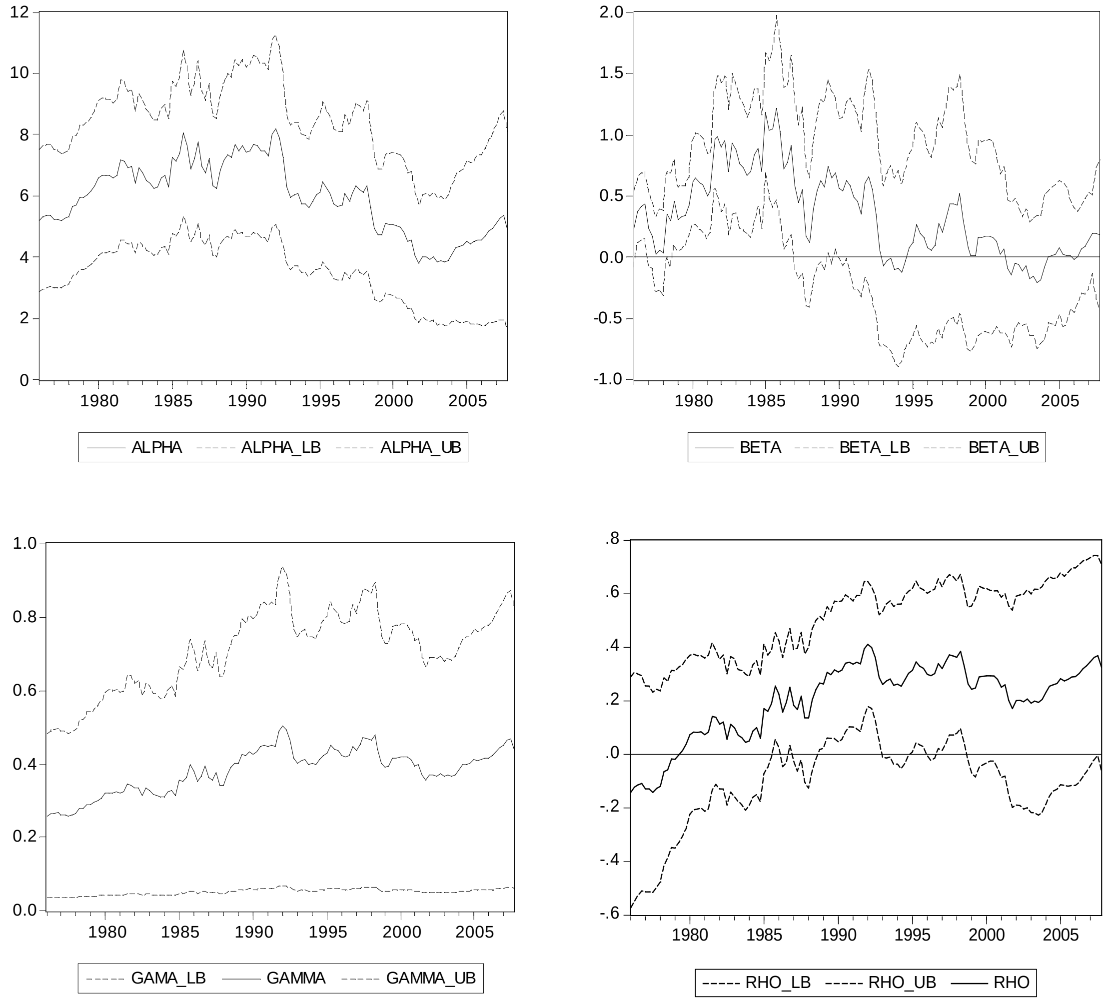
[21] Originally, Schlicht and Ludsteck (2006) start with a derivation of maximum likelihood estimator of parameters *a* based on the idea of orthogonal parameterization, which is described in the Appendix. Then they prove that the weighted least squares estimator is identical to the maximum likelihood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

The estimation procedure proceeds as follows. The iterative procedure has two steps. First, given variances of residuals in both equations in (15), σ^2 and σ_i^2, the coefficients at 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 Estimates not dealing with endogeneity in policy rules: An Example

