Universitat Autònoma de Barcelona
Departament d'Economia Aplicada

Empirical Essays on Monetary Policy

PhD Dissertation

by

Bořek Vašíček

Supervisor:

Prof. Luca Gambetti
Departament d'Economia e Historia Economica
Universitat Autònoma de Barcelona

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Danečkovi

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Introduction

This dissertation is divided into four essays, each of them having its own structure and methodological framework. Although each of the essays making the chapters of the thesis is self-contained, their topics are very closely related. Consequently, the reader will be able to follow the thesis in its unity.

The four essays are empirical and address some relevant issues from monetary economics and policy. In first three essays we aim at the way central banks sets their policy rates, which is usually described by central bank's reaction function or a monetary policy rule (Taylor, 1993), and in the fourth one, we look at the inflation dynamics and the New Keynesian (NK) Phillips Curve (Galí and Gertler, 1999). In the first three papers, we look at different issues related to policy rules such as their nonlinearities, evolution across time, the intensity of interest rate response to inflation, degree of policy inertia as well as how policy changes when it is faced by financial stress.

The monetary policy rule and the NK Phillips Curve became key elements of the NK policy model (e.g. Galí, 2008), which is at present the most influential theoretic framework for analysis of macroeconomic dynamics and monetary policy both in academia and among monetary policy makers. Dynamics Stochastic General Equlibrium models (DSGE), which are the main analytic tools within most central banks (e.g. Area-Wide model of the European Central Bank), are inspired by the NK models. While the traditional Keynesian literature assumed sticky prices and wages in the short-term, it was unable to provide their satisfactory microeconomic explanations. The NK model enriched the original Keynesian literature by microeconomic foundations but at the same time by some important elements such as rational expectations and vertical long-term Phillips curve incompatible with the Keynesian tradition. Price rigidities (both nominal and real), which are result of imperfect competition, enable that monetary policy has impact on real variables in the short term. Such view on monetary policy is quite paradoxically close to monetarism but quite distant from other current macroeconomic branches such as Real Business Cycle Theory or Post-Keynesianism that cast doubt on the ability of monetary policy to affect real economy (Snowdon and Vane, 2005). In spite of the popularity of NK models, the empirical evidence goes behind the theory. The purpose of this thesis is to contribute to the empirics of the NK model.

In each chapter we look at several countries in order to pursue comparative analysis. The reason for this choice is my conviction that empirical macroeconomic analysis can provide reliable results only when one is not limited by data of a particular country and period. Consequently, a comparative study faces a trade-off between the choice of model, method and data that are the most suitable for a particular country, on one hand, and the need to take homogeneous stance that allows comparison of the results, on the other. Cross-country analysis also forces the author to avoid subtle methodological adjustments to achieve “the best” or the most interesting results for a particular country. Given that the thesis is empirical, the methodology of each essay consists in use of a given macroeconomic model, specification of empirical relation that will become subject of analysis, selection of corresponding time series and an application of suitable econometric methods. In general, we look at two kinds of countries: (i) developed economies such as the US or the UK where not only monetary policy has a long-term tradition but also the data are of good quality and the time series cover several decades and (ii) post-transitional economies of new Central-European EU member states where the period for reasonable econometric analysis is substantially limited both by the process of economic transition itself but also by the data availability.

Chapter 1, *How Does Monetary Policy Change? Evidence on Inflation Targeting Countries* is a joint work with J. Baxa and R. Horváth. It examines how the monetary policy setting evolves in time. We aim at a group of inflation targeting (IT) countries (Australia, Canada, New Zealand, Sweden and the United Kingdom) because it seems especially interesting how the de-facto interest rate setting changes once this policy framework is introduced as well as whether the operative monetary policy under this regime is driven by some rule or is rather discretional.

The IT regime gained popularity as a monetary policy framework World and numerous academic studies suggest the gradual convergence of the inflation rate to its target as an optimal monetary policy (Clarida et al., 1999, 2000; Batini and Haldane, 1999). The IT feasibility has been confirmed by the experience of many countries that adopted this framework. However, there is still ongoing discussion what part of the decrease of inflation rate can be attributed to IT and the empirical evidence about the effect of this regime on other variables is mixed. Bernanke et al. (1999) assert that IT does not decrease the cost of disinflation and does not bring credibility to monetary policy immediately. The inflation expectations decrease only gradually over time. Ball (1999) concludes that IT is an efficient policy with respect to inflation and output variance, but it is not superior to simple interest-rate rules. Ball and Sheridan (2003) find that while the economic performance (inflation, output and interest rate variability) improved over time in OECD countries, there is no significant difference between countries pursuing IT and those using other policy arrangements. Moreno and Rey (2006) studying the inflation dynamics in various European countries found that IT had positive impact on the decrease of low frequency dynamic (inflation trend) because it decreases the size of monetary shocks.

In terms of discussion whether IT is rule-based or discretional, it is interesting to recall Bernanke's claim (2003) that most major central banks in fact use constrained discretion. The word "constrained" reflects situation when central banks have explicit commitment of keeping the inflation low and stable. "Discretion" does not consist only in the freedom to choose the operative policy for achievement of the price stability but also conditioned on the fulfillment of the previous objective, the central bank can pursue other targets, in particular cycle stabilization. Svensson (2003) adverts to certain inconsistency between the need to follow policy rule from the commitment perspective and the recommendation not to apply the simple rules mechanically when no specification is provided what is the permitted deviation from the rule.

To examine the change in monetary policy rules across time, we use time-varying parameter model with endogenous regressors with aim to reveal if and how the interest rate setting changed as a consequence of adoption of IT regime without imposing ex-ante any structural breaks. To estimate this model we use novel moment-based estimator (Schlicht and Ludsteck, 2006) that has some desirable computational properties over the common Kalman filter.

The findings of this chapter can be resumed as follows. First, monetary policy rules change rather gradually than abruptly, which points to the importance of applying time-varying estimation models as opposed to switching models. Second, the interest rate smoothing parameter is much lower than what previous time-invariant estimates report. Therefore, the puzzle of high policy inertia found in many empirical studies, which is inconsistent with practical unpredictability of interest rate changes in middle-term (Rudebusch, 2006), seems to be driven by omission of time-variance of response coefficients. Third, the external factors matter most IT countries, albeit the importance of exchange rate diminishes after the IT adoption. Fourth, the response of interest rates to inflation is particularly strong during the periods when central bankers want to break the record of high inflation such as in the U.K. at the beginning of 1980s. Contrary to common wisdom, the response becomes less aggressive after the IT adoption suggesting the positive effect of this regime on anchoring inflation expectations. This result is supported by our complementary finding that inflation persistence as well as policy neutral rate typically decreased after the adoption of inflation targeting.

Chapter 2, *Monetary Policy Rules and Financial Stress:Does Financial Instability Matter for Monetary Policy?* is a joint work with J. Baxa and R. Horváth. It extends the previous chapter examining whether and how main central banks (Australia, Canada, Sweden, the United Kingdom and the US) responded to the episodes of financial stress over the last three decades. Given that the interest rate response to financial stress must be nonlinear and only temporal, the time-varying coefficient model again seems to be the most suitable framework. This methodology accompanied by new financial stress dataset developed by the International Monetary Fund (IMF) allows not only testing whether the central banks responded to financial stress at all but also detecting the periods and type of stress that were for monetary authorities the most worrying.

The financial frictions such as unequal access to credits or debt collateralization were recognized to have important consequences for the transmission of monetary policy as they increase the response of aggregate demand to interest rate changes. Already Fisher (1933) presented the idea that adverse credit-market conditions can cause significant macroeconomic disequilibria. Kiyotaki and Moore (1997) showed that the credit markets are characterized by asymmetric information and principal-agent relationships when the durable assets are not only used in production process but they are often used as collateral for loans. The role of credit market frictions was incorporated to NK model by Bernanke et al. (1999). Their "financial accelerator" mechanism consists in the relation between net worth of borrows and the finance premium they have to pay. This relationship makes that any shock affecting the net worth of borrowers has accelerated affects on aggregate demand and output.

In spite of the potential effects the financial distress can have on real economy, the relation between monetary policy stance and financial stability is not straightforward. The financial stability is mainly not an explicit objective of monetary policy but it is rather retained by separate supervisory function of central banks. Although many DSGE models already include financial sector and model potential financial frictions (Tovar, 2009), the issues of viability to modify monetary policy stance in face of financial stress (Cúrdia and Woodford, 2010) in general, or asset price misalignments in particular (Bernanke and Getler, 1999) are still subject to debate. Moreover, there are other puzzles such as how to measure the financial stress or whether the central banks in reality adjust the policy rates when financial markets are fragile (Borio and Lowe, 2004; Ceccheti and Li, 2008).

The IMF financial stress index (Cardarelli et al., 2009) is a comprehensive stress measure that tracks the overall financial stress alike the distress in different markets (banking sector, stock market and exchange rate market). Combined with time-varying model, we aim to quantify the intensity of the interest rate response to the financial stress. Our findings suggest that central banks usually loosen monetary policy in the face of high financial stress. We argue that all central banks kept its policy rates during the crisis by about 50 basis points on average (250 pb in case of the Bank of England) lower solely due to their response to financial stress (besides the decrease justified by developments in inflation expectations or the output gap). The current global turmoil stands out as an exceptional period both in term of the intensity of the response but also its coincidence for different central banks. There is a certain cross-country and time heterogeneity when we look at central banks' considerations of specific types of financial stress. While most central banks seem to respond to stock market stress and bank stress, exchange rate stress is found to drive the reaction of central banks only in more open economies.

Chapter 3, *Is Monetary Policy in New Member States Asymmetric?* extends the analysis of policy rules towards emerging market economies. In particular, we deal with potential asymmetries in monetary policy conduct of three new EU members (the Czech Republic, Hungary and Poland) that apply inflation targeting that is prone to policy asymmetries.

The empirical studies usually assume that monetary policy responds symmetrically to economic developments, in other words the monetary policy rule is modeled as a linear function. A theoretical underpinning of a linear policy rule is the common linear-quadratic (LQ) representation of macroeconomic models (economic structure is assumed to be linear and the policy objectives to be symmetric, in particular the loss function is quadratic). However, when the assumptions of the LQ framework are relaxed, the optimal monetary policy can be asymmetric, which can be represented by a nonlinear monetary policy rule. For example, a convex Phillips curve implies that the inflationary effects of excess demand are larger than the disinflationary effects of excess supply (e.g. Laxton et al., 1999). Such situation can lead optimizing central bankers to behave asymmetrically (Dolado et al., 2005). On the other hand, asymmetric monetary policy can be related by asymmetric preferences of central bankers. For example, reputation reasons can drive central banks, especially those pursuing inflation-targeting, having an anti-inflation bias and respond more severely when inflation is high or exceeds its target value (Ruge-Murcia, 2004).

To test whether the monetary policy can be described in the three new EU member states as asymmetric, we apply two different empirical frameworks: (i) GMM estimation of models that allow discriminating between the sources of policy asymmetry but are conditioned by specific underlying relations (Dolado et al., 2005: Surico, 2007); and (ii) a flexible framework of sample splitting where nonlinearity enters via threshold variables and monetary policy is allowed to switch between regimes (Hansen, 2000; Caner and Hansen, 2004).

We find little evidence in favor of nonlinearities in economic system, in particular convexity of AS schedule, and therefore no reason for asymmetric policy. On the other hand, there is some evidence of asymmetric policy driven by central bank preferences in terms of inflation rate and the level of interest rate. Unfortunately, the validity of empirical results is conditioned by specific parametric models underlying the estimated nonlinear equations. Therefore, we use as an alternative a method of sample splitting where nonlinearities enter via a threshold variable and monetary policy is allowed to switch between two regimes. Besides the inflation and output gaps we used a financial stress index as competing threshold variables. The threshold effects are most evident with financial stress index. While the Czech and Polish National Banks seem to face the financial stress by decreasing their policy rates, the opposite pattern is found in Hungary.

Chapter 4, *Inflation dynamics and the New Keynesian Phillips curve in Central European countries* seeks to shed light on the inflation dynamics of four new EU members (Czech Republic, Hungary, Poland and Slovakia) and test whether the predominant theoretical model of inflation, the New Keynesian Philips Curve (NKPC), finds support in the data of these countries.

Understanding the nature of the inflation dynamics is very important for the implementation of monetary policy. Unlike the traditional PC that was a purely empirical model, the NKPC, which appeared in the 1990's, stands on elaborated micro-foundations of monopolistic competition and price rigidities (Taylor, 1980; Calvo, 1983). The NKPC links the current price inflation to inflation expectations of economic agents. However, as the data suggest that inflation is persistent process, several attempts were made to give inflation persistence some structural base (Galí and Gertler, 1999; Christiano et al., 2005). The current discussion on the inflation dynamics in general and the NKPC in particular aims at two main issues: (i) whether inflation is mainly backward-looking or forward-looking phenomenon and (ii) what is the principal inflation-forcing variable in the short-term. These characteristics of inflation have important consequences for monetary policy. In particular, if inflation is predominantly forward-looking phenomenon and its dependence on the past (intrinsic persistence) is limited, a credible monetary policy can achieve disinflation at no cost (in terms of real output loss).

Inflation in the NMS has some specific features as compared to developed countries. During most of the 1990s it has been affected by price liberalization. Therefore, inflation rates have been comparatively higher. Moreover, all four countries are small open economies and it is arguable that domestic price developments are influenced by external sources that were practically ignored in original empirical studies on the NKPC (Batini et al., 2005).

This chapter estimates both forward-looking and hybrid NKPC using different alternative forcing variables both domestic (the output gap, unit labor cost) and external (exchange rate, oil prices, foreign inflation rates). Besides assuming rational expectations in GMM framework (Galí and Gertler, 1999), we use survey data on inflation expectations where available (Zhang et al., 2009). Our main results can be resumed as follows. (1) The inflation rates in the four countries hold significant forward-looking component and therefore the current inflation is (at least partially) determined by its future expected value. However, the size and significance of the backward-looking terms suggest that inflation rates in the NMS are rather persistent. (2) The evidence in favor of the real marginal cost as the main inflation-forcing variable is fragile. The output gap performs better only marginally. On the other hand, we find some evidence that the short-term inflation impulses in the NMS can be external (oil prices, exchange rate, inflation rates abroad). (3) Inflation persistence of the NMS implies that anti-inflationist policy could be accompanied by an output loss. The structural model behind the NKPC suggests that the effect of monetary policy on inflation goes via the marginal cost. However, if the marginal cost does affect inflation, monetary policy can influence inflation only via its credibility and its effect on inflation expectations.

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Chapter 1.

How Does Monetary Policy Change?
Evidence on Inflation Targeting Countries

1.1 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) claim 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

[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).

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.

1.2 Related Literature

1.2.1 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. Taylor and Davradakis (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

[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.

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 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.

1.2.2 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 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, 2009; 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

[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.

[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.

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 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]

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[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).

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 US, 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. 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

[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.
[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.

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).

1.3 Empirical Methodology

1.3.1 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:

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

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 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:

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

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

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[9] The policy neutral rate is typically defined as the sum of the real equilibrium rate and expected inflation that is $\bar{r} = rr + \pi^*$.

[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.

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, defining $\alpha = \overline{r} - \beta\pi^*$, and including additional variables, which results in Eq. (3):

$$r_t = (1-\rho)\left[\alpha + \beta\left(\pi_{t+i}\right) + \gamma y_{t+j} + \delta x_t\right] + \rho r_{t-1} + \varepsilon_t \qquad (3)$$

Following Clarida et al. (1998), the intercept α can be restated in terms of equilibrium real interest rate rr such that $\alpha = rr + (1-\beta)\pi^*$. Hence, the time varying intercept encompass both, changes in equilibrium real interest rate caused by economic fundamentals and inflation target.[11] We set i equal to 2 and j to 0.[12] 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).

[11] Clarida et al. (2000, p. 154) note that in absence of further assumptions this approach does not allow identification of economy's equilibrium real rate rr and inflation target π^*. There are various ways how to overcome this identification problem in the literature. Clarida et al. (1998, 2000) assume that rr can be approximate by sample average of real interest rate, Laubach and Williams (2003) use a set of state space models to recover time variation in equilibrium interest rate. Their approach is used by Leigh (2008) as well. In this paper, we focus on the evolution of policy parameters β, γ and δ and we left the decomposition of time-varying intercept for future research.

[12] 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. For the output gap, we assume a contemporaneous response. As we do not have real time data and rely on statistical construction of the output gap, it is likely that besides the prediction error there is also a construction error that both might be magnified if an unobserved forecast is substituted by the output gap estimate for future periods.

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).[13] Following Kim (2006) we can rewrite Eq. (3) as follows:

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

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

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

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

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

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

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

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

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. Eqs. (10) and (11) track 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). 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 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 equations (10) and (11) and save the standardized residuals φ_t and ν_t. In the second step, we estimate Eq. (12) below

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

along with Eqs. (5)–(9). Note that (12) now includes bias correction terms[14], 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.

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

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

The standard framework for second-step estimation (i.e. Eq. (12)) 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.[15] 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, for different initial values, 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).

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:

$$\sum_{t=1}^{T}\zeta_t^2 + \theta_1\sum_{t=1}^{T}\vartheta^2_{1,\mathrm{t}} + \theta_2\sum_{t=1}^{T}\vartheta^2_{2,\mathrm{t}} + \ldots + \theta_\mathrm{n}\sum_{t=1}^{T}\vartheta^2_{\mathrm{n,t}} \tag{13}$$

[14] 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.1.1 in the Appendix).

[15] 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.

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.[16.] 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.[17]

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 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.[18]

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[16] 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.

[17] 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.

[18] 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.

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, 2009; 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),[19] 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 (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

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

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.

1.3.2 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).[20]

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).[21] 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).

1.4 Results

1.4.1 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

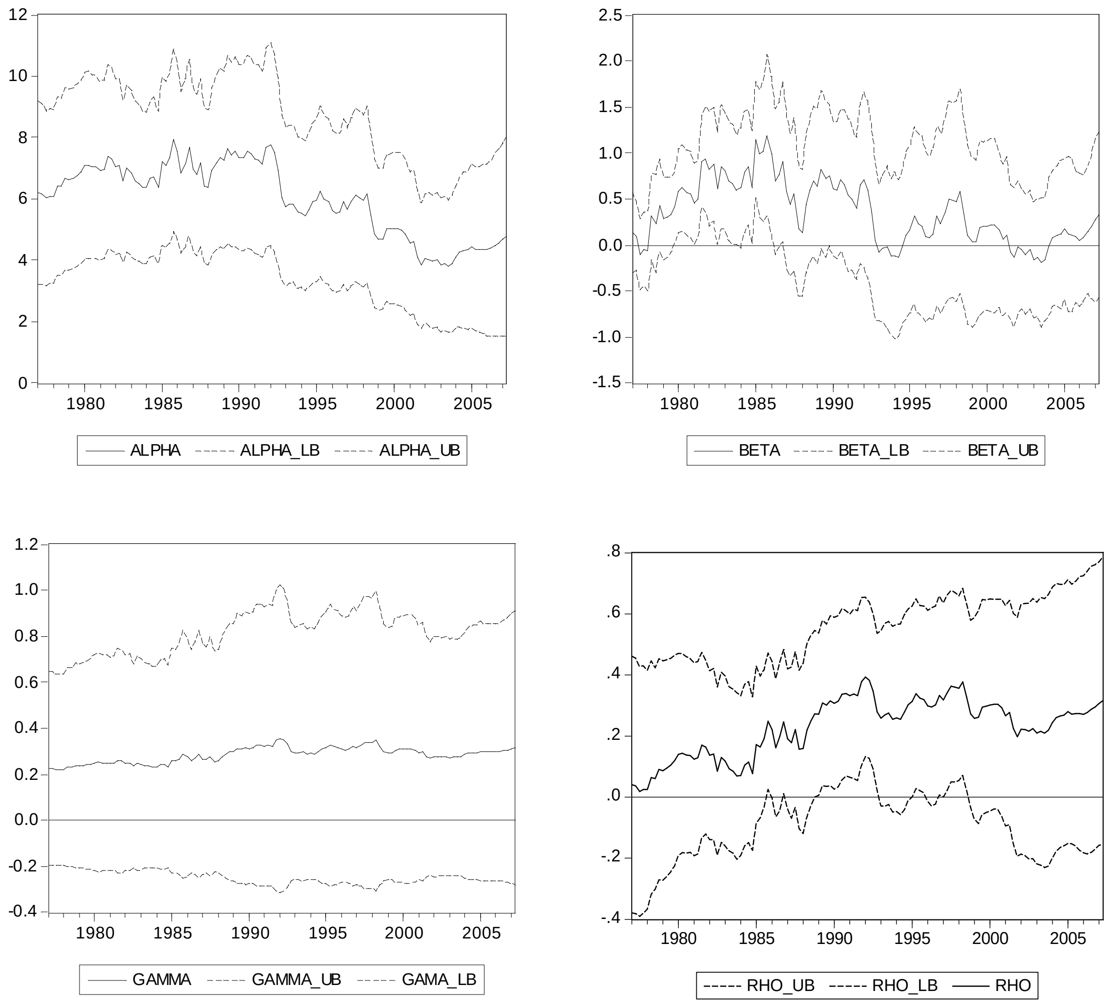
[20] We use year-on-year data, as the inflation target is also defined on a year-on-year basis.

[21] 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.

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 1.1: Time-varying response coefficients in baseline (closed economy) policy rule, the UK



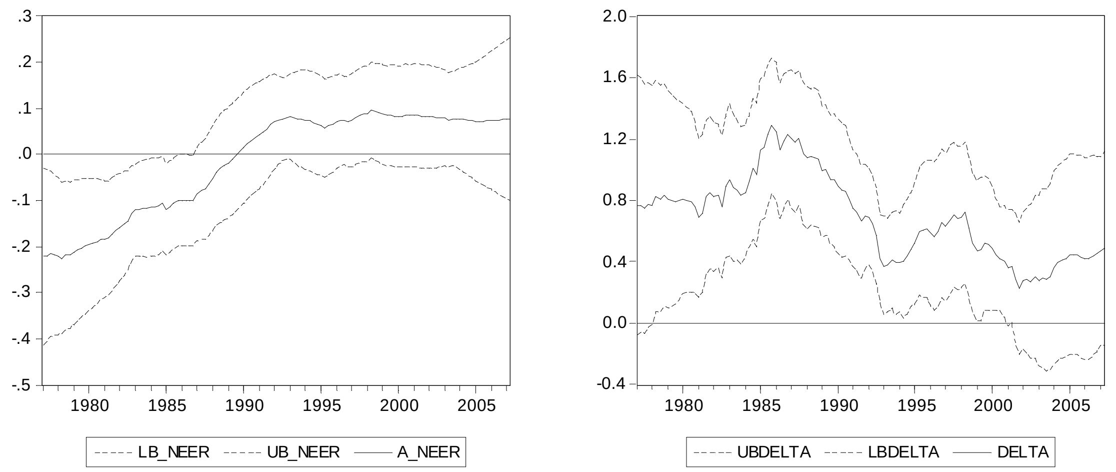
Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The upper-left graph depicts the evolution of the neutral rate. The upper-right graph depicts the evolution of the response of interest rates to inflation. The lower-left graph depicts 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.[22] 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). It was particularly strong during the 1980s and subsequently its importance

[22] 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.

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 1.2: Time-varying response coefficients in augmented (open economy) policy rule, theUK



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.1.1).[23] 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.

1.4.2 New Zealand

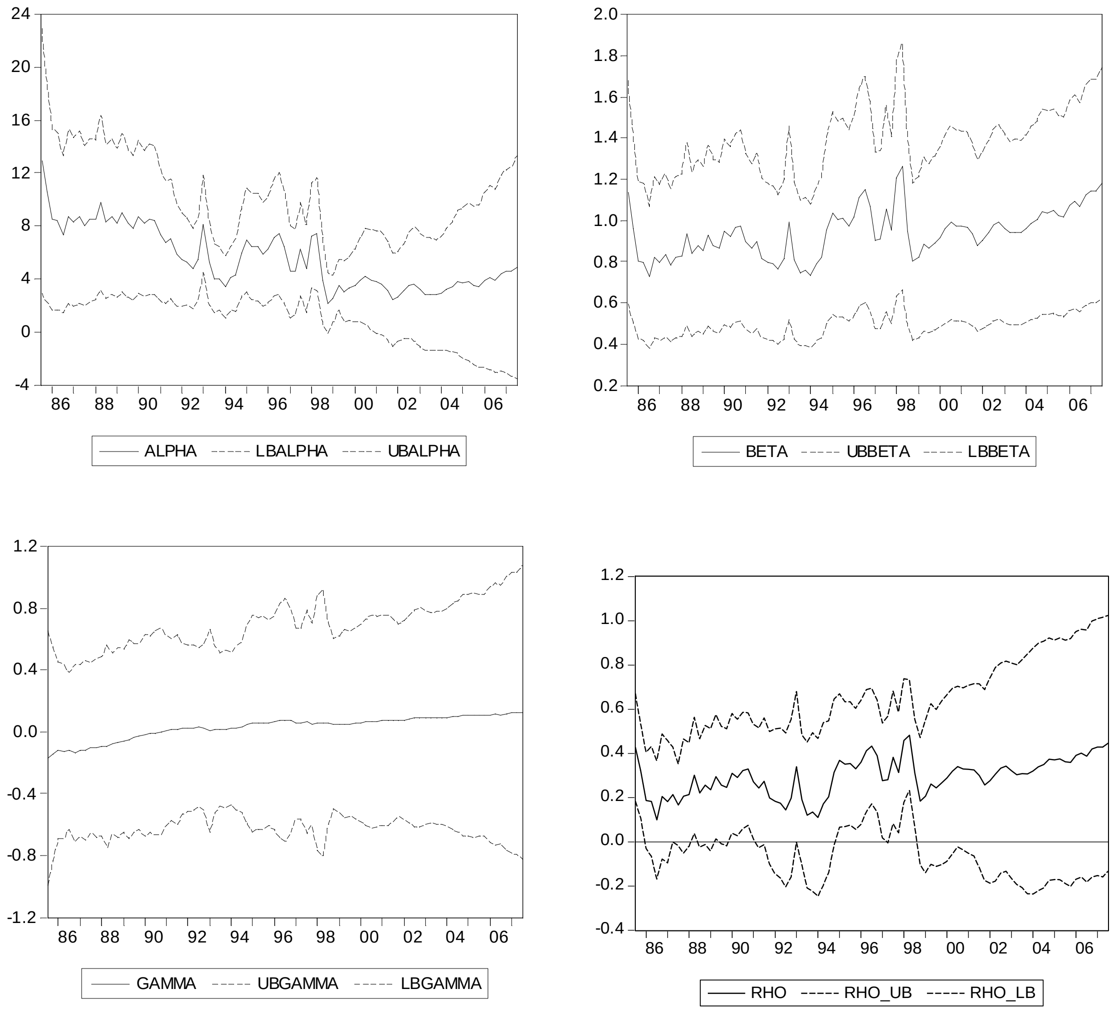
New Zealand was the first country in the world to introduce inflation targeting, doing so by means of the Federal Bank Act signed in March 1990.[24] Our results indicate that the response of the RBNZ to expected inflation was very close to unity throughout the sample period (1985–2007). In fact, it is clearly visible that the interest rate and inflation series move together very closely. However, in Figure 1.3 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.[25] 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.

[23] 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.

[24] 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.

[25] 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.

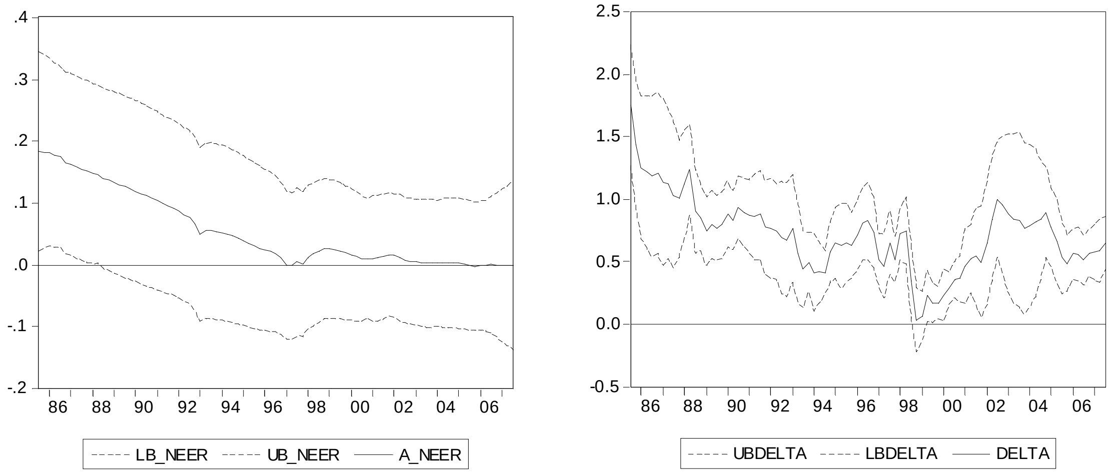
Figure 1.3: Time-varying response coefficients in baseline (closed economy) policy rule, New Zealand



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 1.4: Time-varying response coefficients in augmented (open economy) policy rule, New Zealand



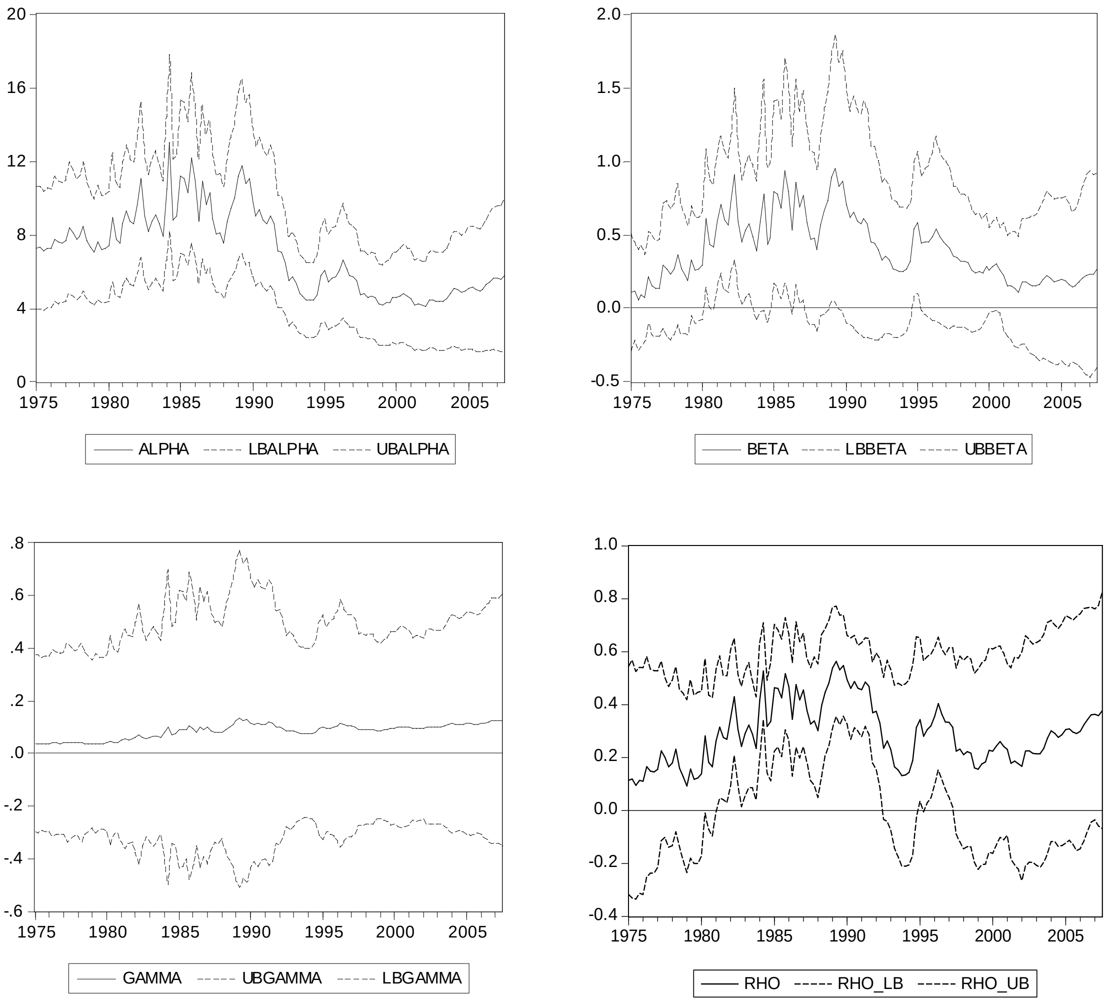
Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The left-hand graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (the deviation from the HP trend). The right-hand graph depicts the evolution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

1.4.3 Australia

Our results for Australia are available in Figures 1.5 and 1.6. 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.