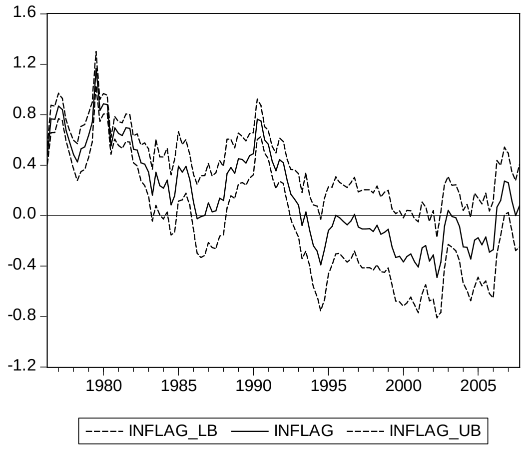
Figure A.1 – Time-varying response coefficients in baseline (closed economy) policy rule, the U.K.



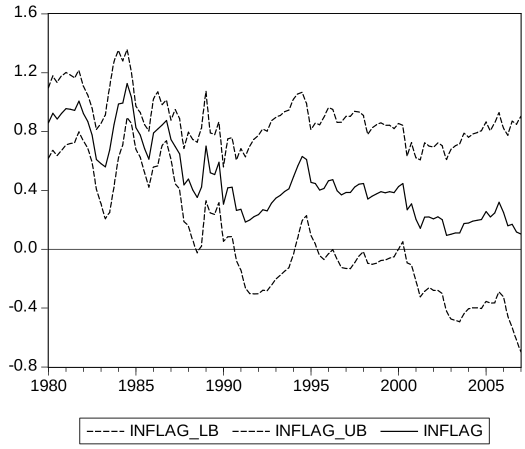
Note: 90% confidence bands; model without bias correction terms, i.e. not dealing with endogeneity in monetary policy rules.

Figure A.2 – Time-varying response coefficients in AR(1) model for inflation

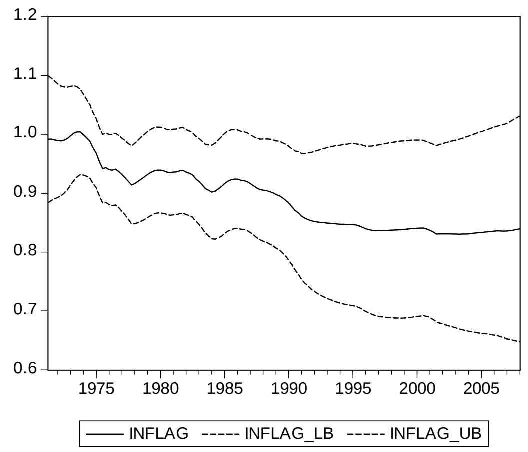
U.K.



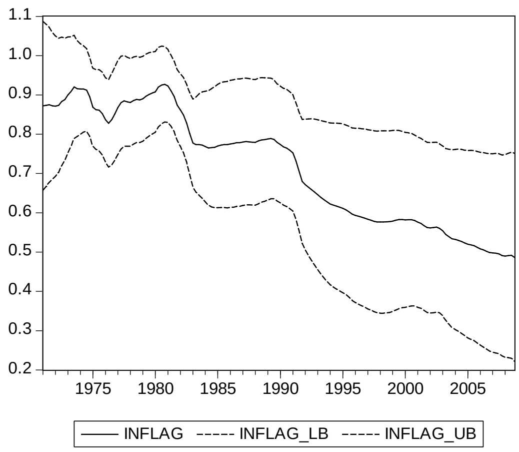
New Zealand



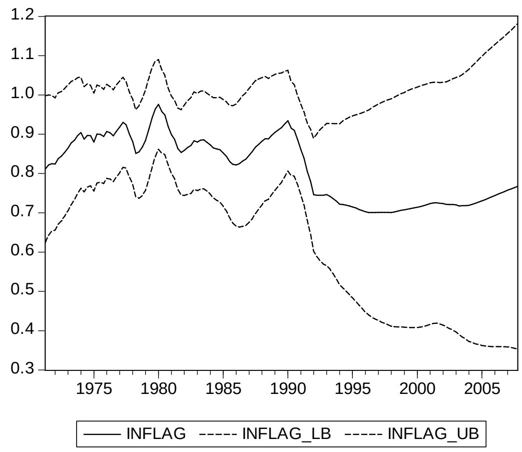
Australia



Canada



Sweden



Note: 90% confidence bands; the coefficient of AR(1) term (i.e. model $\pi_t = \alpha + \pi_{t-1}$) for inflation series employed in previous analysis for each country.