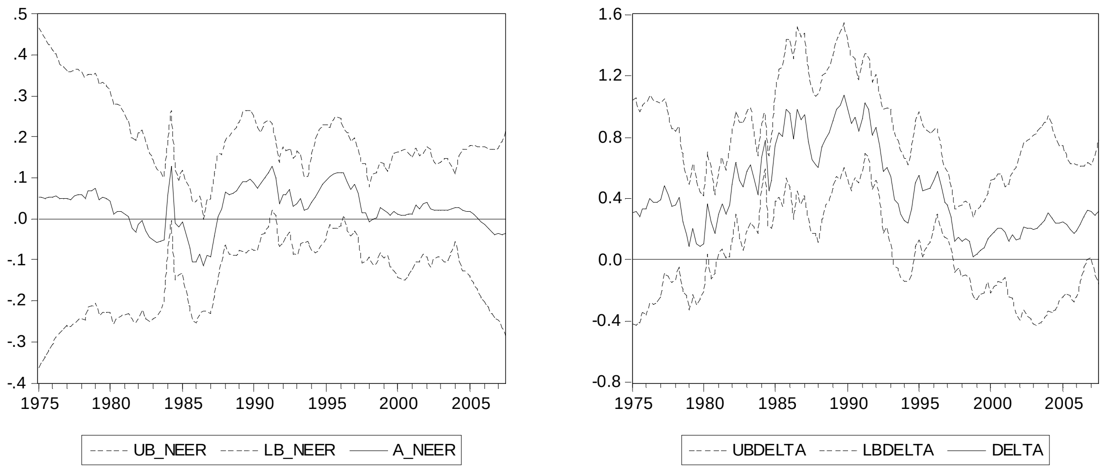
Figure 1.5: Time-varying response coefficients in baseline (closed economy) policy rule, Australia



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 1.6: 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.

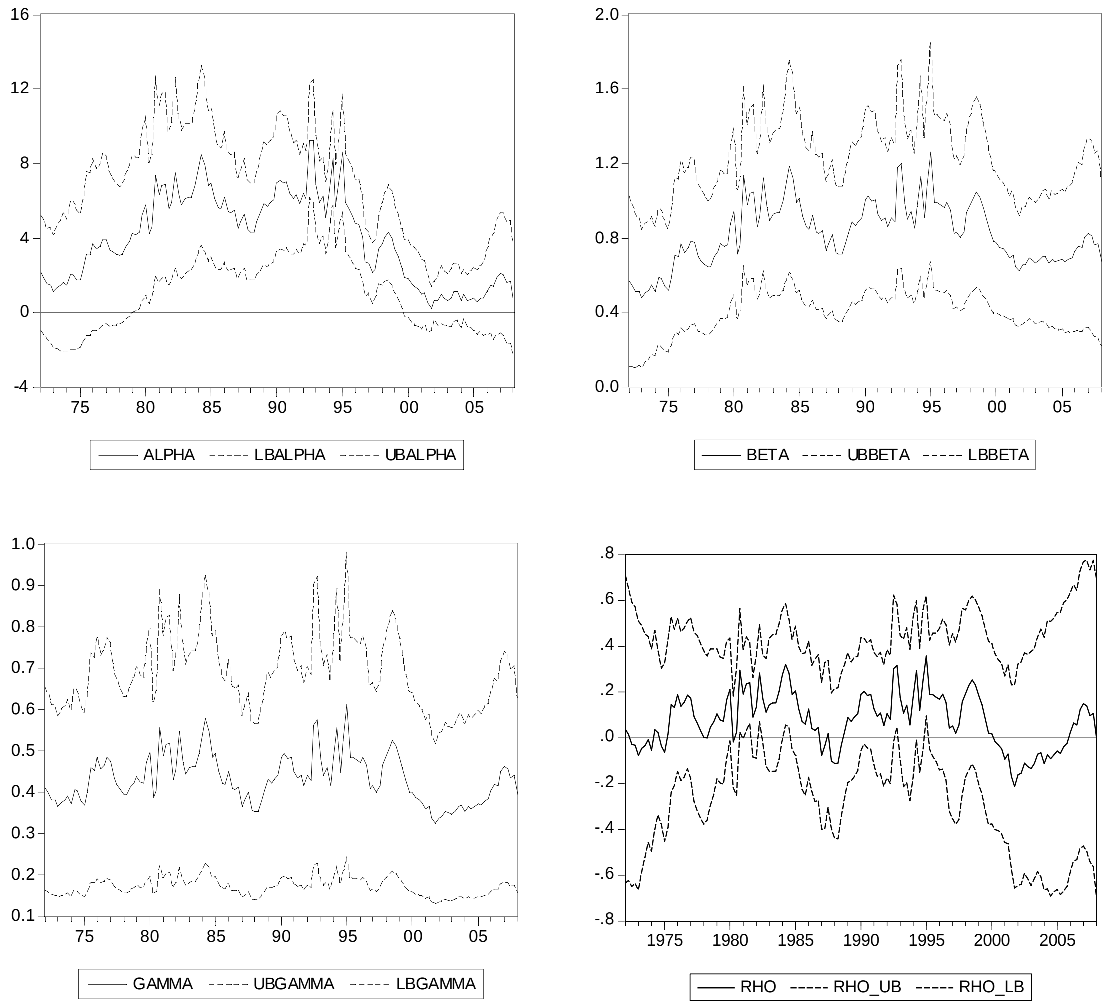
1.4.4 Canada

The monetary policy rule estimates for Canada are presented in Figures 1.7 and 1.8. 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 1.7). 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).[26]

[26] 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 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 1.7: 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

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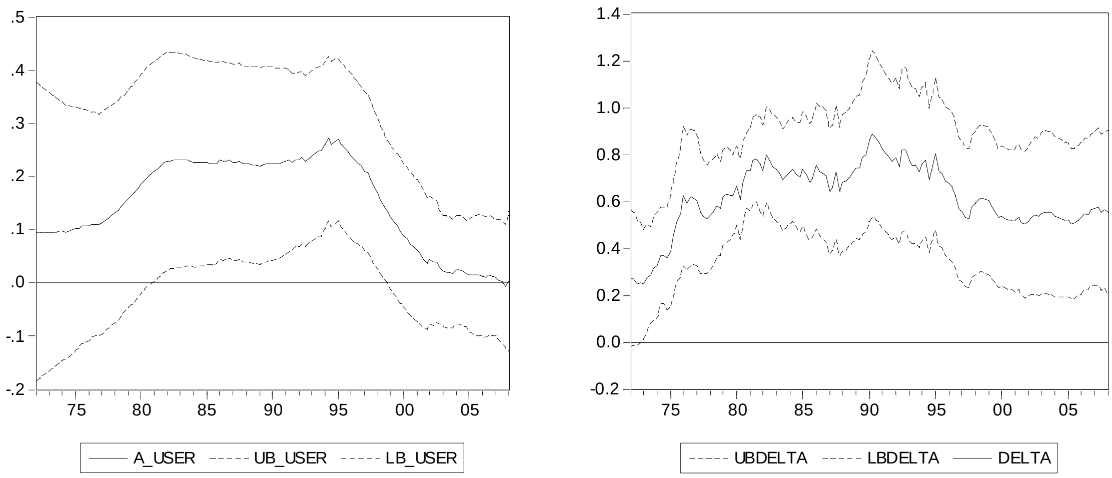
the inflation rate adjusted for the transitory effect of the exchange rate on import prices. However, there is no indication that the BoC actually ever used the MCI for practical policy making, and it ceased to publish it in 2006.

rate. The 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 1.8: Time-varying response coefficients in augmented (open economy) policy rule, Canada



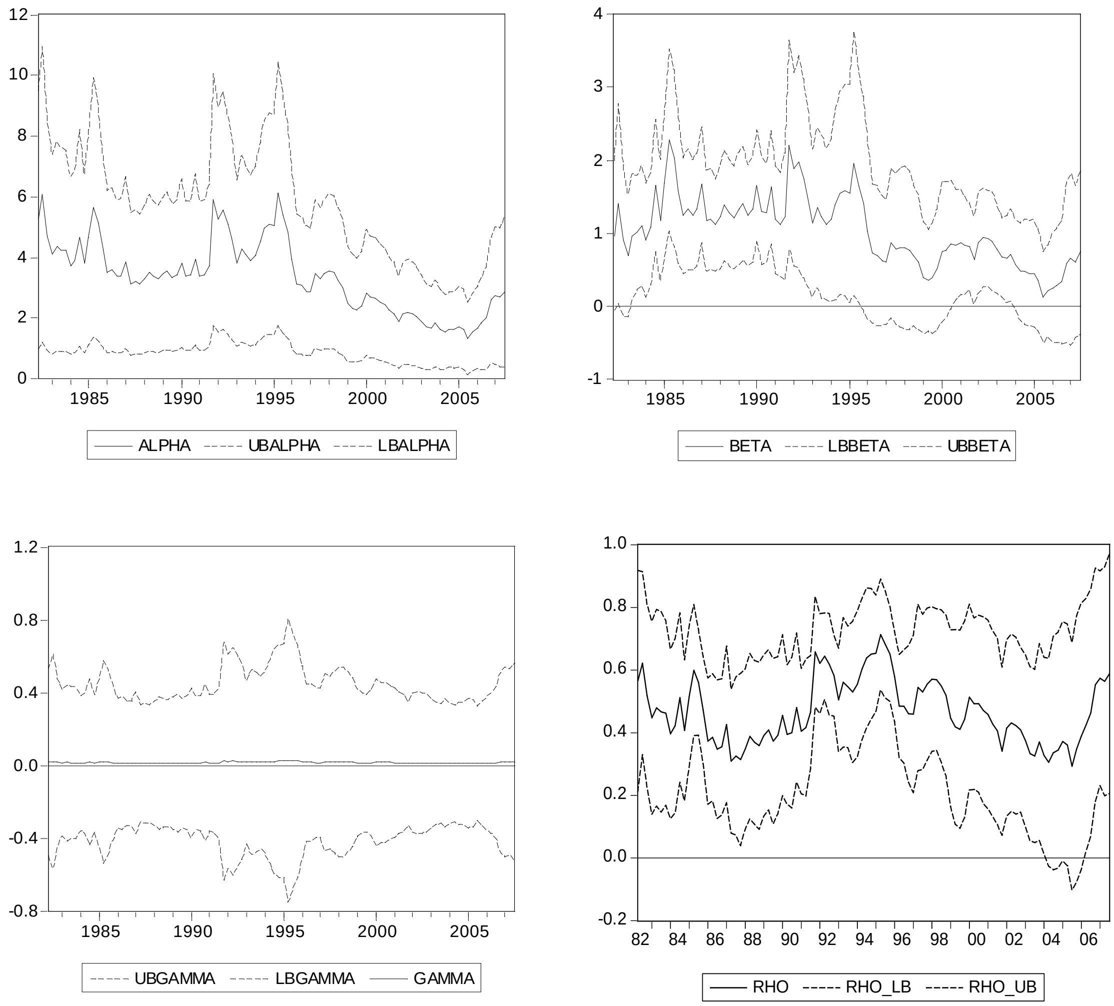
Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The left-hand graph depicts the evolution of the response of interest rate to the nominal effective exchange rate (the deviation from the HP trend). The right-hand graph depicts the evolution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

1.4.5 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.[27]

[27] 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 1.9: Time-varying response coefficients in baseline (closed economy) policy rule, Sweden

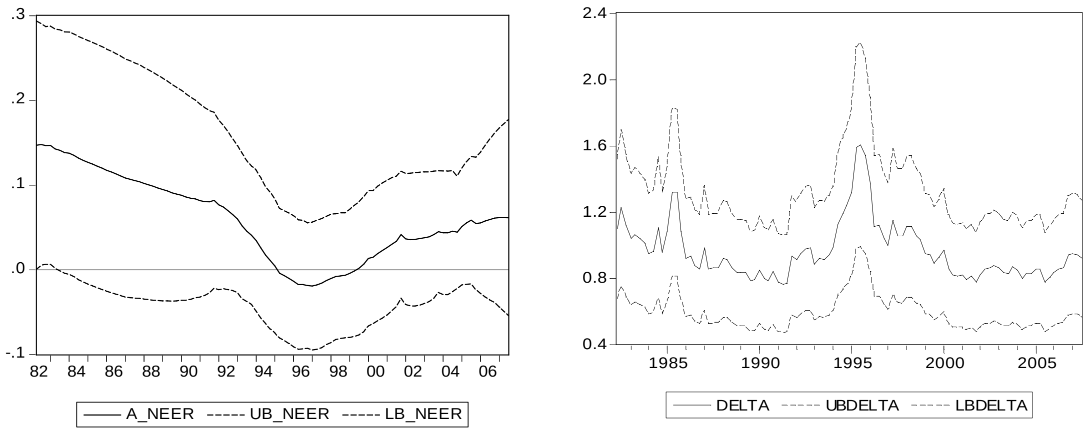


Note: 95% confidence bands; model with bias correction terms, i.e. dealing with endogeneity in monetary policy rules. The upper-left graph depicts the evolution of the neutral rate. The upper-right graph depicts the evolution of the response of interest rates to inflation. The lower-left graph depicts 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 1.10: 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.

1.4.6 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 the 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.

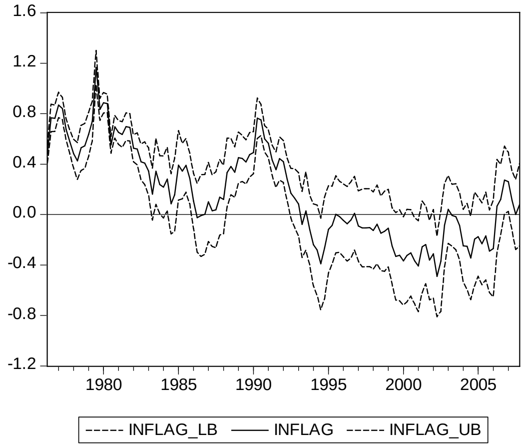
To shed some light on this issue, we used our estimation framework and fitted the AR(1) model with drift to inflation series, allowing the coefficient on lagged inflation as well as the constant to be time-varying. Our results (as reported in Figure 1.11) indicate that inflation persistence decreased 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.[28] In addition, the results reported in Figure 1.11 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 (for the US see e.g. Benati and Surico, 2009; Canova and Gambetti, 2008; Gambetti et al., 2008; Sims and Zha, 2006; and for the UK and the Euro area Canova et al., 2007). 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.

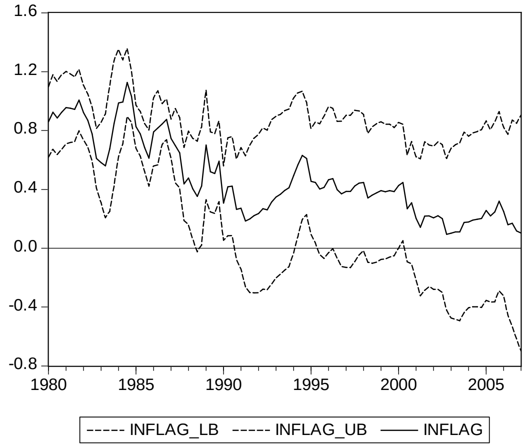
[28] See Hondroyiannis et al. (2009) on the estimation of the Phillips curve within a somewhat different time-varying framework.

Figure 1.11: Time-varying response coefficients in AR(1) model for inflation

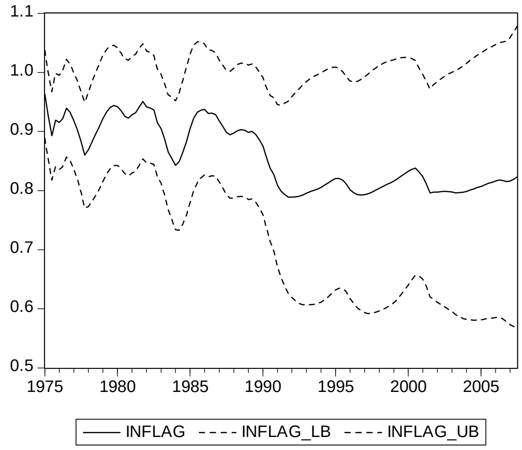
UK



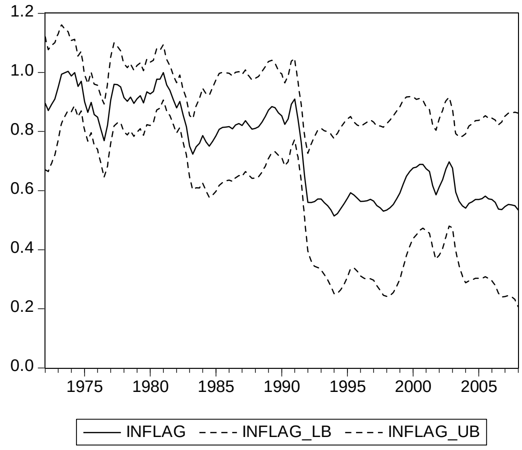
New Zealand



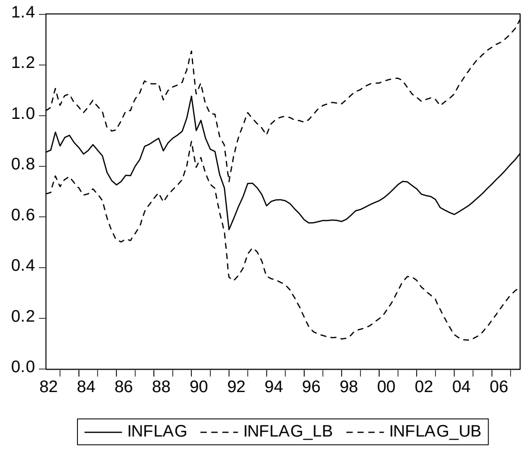
Australia



Canada



Sweden



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.

1.5 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

A.1.1 The VC Method

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

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

Define the following matrices:

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

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.

A.1.1.1 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.[29]

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

[29] 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).

A.1.1.2 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.[30]

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

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

[30] 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.

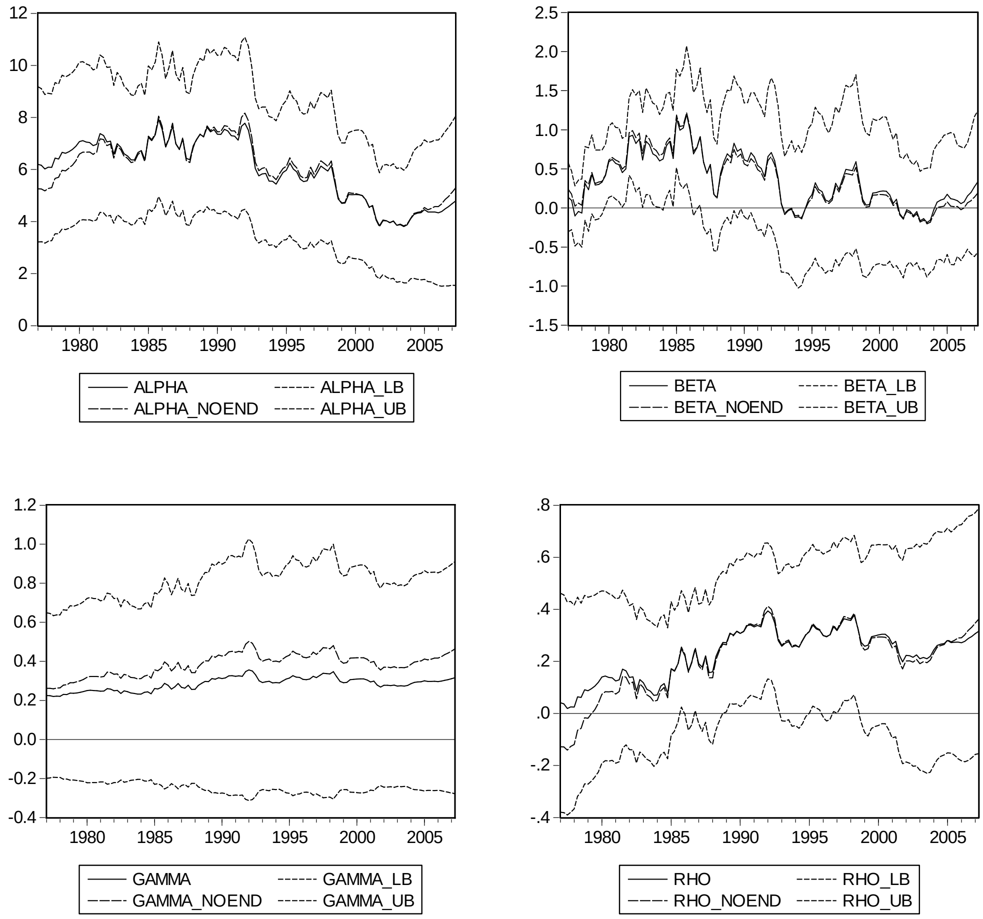
which is used for the estimation of coefficients $\hat{a}_t$.

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).

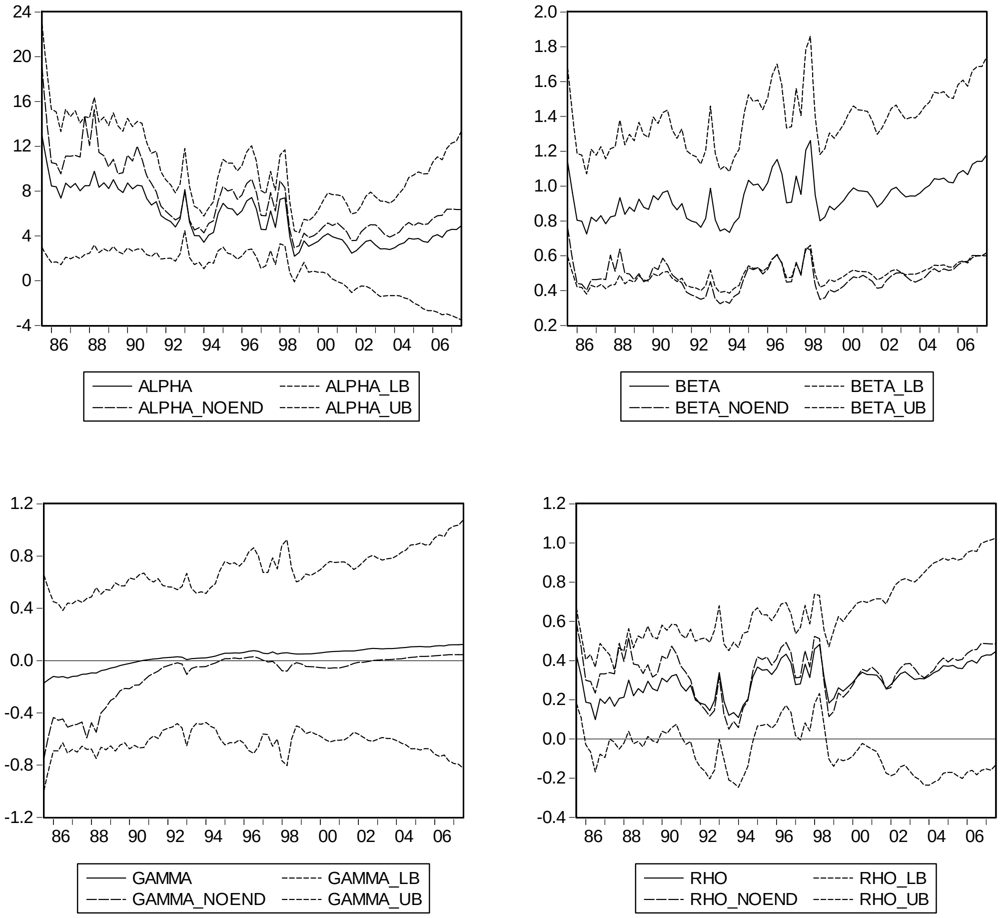
A.1.2 The effect of endogeneity in the policy rule

Figure A.1.1: Time-varying response coefficients in baseline (closed economy) policy rule, UK



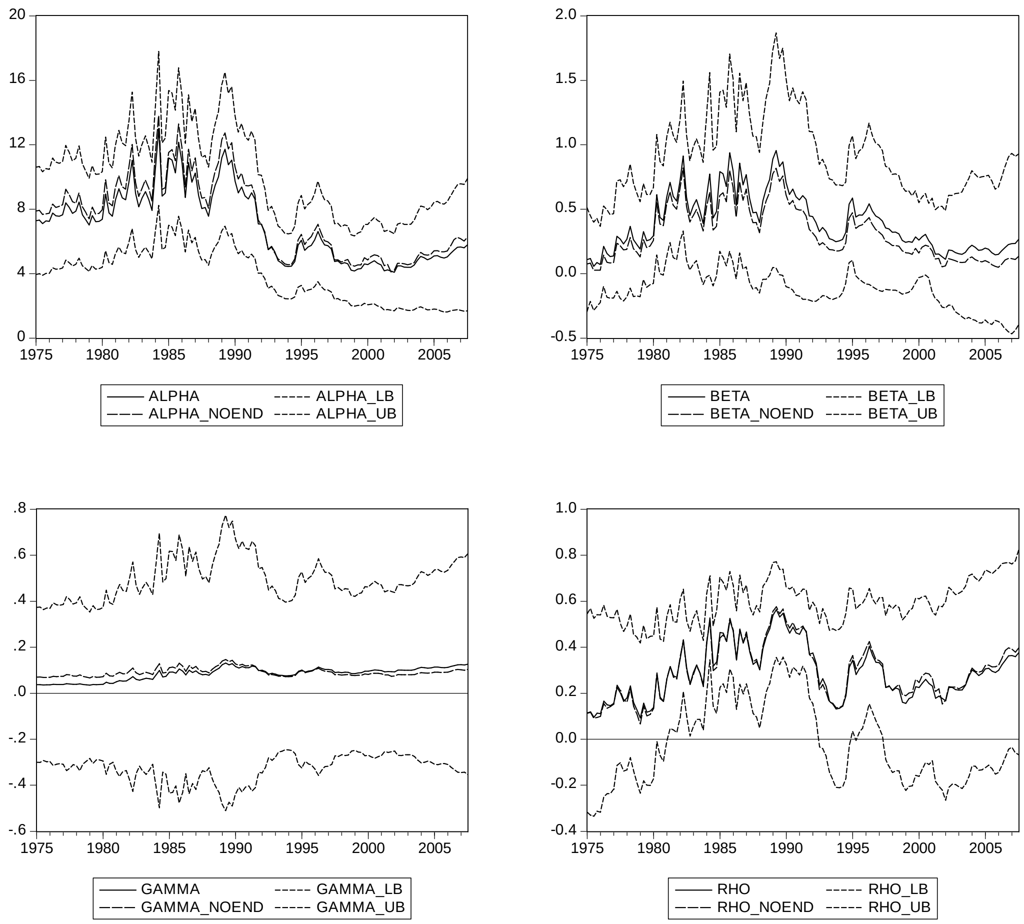
Note: Model with bias correction terms (solid line); model not dealing with endogeneity in monetary policy rules (dashed line). The 95% confidence bands correspond to the model with bias correction terms.

Figure A.1.2: Time-varying response coefficients in baseline (closed economy) policy rule, New Zealand



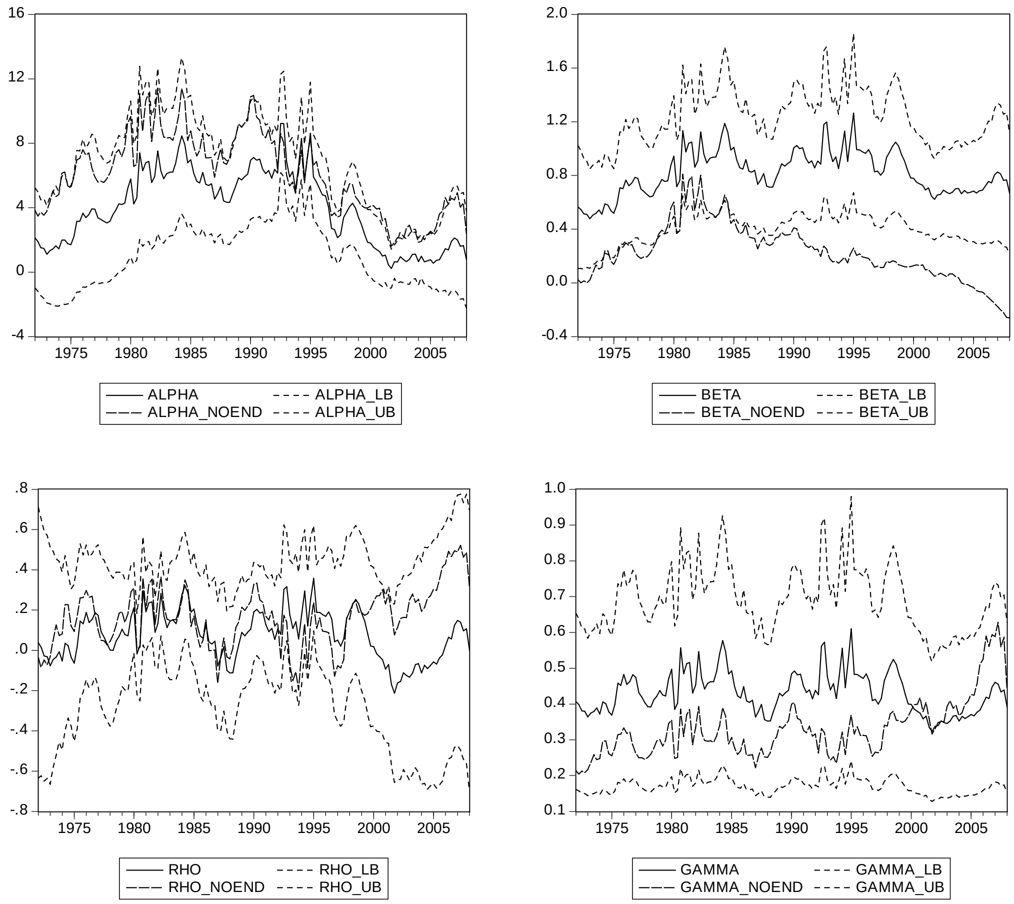
Note: Model with bias correction terms (solid line); model not dealing with endogeneity in monetary policy rules (dashed line). The 95% confidence bands correspond to the model with bias correction terms.

Figure A.1.3: Time-varying response coefficients in baseline (closed economy) policy rule, Australia



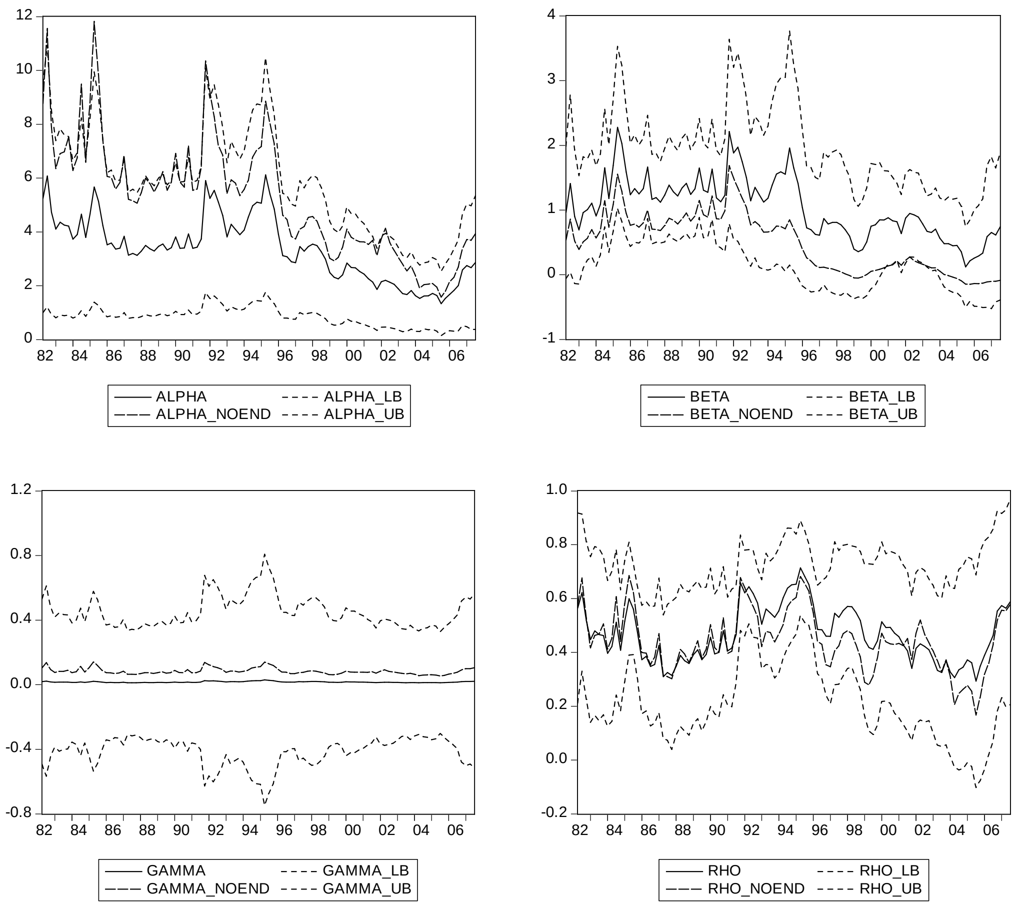
Note: Model with bias correction terms (solid line); model not dealing with endogeneity in monetary policy rules (dashed line). The 95% confidence bands correspond to the model with bias correction terms.

Figure A.1.4: 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.1.5: 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.1.1: 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%.

Chapter 2.

Monetary Policy Rules and Financial Stress: Does Financial Instability Matter for Monetary Policy?

2.1 Introduction

The recent financial crisis has intensified the interest in exploring the interactions between monetary policy and financial stability. Official interest rates were driven sharply to historic lows and many unconventional measures were used to pump liquidity into the international financial system. Central banks pursued monetary policy under high economic uncertainty coupled with large financial shocks in many countries. The financial crisis also raised new challenges for central bank policies, in particular how to operationalize the issues related to financial stability for monetary policy decision-making (Goodhart, 2006; Borio and Drehmann, 2009).

This paper seeks to analyze whether and how central banks reacted to the periods of financial instability, and in particular whether and how the interest-setting process evolved in response to financial instability over the last three decades. The monetary policy of central banks is likely to react to financial instability in a non-linear way (Goodhart et al., 2009). When a financial system is stable, the interest rate setting process largely reflects macroeconomic conditions and financial stability considerations enter the monetary policy discussions only to a limited degree. On the other hand, central banks may alter its monetary policy to reduce financial imbalances if these become severe. In this respect, Mishkin (2009) questions the traditional linear-quadratic framework[1] when financial markets are disrupted and puts forward the arguments for replacing it by nonlinear dynamics, describing the economy and the non-quadratic objective function resulting in the non-linear optimal policy.

To deal with the complexity of monetary policy and financial stability nexus as well as to evaluate monetary policy in a systematic manner, this paper employs the recently developed time-varying parameter estimation of monetary policy rules appropriately accounting for endogeneity in policy rules. This flexible framework, together with a new comprehensive financial stress dataset developed by the International Monetary Fund, will allow not only

[1] Linear behavior of the economy and the quadratic objective function of monetary authority.

testing whether the central banks responded to financial stress but also the quantification of the magnitude of this response and the detection of the periods and types of stress that were the most worrying for monetary authorities.

Although theoretical studies disagree about the role of financial instability for central bank interest rate setting policy, our empirical estimates of time-varying monetary policy rules of the US Fed, the Bank of England (BoE), Reserve Bank of Australia (RBA), Bank of Canada (BoC) and Sveriges Riksbank (SR) shows that central banks often alter the course of its monetary policy in the face of high financial stress, mainly by decreasing their policy rates. However, the size of this response varies substantially over time as well as across countries. There is a certain across country and time heterogeneity as well when we look at central banks' consideration of specific types of financial stress: Most of them seemed to respond to stock market stress and bank stress, and exchange rate stress drives central bank reactions only in more open economies

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

2.2 Related Literature

First, this section gives a brief overview of the theory as well as empirical evidence on the relationship between monetary policy (rules) and financial instability. Second, it provides a short summary of various measures of financial stress.

2.2.1 Monetary policy (rules) and financial instability – some theories

Financial friction, such as an unequal access to credits or debt collateralization, were recognized to have important consequences for monetary policy transmission and Fisher (1933) has already presented the idea that adverse credit-market conditions can cause significant macroeconomic disequilibria.

During the last two decades, the effects of monetary policy have been studied mainly within New Keynesian (NK) dynamic stochastic general equilibrium (DSGE) models, which assume the existence of nominal rigidities. The common approach to incorporate financial market friction within the DSGE framework is to introduce the financial accelerator mechanism (Bernanke et al., 1996, 1999), implying that endogenous developments in credit markets work to amplify and propagate shocks to the macroeconomy. Tovar (2009) emphasizes that the major weakness of the financial accelerator mechanism is that it only deals with one of many possible financial frictions. Goodhart et al. (2009) notes that many NK DSGE models lack the financial sector completely or modeled it in a rather embryonic way. Consequently, more recent contributions within this stream of literature examined other aspects of financial frictions such as the balance sheets in the banking sector (Choi and Cook, 2004), portfolio choice issue with complete (Engel or Matsumoto, 2009) or incomplete markets (Devereux and Sutherland, 2007) or collateral constraints (Iacovello and Neri, 2010).[2]

A few studies focus more specifically on the relationship between the monetary policy stance (or the monetary policy rule) and financial stability. However, they do not arrive at a unanimous view on whether a monetary policy rule should include some measure of financial stability. Brousseau and Detken (2001) present an NK model where a conflict arises between short-term price stability and financial stability due to a self-fulfilling belief linking the stability of inflation to the smoothness of the interest rate path and suggests that monetary policy should react to financial instability. Akram et al. (2007) investigate the macroeconomic implications of pursuing financial stability within a flexible inflation-targeting framework. Their model using policy rule, augmented with financial stability indicators, shows that the gains of such an augmented rule vis-à-vis the rule without financial stability indicators highly depends on the nature of the shocks. Akram and Eitrheim (2009) build on the previous framework, finding some evidence that the policy response to housing prices, equity prices or credit growth can cause high interest rate volatility and actually lower financial stability in terms of indicators that are sensitive to interest rates. Ceccheti and Li (2008) show in a static and dynamic setting that a potential conflict between monetary policy and financial supervision can be avoided if the interest rate rule takes (procyclical) capital adequacy requirements into account, in particular that the policy interest rates are lowered when financial stress is high. Bauducco et al. (2008) extends the current benchmark NK model to include financial systems and firms that require external financing. Their simulations show that if a central bank responds to financial instability by policy easing it achieves better inflation and output stabilization in the short term at the cost of greater inflation and output

[2] The survey of this literature is provided by Tovar (2009).

volatility in the long term and vice versa. For the US Fed Taylor (2008) proposes a modification of a standard Taylor rule to incorporate adjustments to credit spreads. Teranishi (2009) derives a Taylor rule augmented by the response to credit spreads as an optimal policy under heterogeneous loan interest rate contracts. He finds that the policy response to a credit spread can be both positive and negative depending on the financial structure. However, he also puts forward that when nominal policy rates are close to zero a commitment rather than discretional policy response is the key for reducing the credit spreads. Christiano et al. (2008) suggest augmenting the Taylor rule with aggregate private credit and find that such a policy would raise welfare by reducing the magnitude of the output fluctuations. Cúrdia and Woodford (2010) develop an NK DSGE model with credit frictions to evaluate the performance of alternative policy rules that are augmented by a response to credit spreads and to aggregate the volume of private credit in the face of different shocks. They argue that the response to credit spreads can be welfare improving, but the optimal size of such a response is likely rather small. Like Teranishi (2009), they find little support for augmenting a Taylor rule by the credit volume given — that the size and even the sign of the desired response is sensitive to the sources of shock and their persistence — which is information that is not always available during operational policy making.

The related stream of literature focuses on a somewhat narrower issue of whether or not monetary policy should respond to asset prices. Bernanke and Gertler (1999, 2001) argue that the stabilization of inflation and output provides a substantial contribution to financial stability and there are little if any gains to responding to asset prices. Faia and Monacelli (2007) extend the model developed by Bernanke and Gertler (2001) by a robust welfare metric confirming that strict inflation stabilization offers the best solution. Cecchetti et al. (2000) takes the opposite stand arguing that developments in asset markets can have a significant impact on both inflation and real economic activity, and central banks might achieve better outcomes considering the asset prices provided they are able to detect their misalignments. Borio and Lowe (2002) support this view claiming that financial imbalances can build up even in a low inflation environment, which is normally favorable to financial stability. The side effect of low inflation is that excess demand pressures may first appear in credit aggregates and asset prices rather than consumer prices, which are normally considered by the policy makers. Gruen et al. (2005) argues that responding to an asset bubble is feasible only when the monetary authority is able to make a correct judgment about the process driving the bubble. Roubini (2006) and Posen (2006) provide the summary of this debate from a policy perspective.

2.2.2 Monetary policy (rules) and financial instability – empirical evidence

The empirical evidence on central banks' reaction to financial instability is rather scant. Following the ongoing debate about whether central banks should respond to asset price volatility (e.g. Bernanke and Gertler, 1999, 2001; Cecchetti et al., 2000; Bordo and Jeanne, 2002), some studies tested the response of monetary policy to different asset prices, most commonly to stock prices (Rigobon and Sack, 2003; Siklos and Bohl, 2008; Fuhrer and Tootel, 2008). They find some evidence that asset prices either entered the policy information set (because they contain information about future inflation) or that some central banks were directly trying to offset its disequilibria.[3] All these papers estimate time-invariant policy rules, which means that they test a permanent response to these variables. However, it seems more plausible that if central banks respond to asset prices, they do it only when their misalignments are substantial, in other words their response is asymmetric. There are two additional controversies related to the effects of asset prices on monetary policy decisions: (i) The first concerns the measure, in particular whether the stock market index that is typically employed is sufficiently representative or whether some other assets, in particular the housing prices, should be considered as well, and (ii) the second issue is related to the (even ex-post) identification of the asset price misalignment. Finally, it is likely that the perception of misalignments is influenced by general economic conditions and that a possible response could evolve over time. Detken and Smets (2004) summarize some stylized facts on macroeconomic and monetary policy developments during asset price booms. Overall they find that monetary policy was significantly looser during the high-cost booms that were marked by the investment and real estate prices crash in the post-boom periods.

A few empirical studies measure monetary policy response using broader measures of financial imbalances. Borio and Lowe (2004) estimate the response of four central banks (Reserve Bank of Australia, Bundesbank, Bank of Japan and the US Fed) to imbalances proxied by the ratio of private sector credit to GDP, inflation-adjusted equity prices and their composite. They find either negative or ambiguous evidence for all countries except for the USA confirming that the Fed responded to financial imbalances in an asymmetric and reactive

[3] A similar but somewhat less polemic debate applies to the role of the exchange rate, especially for small open economies (Taylor, 2001).

way, i.e. that the federal fund rate was disproportionally lowered in the face of imbalance unwinding but it was not tightened beyond normal as imbalances built up. Ceccheti and Li (2008) estimate a Taylor rule augmented by a measure of banking stress, in particular a deviation of leverage ratios (total loans to the sum of equity and subordinated debt; total assets to the sum of bank capital and reserves) from its Hodrick-Prescott trend. They find some evidence that the Fed adjusted the interest rate in order to counteract the procyclical impact of a bank's capital requirements, while the Bundesbank and the Bank of Japan did not. Bulíř and Čihák (2008) estimate the monetary policy response to seven alternative measures of financial sector vulnerability (crisis probability, time to crisis, distance to default or credit default swap spreads) in a panel of 28 countries. Their empirical framework is different in the sense that the monetary policy stance is proxied along the short-term interest rate by measures of domestic liquidity and external shocks are controlled for. In the panel setting, they find a statistically significant negative response to many variables representing vulnerability (policy easing) but surprisingly not in country-level regressions. Belke and Klose (2010) investigate the factors behind the interest rate decision of the ECB and the Fed during the current crisis. They conclude that the estimated policy rule was significantly altered only for the Fed and they put forward that the ECB gave greater weight on inflation stabilization at the cost of some output loss.

2.2.3 Measures of financial stress

The incidence and determinants of different types of crises have been typically traced in the literature by a means of narrative evidence (expert judgment). This was sometimes complemented with selected indicators (the exchange rate devaluation, the state of foreign reserves) that point to historical regularities (e.g. Eichengreen and Bordo, 2002; Kaminsky and Reinhart, 1999; Reinhart and Rogoff, 2009; Laeven and Valencia, 2008). The empirical studies (e.g. Goldstein et al., 2000) used binary variables that were constructed based on these narratives.

Consequently, some contributions strived to provide more data-driven measures of financial stress. Most of the existing stress indices are based on high-frequency data but they differ in the selected variables (bank capitalization, credit ratings, credit growth, interest rate spreads or volatility of different asset classes), country coverage and the aggregation method. An important advantage of continuous stress indicators is that it may reveal periods of small-scale stress that did not result in full-blown crisis and were neglected in studies based on binary crisis variables.

The Bank Credit Analyst (BCA) reports a monthly financial stress index (FSI) for the USA that is based on the performance of banking shares as compared to whole stock market, credit spreads and the slope of the yield curve, and the new issues of stocks, bonds and consumer confidence. JP Morgan calculates a Liquidity, Credit and Volatility Index (LCVI) based on seven variables: the US Treasury curve error (standard deviation of the spread between on-the-run and off-the-run US Treasury bills and bonds along the entire maturity curve), the 10-year US swap spread, US high-yield spreads, JP Morgan's Emerging Markets Bond Index, foreign exchange volatility (weighted average of 12-month implied volatilities of several currencies), the Chicago Board of Exchange equity volatility index VIX, and the JP Morgan Global Risk Appetite Index.

Illing and Liu (2006) develop a comprehensive FSI for Canada. Their underlying data covers equity, bond and foreign exchange markets as well as the banking sector. They use a standard measure and refined measure of each stress component, where the former refers to the variables and their transformations that are commonly found in the literature, while the latter incorporates the adjustments that allow for better extraction of the information about stressful periods. They explore different weighting schemes to aggregate the individual series (factor analysis, size of a corresponding market on total credit in economy, variance-equal weighting). At last, they perform an expert survey to identify periods that were perceived as especially stressful, confirming that the FSI matches these episodes very well.

Carlson et al. (2008) propose for the Fed Board of Governors a framework similar to the option pricing model (Merton, 1974) that aims to provide a distance-to-default of the financial system, the so-called Index of Financial Health. The method uses the difference between the market value of a firm's assets and liabilities and the volatility of the asset's value in order to measure the proximity of a firm's assets being exceeded by its liabilities. They apply this measure to 25 of the largest US financial institutions, confirming its impact on capital investments in the US economy. The FED of Kansas City developed the Kansas City Financial Stress Index (Hakkio and Keeton, 2009) that is published monthly and is based on eleven variables (seven spreads between different bond classes by issuers, risk profiles and maturities, correlations between returns on stocks and Treasury bonds, expected volatility of overall stock prices, volatility of bank stock prices and a cross-section dispersion of bank stock returns) that are aggregated by principal component analysis.

Finally, the International Monetary Fund (IMF) recently published financial stress indices for various countries. Cardarelli et al. (2009) propose a comprehensive index based on high-frequency data where the price changes are measured with respect to its previous levels or trend value. The underlying variables are standardized and aggregated into a single index (FSI) using variance-equal weighting for each country and period. The FSI has three subcomponents: the banking sector (the slope of the yield curve, TED spread, the beta of banking sector stock), securities markets (corporate bonds spread, stock market returns and time-varying volatility of stock returns) and exchange rate (time-varying volatility of NEER changes). Balaskrishnan et al. (2009) modify the previous index to account for specific conditions of emerging economies; on one hand including the measure of exchange rate pressures (currency depreciation and decline in foreign reserves) and sovereign debt spread, on the other hand downplaying the banking sector measures (slope of the yield curve and TED spread).[4] We will use the former index given its comprehensiveness as well as its availability for different countries (see more details below).

2.3 Data and Empirical Methodology

2.3.1 The dataset

We use monthly data (quarterly for Sweden and Canada, more details below). The sample period slightly varies due to data availability (the US 1981:1 M – 2009:6M; the UK 1981:1M -2009:3M; Australia 1983:3M – 2009:5M; Canada 1981:1Q – 2008:4Q; Sweden 1984:2Q – 2009:1Q).

The dependent variable is typically the interest rate closely related to the official (censored) policy rate, in particular the federal fund rate (3M) for the USA, the discount rate (3-month treasury bills) for the UK, Canada and Sweden, and the 3-month RBA accepted bills rate for Australia. It is evident that policy rate is not necessarily the only instrument that the central bank uses, especially during the 2008-2009 global financial crisis when many unconventional measures were implemented (see Borio and Disyatat, 2009; Reis, 2010). So as to address this

[4] The IMF Financial Stress Index has been recently applied by Melvin and Taylor (2009) to analyze the exchange rate crises.

issue in terms of estimated policy rules, for robustness check we use the interbank interest rate (at a maturity of 3 months). While both rates are used in empirical papers on the monetary policy rules estimation without great controversy, the selection of the interest rate turns more delicate during financial stress periods (Taylor, 2008). While the former is more directly affected by genuine monetary policy decisions (carried by open market operations), the latter additionally includes liquidity conditions on the interbank markets and as such can be affected by some unconventional policies, though these are usually insulated (often intentionally) from policy interest rates.[5] This represents a drawback but also a potential advantage of this alternative dependent variable. On the one hand, the changes in official policy rate may not fully pass through into interbank interest rates, in particular when the perceived counterparty risk is too high and the credit spreads widen (see Taylor and Williams, 2009). On the other hand the interbank rate may also incorporate the impact of policy actions such as quantitative easing aimed to supply additional liquidity into the system.[6]

The inflation is measured as a year-on-year change of CPI, apart from the United States where we use the personal consumption expenditures price index (PCE) and Sweden where the underlying inflation CPIX (excludes households' mortgage interest expenditure and the direct effects of changes in indirect taxes and subsidies from the CPI) is used.[7] The output gap is proxied by the gap of the seasonally adjusted industrial production index derived by the Hodrick-Prescott filter with smoothing parameter set to 14400. For Sweden and Canada, where we use quarterly data, the output gap was taken as reported in the OECD Economic Outlook (production function method based on NAWRU — non-accelerating wages rate of unemployment).

[5] Borio and Disyatat (2009) characterize the unconventional policies as policies that affect the central bank's balance sheet size and composition and that can be insulated from the interest rate policy (so-called "decoupling principle"). One common example of such a policy (not necessarily used during the time of crisis) is a sterilized exchange rate intervention. Given that we aim not for a single episode of stress but rather want to identify whether monetary authorities deviated from its systematic pattern (the policy rule) during these periods (by responding to indicators of financial stress), we need to use a consistent measure of policy actions that are adjusted during the periods of financial stress, though other measures can be in place as well. Therefore, we assume that the monetary policy stance is fully reflected in the interest rate, and we are aware that it might be subject to a downward bias on the financial stress coefficient. The reader may want to interpret our results on the importance of financial stress on interest rate setting as a conservative estimate.

[6] There are other policy measures that can be used as a reactive or pre-emptive response to financial stress such as regulatory or administrative measures, although their effects are likely to appear only in the longer term and cannot be reasonably included in our empirical analysis.

[7] For Australia, the monthly CPI is not available because both the Reserve Bank of Australia and the Australian Bureau of Statistics only publish quarterly data. The monthly series was obtained using a linear interpolation of the CPI index.

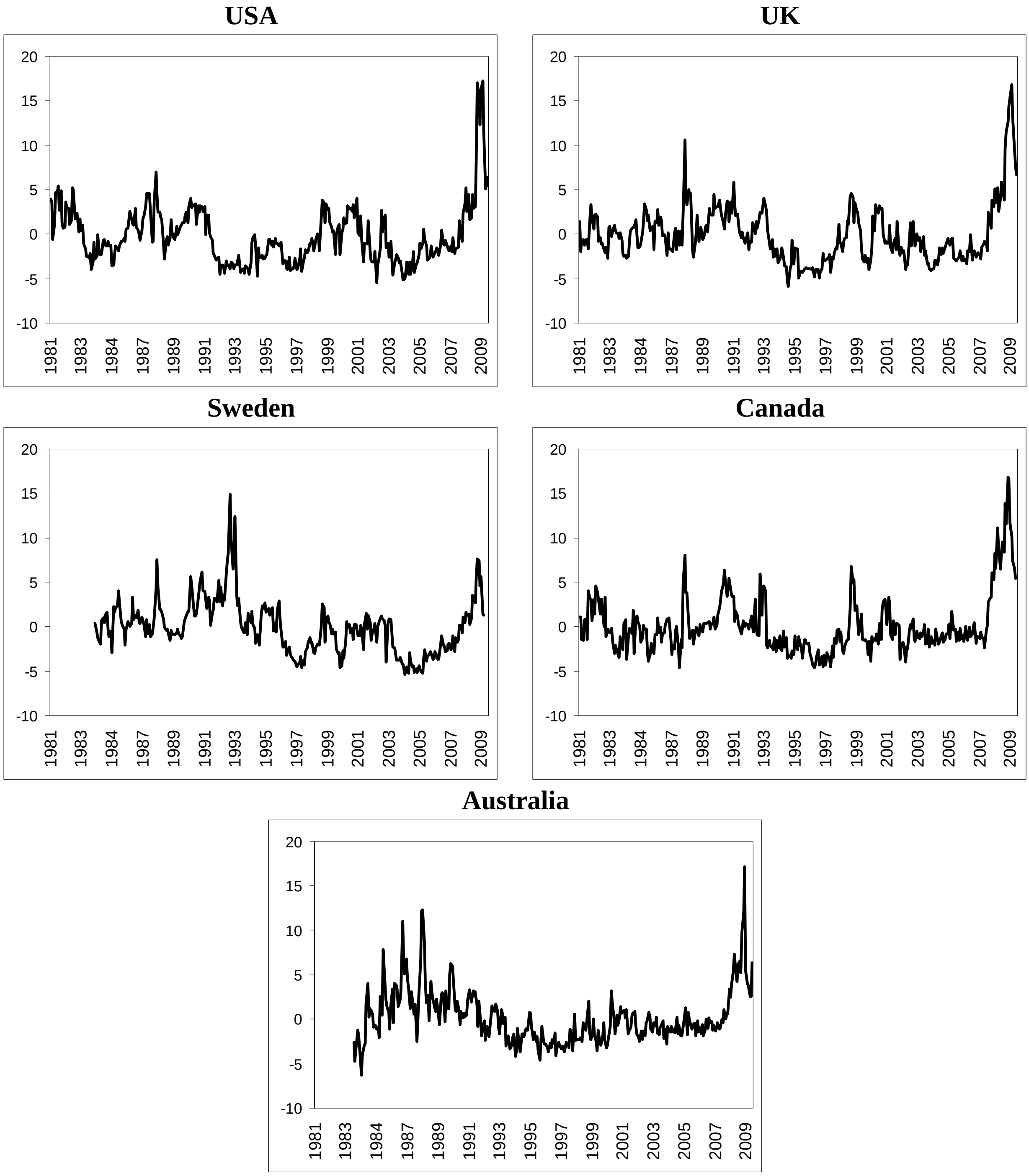
We proxy the financial stress by means of the FSI provided recently by the IMF (Cardarelli et al., 2009), which is a consistent measure for a wide range of countries but at the same time is comprehensive enough to track stress of a different nature. It includes the main components of financial stress in an economy and is available for a reasonably long period to be used for our empirical analysis (see Figure 2.1). We use both the overall index, which is a sum of seven components, as well as each sub-index and component separately:

- (i) Banking-related sub-index components: the inverted term spread (the difference between short-term and long-term government bonds), TED spread (the difference between interbank rates and the yield on treasury bills), banking beta (12-month rolling beta, which is a measure of the correlation of banking stock returns to total returns in line with the CAPM);
- (ii) Securities-markets related sub-index components: corporate bond spread (the difference between corporate bonds and long-term government bond yields), stock markets returns (monthly returns multiplied by -1), time-varying stock return volatility from the GARCH(1,1) model, and
- (iii) Foreign exchange related sub-index: the time-varying volatility of monthly changes in NEER, from the GARCH (1,1) model.

We have examined the various alternative methods of aggregating the components: simple sum, variance-equal weighting, and PCA weighting, but failed to uncover any systematic differences among these in terms of the values of the overall index and consecutively in the empirical results.

The use of a composite index has a number of benefits. First, it approximates the evolution of financial stress caused by different factors and thus it is not limited to one specific type of instability. Second, the inclusion of additional variables into the stress index does not affect the evolution of the indicator markedly (Cardarelli et al., 2009). Third, the composition of the indicator allows breaking down the reactions of the central bank with respect to different stress sub-components. Nevertheless, one has to be cautious about the interpretation. The composite indicator might suggest a misleading interpretation as long as the stress is caused by variables not included within the FSI but rather are highly correlated with some sub-component. An example is the case of Sweden during the ERM crisis. At the time of the crisis, Sweden maintained a fixed exchange rate and Riksbank sharply increased interest rates in order to sustain the parity. However, this is not captured by the exchange rate sub-component of the FSI, which measures the exchange rate volatility, because the volatility was actually close to zero. A closer look at the data shows that this period of stress is captured by the inverted term structure; hence it is incorrectly attributed to bank stress. A similar pattern can be observed for the UK, where the FSI increases after the announcement of the withdrawal from the ERM.

Figure 2.1: IMF Financial Stress Indicator



Note: The figure presents the evolution of the IMF stress index over time. Higher numbers indicate more stress (see Cardarelli et al., 2009).

2.3.2 The empirical model

Following Clarida et al. (1998, 2000), most empirical studies assume that the central bank sets the nominal interest rate in line with the state of the economy typically in a forward-looking manner:

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

where r_t^* denotes the targeted interest rate, $\overline{r}$ is the policy neutral rate[8], π_{t+i} stands for the central bank forecast of the yearly inflation rate i indicates periods ahead based on an information set Ω_t used for interest rate decision available at time t, and π_{t+i}^* is the central bank's inflation target.[9] y_{t+j} represents a measure of the output gap.

In addition, we allow for the interest rate smoothing of central banks, which is in empirical studies tracked by the simple partial-adjustment mechanism:

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

where $\rho \in \left[0,1\right]$ is the smoothing parameter. Although it seems plausible that the central banks adjust the interest rate sluggishly to the targeted value, there is an ongoing controversy as to whether this parameter represents genuine policy inertia or reflects empirical problems related to omitted variables, dynamics or shocks (see e.g. Rudebusch, 2006). The linear policy rule in Eq. (1) can be obtained as the optimal monetary policy rule in the LQ framework where the central bank aims only at price stability and economic activity. Bauducco et al. (2008) propose an NK model with a financial system where the monetary policy has privileged information (given its supervisory function) on the health of the financial sector. In such a setting, the common policy rule represented by Eq. (1) shall be augmented by variables representing the health of the financial sector. Following this contribution we consider the forward-looking rule where central banks may respond to a comprehensive measure of financial stress rather than stress in a particular segment (Bulíř and Čihák, 2008). Therefore, we substitute Eq. (2) into Eq. (1), eliminate unobserved forecast variables, pass the inflation forecast to the generic intercept α and include measures of the financial stress described above, which results in Eq. (3):

[8] A policy-neutral rate is typically defined as the sum of the real equilibrium rate and expected inflation.
[9] An explicit definition of an inflation target exists only for countries with an inflation targeting (IT) regime. Most empirical studies assume, in line with Taylor (1993), that this target does not vary in time and can be omitted in the empirical model.

$$r_t = (1-\rho)\left[\alpha + \beta(\pi_{t+i}) + \gamma y_{t+j}\right] + \rho r_{t-1} + \delta x_{t+k} + \varepsilon_t \qquad (3)$$

While in Eq. (1) the term α coincides with the policy neutral rate $\bar{r}$, its interpretation is not straightforward in our case. First, it is time varying as it encompass both changes in equilibrium real interest rate and inflation target. Second, the empirical model is augmented by additional variable, the measure of financial stress. Note that the financial stress index x_{t+k} does not appear within the square brackets because it is not a variable that determines the target interest rate r_t^* but it is rather a factor such as the lagged interest rate, i.e. it may explain why the actual interest rate r_t deviates from the target. Moreover, placing it in the regression on the same level as a lagged interest rate, we can directly test whether this variable representing ad-hoc policy decisions decreases the interest rate inertia ρ as suggested by Mishkin (2009). The common logic also suggest that the coefficients ρ and δ shall move in the opposite direction because the central bank either smoothes the interest rate changes or adjusts the rates in the face of financial stress. In the latter case, the response is likely to be quick and substantial. We set *i* equal to 2, *j* equal to 0 and *k* equal to -1.[10] 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).

The empirical studies on monetary policy rules have moved from using time-invariant estimates (Clarida et al., 1998) through sub-sample analysis (Taylor, 1999; Clarida et al., 2000) towards more complex methods that allow an assessment of the evolution in the conduct of monetary policy. There are two alternative methods to modeling structural changes in monetary policy rules that occur on an unknown date: (i) regime switching models, in particular the state-dependent Markov switching models (Valente, 2003; Assenmacher-Wesche, 2006; Sims and Zha 2006) and (ii) state-space models, where the changes are characterized by smooth transitions rather than abrupt switches (Boivin, 2006; Elkhoury,

[10] 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 in longer horizons. In case of the output gap, we instead assume a backward-looking reaction. The reason is that in the absence of real time data we have to rely on the output gap construction of statistical methods. It is arguable that besides the prediction error there is also a construction error that both might be magnified if an unobserved forecast is substituted by the output gap estimate for future periods. At last, we assume that central bankers' response (if any) to financial stress is rather immediate (see Mishkin, 2009). Therefore, we use one lag of the FSI and its subcomponents in the benchmark case. However, as a robustness check we allow for different lags and leads, allowing the central bankers' response to be preemptive rather than reactive.

2006; Kim and Nelson, 2006; Trecrocci and Vasalli, 2009). As argued in Baxa et al. (2010), we consider the second approach as preferable for the estimation of policy rules given that it is more flexible and allows the incorporation of a simple correction of endogeneity (Kim, 2006; Kim and Nelson, 2006), which is a major issue in forward-looking policy rules estimated from ex-post data.[11] The state-space approach or time-varying coefficient model seems also suitable when one wants to evaluate the effect of factors such as financial stress that can, for a limited length of time, alter (rather than permanently change) the monetary policy conduct.

The state-space models are commonly estimated by means of a maximum likelihood estimator via the Kalman filter or smoother. Unfortunately, this approach has several limitations that can turn problematic in applied work. First, the results are somewhat sensitive to the initial values of the parameters, which are usually unknown, especially in the case of variables whose impact on the dependent variable is not permanent and whose size is unknown, which is the case of financial stress and its effect on interest rates. Second, the log likelihood function is highly nonlinear and in some cases, optimization algorithms fail to minimize the negative of the log likelihood. In particular, it can either fail to calculate the Hessian matrix throughout the iterations process or, when the likelihood function is approximated to facilitate computations, covariance matrix of observation vector can get singular for provided starting values. The alternative is a moment-based estimator proposed by Schlicht (1981, 2005) and Schlicht and Ludsteck (2006), which is employed in our paper and briefly described below. This framework is flexible enough so as to incorporate the endogeneity correction proposed by Kim (2006).

Kim (2006) shows that the conventional time-varying parameter model delivers inconsistent estimates when explanatory variables are correlated with the disturbance term and proposes an estimator of the time-varying coefficient model with endogenous regressors. The endogeneity may arise not only in forward-looking policy rules based on ex-post data (Kim and Nelson, 2006, Baxa et al., 2010), but also in the case of variables that have a two-sided relation with monetary policy. Financial stress unquestionably enters this category. Following Kim (2006) we rewrite Eq. 3 as follows:

[11] The time-varying parameter model with the specific treatment of endogeneity is still relevant when real-time data are used (Orphanides, 2001). When the real-time forecast is not derived under the assumption that nominal interest rates will remain constant within the forecasting horizon (Boivin, 2006) or in the case of measurement error and heteroscedasticity (Kim et al., 2006).

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

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

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

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

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

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

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

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

$$x_{t+k} = Z'_{t-m}o + \sigma_\iota \iota_t,\ \iota_t \sim i.i.d.N(0,1) \tag{12}$$

The measurement Eq. (4) of the state-space representation is the monetary policy rule. The transitions in Eqs. (5)-(9) describes the time-varying coefficients as a random walk process without drift.[12] Eqs. (10)-(12) track the relationship between the potentially endogenous regressors (π_{t+i}, y_{t+j} and x_{t+k}) and their instruments, Z_t. We use the following instruments: π_{t-1}, π_{t-12} (π_{t-4} for CAN and SWE), y_{t-1}, y_{t-2}, r_{t-1}, and when k≥ 0 also x_{t-1} and x_{t-2}). Unlike Kim (2006), we assume that the parameters in Eqs. (10)-(12) are time-invariant. The correlation between the standardized residuals φ_t, ν_t and ι_t and the error term ε_t is $\kappa_{\varphi,\varepsilon}$, $\kappa_{\nu,\varepsilon}$ and $\kappa_{\iota,\varepsilon}$ respectively (note that σ_φ, σ_ν, and σ_ι are the standard errors of φ_t, ν_t and ι_t, respectively). The consistent estimates of the coefficients in Eq. (4) are obtained in two steps. In the first step, we estimate Eqs. (10)-(12) and save the standardized residuals φ_t, ν_t and ι_t. In the second step, we estimate Eq. (13) below along with Eqs. (5)-(9). Note that Eq. (13) now includes bias correction terms, i.e. (standardized) residuals from Eqs. (10)-(12), to address the aforementioned endogeneity of the regressors. Consequently, the estimated parameters in Eq. (13) are consistent, as ζ_t is uncorrelated with the regressors.

$$r_t = (1-\rho_t)\left[\alpha_t + \beta_t(\pi_{t+2}) + \gamma_t y_{t-1}\right] + \rho_t r_{t-1} + \delta_t x_{t-1} + \kappa_{\nu,\varepsilon}\sigma_\varepsilon \nu_t + \kappa_{\varphi,\varepsilon}\sigma_\varepsilon \varphi_t + \kappa_{\iota,\varepsilon}\sigma_{\varepsilon,t}\iota_t$$
$$+\zeta_t,$$
$$\zeta_t \sim N\left(0,(1-\kappa^2_{\nu,\varepsilon} - \kappa^2_{\varphi,\varepsilon} - \kappa^2_{\iota,\varepsilon})\sigma^2_{\varepsilon,t}\right) \tag{13}$$

[12] Note that while a typical time-invariant regression assumes that $a_t = a_{t-1}$, in this case it is assumed that $E[a_t] = a_{t-1}$.

As we noted before, instead of the standard framework for second-step estimation, the maximum likelihood estimator via the Kalman filter (Kim, 2006), we use an alternative estimation framework, the "varying coefficients" (VC) method (Schlicht, 1981; Schlicht, 2005; Schlicht and Ludsteck, 2006). This method is a generalization of the ordinary least squares approach that, instead of minimizing the sum of the squares of residuals $\sum_{t=1}^{T}\zeta^2$, uses the minimization of the weighted sum of squares:

$$\sum_{t=1}^{T}\zeta^2+\theta_1\sum_{t=1}^{T}\vartheta_1^2+\theta_2\sum_{t=1}^{T}\vartheta_2^2+\ldots+\theta_{\mathrm{n}}\sum_{t=1}^{T}\vartheta_{\mathrm{n}}^2 \tag{14}$$

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

- (i) it does not require knowledge of initial values even for non-stationary variables prior to the estimation procedure. Instead, both the variance ratios and the coefficients are estimated simultaneously,
- (ii) the property of the estimator, that the time averages of estimated time-varying coefficients are equal to its time-invariant counterparts, permits the easy interpretation of the results in relation to time-invariant results,
- (iii) it coincides with the MLE estimator via the Kalman filter if the time series are sufficiently long and if the variance ratios are properly estimated.

However, this method suffers certain limitations of its own. In particular: (a) it requires that the time-varying coefficients are described as random walks and (b) the shocks in time-varying coefficients ϑ_t are minimized (see Eq. (14)). While this does not represent a major problem for the estimation of the coefficients of common variables such as inflation, where the monetary policy response is permanent, it can lead to a loss of some information about ad-hoc response factors in monetary policy-making that are considered by central bankers only infrequently but once they are in place the policy response can be substantial. A financial

[13] See Schlicht and Ludsteck (2006) and Baxa et al. (2010) for more details.

stress indicator x_{t+k} seems to be this kind of factor. The way to deal with this problem is the estimation-independent calibration of the variance ratios in Eq. (14) such that the estimated coefficient is consistent with economic logic, i.e. it is mostly insignificant and it can turn significant (with no prior restriction on its sign) during the periods of financial stress, i.e. when the financial stress indicator is different from zero. Therefore, we first estimate Eq. (13) using the VC method first and study whether the resulting coefficients at the FSI correspond to economic intuition, and especially whether the coefficient is not constant or slowly moving (a so-called pile-up problem; see Stock and Watson, 1998). When this problem occurs, we compare the results with models where *k* belongs to (-2, -1, 0, 1, 2) and calibrates the variance ratios in Eq. (13) by the variance ratios estimated for the model with the largest variances in the FSI. This step was necessary for Australia and Sweden. The Taylor rule coefficients were compared with the initial estimates and they were consistent in both cases.[14]

The results of our empirical analysis should reveal whether central banks adjusted its interest rate policy in the face of financial stress. However, the time-varying framework also allows inferring whether any response to financial stress led to the temporal dismissal of other targets, in particular the inflation rate. Therefore, we are mainly interested in the evolution of the financial stress coefficient δ_t. We expect it to be mostly insignificant or zero given that episodes of financial stress are rather infrequent and even if they occur the monetary authorities may not always respond to them. Moreover, the size of the estimated coefficient does not have any obvious interpretation since the FSI is a composite indicator normalized to have a zero mean. Consequently, we define the stress effect as a product of the estimated coefficient δ_t and the value of IMF financial stress index x_{t+k}. The interpretations of the stress effect is straightforward: it shows the magnitude of interest rate reactions to financial stress in percentage points or, in other words, the deviation from the target interest rate as implied by the macroeconomic variables due to the response to financial stress.

[14] Stock and Watson (1998) propose a medium unbiased estimator for variance in the time-varying parameter model but its application is straightforward only in the case of one time-varying coefficient and, more importantly, it requires the variables being stationary.

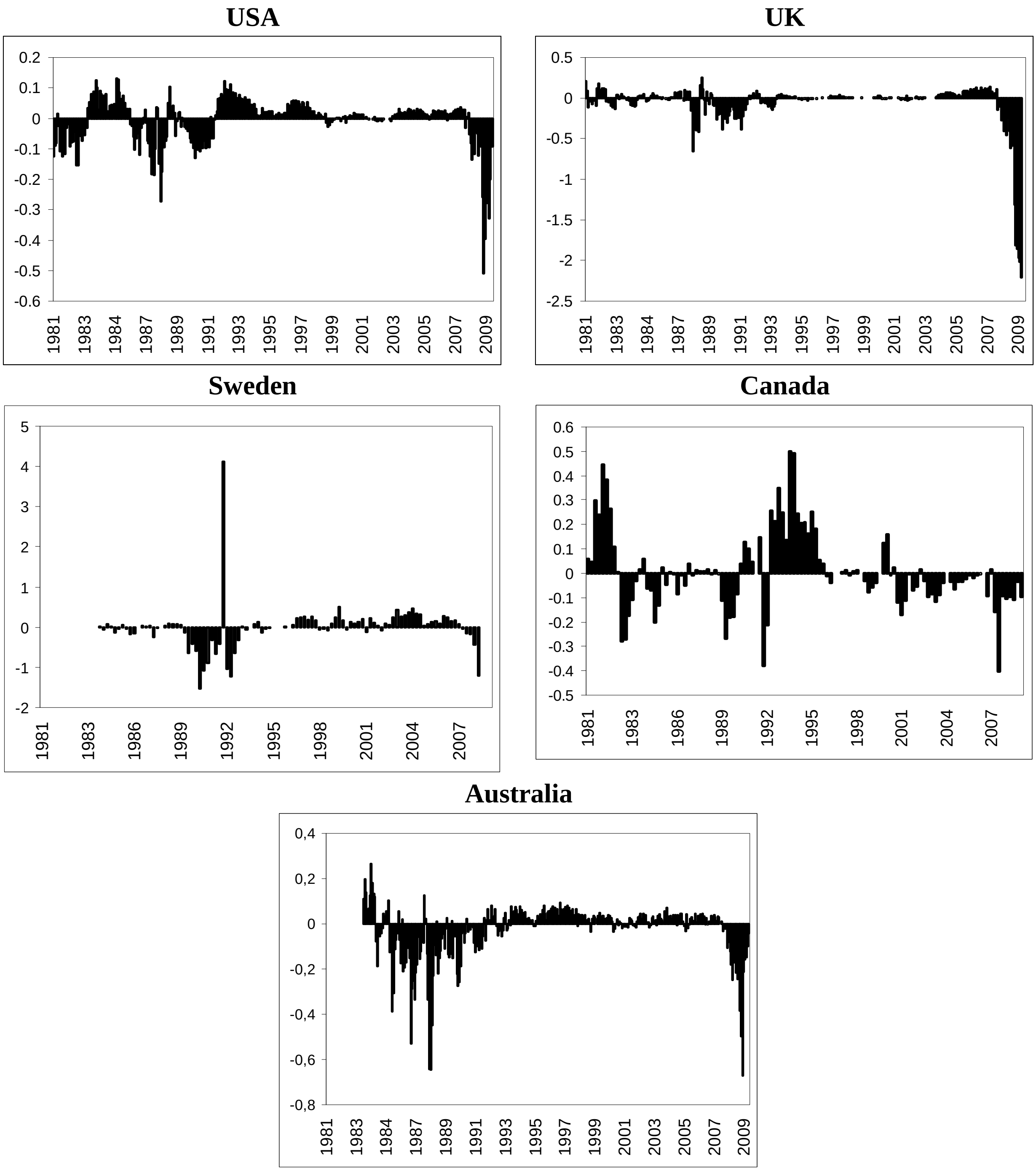
2.4 Results

This section summarizes our results on the effect of financial stress on interest rate setting. First, the results on the effect of the overall measure of financial stress on interest rate setting are presented. Second, the effect of specific components of financial stress on monetary policy is examined. Third, we briefly comment on the monetary policy rule estimates that served as the input for the assessment of financial stress effects. Finally, we perform a series of robustness checks.

Figure 2.2 presents our results on the effect of financial stress on interest rate setting in all five countries (labeled as *financial stress effect* hereinafter). Although there is some heterogeneity across countries, some global trends on the effect of financial stress are apparent. While in good times, such as in the second half of the 1990s, financial stress has virtually no effect on interest rate setting or it is slightly positive,[15] the reaction of monetary authorities to financial stress was highly negative during the 2008-2009 global financial crisis. While the previous evidence on the effect of financial stress on monetary policy is somewhat limited, our results broadly confirm the time-invariant findings of Cecchetti and Li (2008), which show that the US Fed adjusted the interest rates to the procyclical impact of bank capital requirements in 1989-2000. Similarly, Belke and Klose (2010) estimate the Taylor rule on two sub-samples (before and during the 2008-2009 global financial crisis) and find that the Fed reacted systematically not only to inflation and output gap, but to asset prices, credit and money as well.

[15] Note that the positive effect of financial stress on interest rate setting is to some extent a consequence of scaling the financial stress indicator; its zero value corresponds to a long run average stress. Hence, we do not place much attention onto the positive values of stress unless caused by the temporarily positive and significant regression coefficient associated with the FSI.

Figure 2.2: The Effect of Financial Stress on Interest Rate Setting



Notes: The figure depicts the evolution of the financial stress effect. The stress effect (y-axis) is defined as the product of the estimated coefficient on the financial stress indicator in monetary policy rule and the value of the IMF financial stress indicator. The stress effect shows the magnitude of the interest rate reaction to financial stress in percentage points.

The size of financial stress effects on interest rate setting during the recent financial crisis is somewhat heterogeneous with the strongest reaction found for the UK The results suggest that all central banks, except the Bank of England, kept its policy rates by about 50-100 basis points lower, as compared to the counterfactual of no reaction to financial stress. The size of this effect for the UK is assessed to be about three times stronger (i.e. 250 basis points). This implies that about 50 percent of the overall policy rate decreased during the recent financial crisis was motivated by the financial stability concerns in the UK (10-30 percent in the remaining sample countries), while the remaining half falls on unfavorable developments in domestic economic activity. This finding is interesting when confronted with the BoE's very low consideration of expected inflation over the last decade (found Baxa et al., 2010, using the time-varying model and in Taylor and Davradakis, 2006, in the context of the threshold model) that further decreased during the current crisis similarly as the USA but less so for the other central banks. It is also evident that the magnitude of the response is unusual for all five central banks. Yet, the results for Australia, Canada and Sweden show that a similar magnitude of the response to financial stress recorded during the recent financial crisis was already seen in previous periods of high financial stress.

Given that the 2008-2009 global crisis occurred right at the end of our sample (there is a peak in the stress indicator of 5 standard deviations that has not returned to normal values yet), we have performed an additional check to avoid possible end-point bias. In particular, we have run our estimation excluding the observation from the period of 2008-2009 crisis: These results (are available upon request) confirmed the robustness of the reported findings. With regards to the effect of the current crisis, the largest uncertainty is associated with the results for Canada, for which the shortest data sample was available, and it ends in the fourth quarter of 2008. When the possibility of a preemptive reaction of the central bank to financial stress is considered, the effect of financial stress in the current crisis is somewhere at 1-2 percent (see Appendix 3). These additional results suggest the response of the Bank of Canada is rather underestimated.

The question of which components of financial stress influence interest rate setting is addressed in Figure 2.3. Some heterogeneity across countries is again apparent; although it seems that bank stress and stock market stress dominated the central bankers in less open economies. On the other hand, exchange rate stress matters in more open economies such as Canada and Sweden.

More specifically, the US Fed seemed to be worried about financial instability, especially during the 1980s. We can see that the main concern in the early 1980s was banking stress, which is arguably related to the Savings and Loans crisis. Another concern was that of stock market stress in particular during the stock market crash of 1987 when interest rates were lower by 30 b.p. with respect to the benchmark case.

The Bank of England was, in general, much more perceptive to financial stress. We find its response mainly to stock market stress again notably in 1987. Interestingly, we find little response to exchange rate stress, not even during the 1992 ERM crisis. Nevertheless, it has to be emphasized that the interest rate reaction to speculative attack was subdued in comparison to, for example, the Riksbank (Buiter et al., 1998). The coefficient at the FSI remains constant until the devaluation of the pound sterling in September 1992. Since then the effect of financial stress on interest rate setting approaches zero from originally negative values. Besides this, the response of the Bank of England to inflation has decreased. From this perspective, it seems the pound sterling's withdrawal from the ERM allowed for both a more rule-based and less restrictive monetary policy.

The reaction of the Riksbank to the ERM crisis was different. First, after a series of speculative attacks on the Swedish krona in mid-September 1992, the Riksbank still tried to keep the current fixed exchange rate in place and the marginal interest rate jumped up 500 percent in order to offset the outflow of liquidity and other speculative attacks (see the large positive stress effect on the interest rate in 1992 in Figure 2.2). However, not even such an increase was sufficient and the fixed exchange rate had to be abandoned later in November. [16]

The Reserve Bank of Australia significantly loosened its policy during the 1980s, which can be attributed to the stress in the banking sector with an exception of the reaction to the stock market crash in 1987 (see Figure 2.3).

The exchange rate as well as bank stress seems to matter for interest rate considerations at the Bank of Canada. Interestingly, the results suggest that the Bank of Canada often responded to higher exchange rate stress by monetary tightening. A possible explanation for this finding

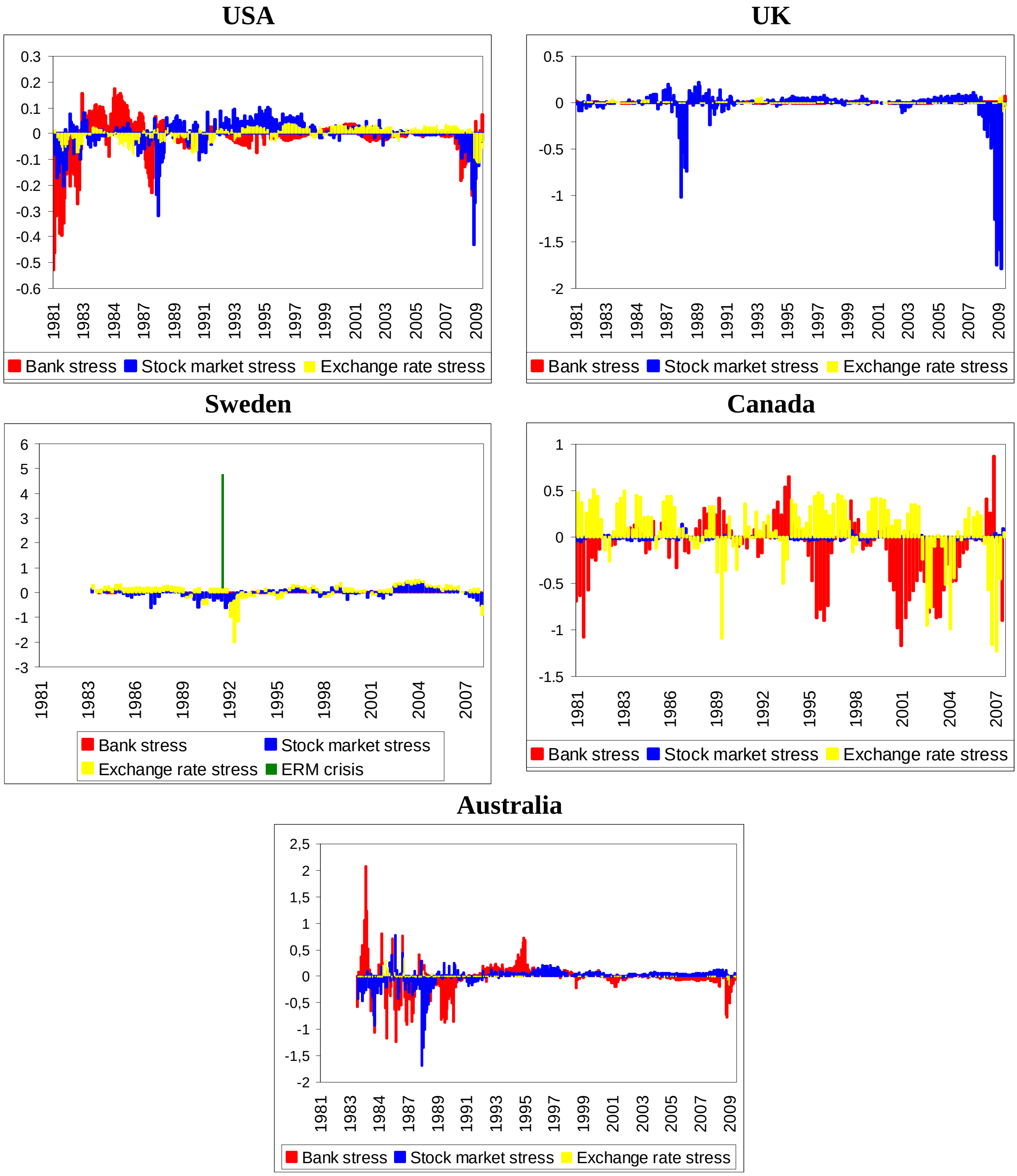
[16] For Sweden, we add a dummy variable for the third quarter of 1992 (ERM crisis) to Eq. (13). At this time the Swedish central bank forced upward short-term interest rates in an effort to keep the krona within the ERM. From the perspective of our model, it was a case of a strong positive reaction to the actual stress that lasted only one period. When this dummy variable was not included, the model with a lagged value of the FSI was unable to show any link between stress and the interest and estimates of other coefficients were inconsistent with economic intuition.

could be that the Canadian central bank tightened the policy when the currency stabilized at the level that the monetary authority considered to be undervalued.

We would like to highlight a comparison of Figures 2.2 and 2.3. First, it should be noted that the positive response to one stress subcomponent may cancel out against a negative response to another one, making the response to the overall stress negligible (as in the case of Canada). Second, the stress effects related to individual subcomponents need not necessarily sum up to the stress effect related to the entire FSI.

All in all, the results suggest that the central bank tends to react to financial stress and different components of financial stress matter in different time periods. The effect of financial stress on interest rate setting is found to be virtually zero in good times and economically sizable during the period of high financial stress.

Figure 2.3: The Effect of Financial Stress Components on Interest Rate Setting: Bank Stress, Exchange Rate Stress and Stock Market Stress



Notes: The figure depicts the evolution of the components of the financial stress effect, namely bank stress effect, exchange rate stress effect and stock market stress effect. The stress effect (y-axis) is defined as the product of the estimated coefficient on the given component of the financial stress indicator in monetary policy rule and the value of the corresponding component of the IMF financial stress indicator. The stress effect shows the magnitude of the interest rate reaction to financial stress in percentage points.

Next, we briefly comment on our monetary policy rule estimates. The plot of the evolution of estimated parameters over time for all countries is available in Appendix 1. First, these results indicate that the interest rate smoothing is much lower than what time-invariant estimates of monetary policy rules typically suggest (see for example, Clarida et al., 1998, 2000). Our estimates of interest rate smoothing seem to be reasonable given the recent critique of Rudebush (2006), who argues that the degree of interest rate smoothing is rather low. Moreover, for some central banks such as the RBA and the BoE in 2010 or the Sveriges Riksbank in late 1980 we find support for Mishkin's (2009, 2010) argument that central banks are less inertial during the crisis.[17] Second, the response of interest rates on inflation is particularly strong during the periods when central bankers want to break the record of high inflation such as in the UK or in Australia at the beginning of 1980s and it becomes less aggressive in the low inflation environment with subdued shocks. Third, some central banks are also found to react to output gap developments with the estimated parameter to be slightly positive on average (see more detailed results in Baxa et al., 2010).

Finally, we perform a battery of robustness checks. First, following the argument put forward above, we use interbank interest rates as a dependent variable. These results are reported in Figures A.2.1-A.2.2. We can see that the overall stress effect on the interbank rate is larger for the US during the current crisis, where it explains 2 percent of the decrease of the interbank interest rate. For Sweden, we have found a strong positive effect of exchange rate volatility in the late 1980s; this might be linked to the aim of the central bank to keep the exchange rate fixed. In other cases, there is no substantial difference between the benchmark results and results obtained using this alternative dependent variable.

Second, in the benchmark model and all the results reported so far we use the first lag of the FSI in policy rule estimation. We stimulate this choice by the use of monthly data, the frequency of monetary policy meetings of most central bank boards and the assumption that the policy actions are likely to be implemented in a timely fashion. In addition, we employ different lags and leads, in the latter case allowing the policy to be preemptive rather than reactive. In this case, we use the future realized value of the FSI as a proxy of a central bank's expectation (in a similar manner as to how it is routinely executed for inflation expectations) and consequently treat the FSI as an endogenous variable (see Figure A.3.1 for the results). In order to get comparable results, we calibrate the variance ratios by the same values as in the baseline specification. Although we find rather mixed evidence on preemptive policy actions,

[17] The correlation coefficient of the estimated time-varying coefficient of a lagged interest rate ρ and the financial stress index δ is -0.79 for Australia, 0.21 for Canada, -0.20 for Sweden, -0.68 for the UK and 0.60 for the US

which can be also related to the inadequacy of proxying the expected values of financial stress by actual values of the financial stress indicator as well as the fact that a central bank might not react to the stress preemptively, the reaction to financial stress in the current crisis is strongly negative for both expected or observed stress.

Third, we further break down the FSI sub-indices to each underlying variable to evaluate their individual contribution.[18] The corresponding stress effects appear in Figures A.4.1-A.4.2. Breaking down stock-market related stress, we find that the US Fed and the BoC react to the corporate bond spread, whereas the BoE and Sveriges Riksbank are more concerned by stock returns and volatility. While the RBA seem to be concerned with both corporate bond spreads and stock market volatility in the 1980s, the role of stock-related stress had substantially decreased by then. As far as bank-related stress is concerned, the TED spread plays a major role in all countries apart from the UK where the largest proportion of the effect on the interest rate can be attributed to an inverted term structure.

Fourth, since the verifications related to our econometric framework to obvious alternatives such as, first, the use of a maximum likelihood estimator via the Kalman filter instead of the moment-based time-varying coefficient framework of Schlicht and, second, the use of a Markov-switching model instead of a state-space model were provided in Baxa et al. (2010), we estimate simple time-invariant monetary policy rules for each country by the generalized method of moments, including various subsamples. This simple evidence reaffirms that the analyzed central banks seem to pay attention to overall financial stress in the economy. The FSI is statistically significant, with a negative sign of a magnitude between 0.05-0.20 for all countries. On the other hand, the coefficient of its sub-components often are not, and the exchange rate subcomponent in some cases has a positive sign. These results, which are available upon request, confirm that to understand the interest rate adjustment to financial stress, one should rely on a model allowing a differential response across time.

[18] This applies only to the banking and the stock market sub-components because the foreign exchange sub-component is represented by a single variable.

2.5 Concluding Remarks

The 2008-2009 global financial crisis awoke a significant interest in exploring the interactions between monetary policy and financial stability. This paper aimed to examine in a systematic manner whether and how the monetary policy of selected main central banks (US Fed, the Bank of England, Reserve Bank of Australia, Bank of Canada and Sveriges Riksbank) responded to episodes of financial stress over the last three decades. Instead of using individual alternative measures of financial stress in different markets, we employed the comprehensive indicator of financial stress recently developed by the International Monetary Fund, which tracks overall financial stress as well as its main subcomponents, in particular banking stress, stock market stress and exchange rate stress.

Unlike a few existing empirical contributions that aim to evaluate the impact of financial stability concerns on monetary policymaking, we adopt a more flexible methodology that allows for the response to financial stress (and other macroeconomic variables) to change over time as well as deals with potential endogeneity (Kim and Nelson, 2006). The main advantage of this framework is that it does not only enable testing whether the central banks responded to financial stress at all, but it also detects the periods and types of stress that were the most worrying for monetary authorities. Our results indicate that central banks truly change their policy stance in the face of financial stress, but that the magnitude of such responses varies substantially over time. As expected, the impact of financial stress on interest rate setting is essentially zero most of the time when the levels of stress are very moderate. However, most central banks loosen monetary policy when the economy faces high financial stress. There is a certain cross-country and time heterogeneity when we look at central banks' considerations of specific types of financial stress. While most central banks seem to respond to stock market stress and bank stress, exchange rate stress is found to drive the reaction of central banks only in more open economies

Consistently with our expectations, the results indicate that a sizeable fraction of monetary policy easing during the 2008-2009 financial crisis can be explained by a direct response to the financial stress above what could be attributed to a decline in inflation expectation and output below its potential. Although, the size of the financial stress effect differs by country. The result suggests that all central banks, except the Bank of England, kept their policy rates at 50 basis points lower on average solely due to financial stress present in the economy.

Interestingly, the size of this effect for the UK is assessed at about five times stronger (i.e. 250 basis points). This implies that about 50 percent of the overall policy rate decrease during the recent financial crisis was motivated by financial stability concerns in the UK (10-30 percent in the remaining sample countries), while the remaining half falls on unfavorable developments in domestic economic activity. For the US Fed, the macroeconomic developments themselves (a low-inflation environment and an output substantially below its potential) explain the major fraction of the policy interest rate decreases during the crisis, leaving the further response to financial stress to be constrained by zero interest rate bound.

All in all, our results point to the usefulness of augmenting the standard version of monetary policy rule by some measure of financial conditions to get a better understanding of the interest rate setting process, especially when financial markets are not stable. The empirical results suggest that the main central banks considered in this study altered the course of their monetary policy in the face of financial stress. The recent crisis seems truly to be an exceptional period in the sense that the response to financial instability was substantial and coincided in all analyzed countries, which is evidently related to intentional policy coordination absent in previous decades. However, we have also seen that previous idiosyncratic episodes of financial distress were, at least in some countries, followed by monetary policy responses of similar if not higher magnitude.

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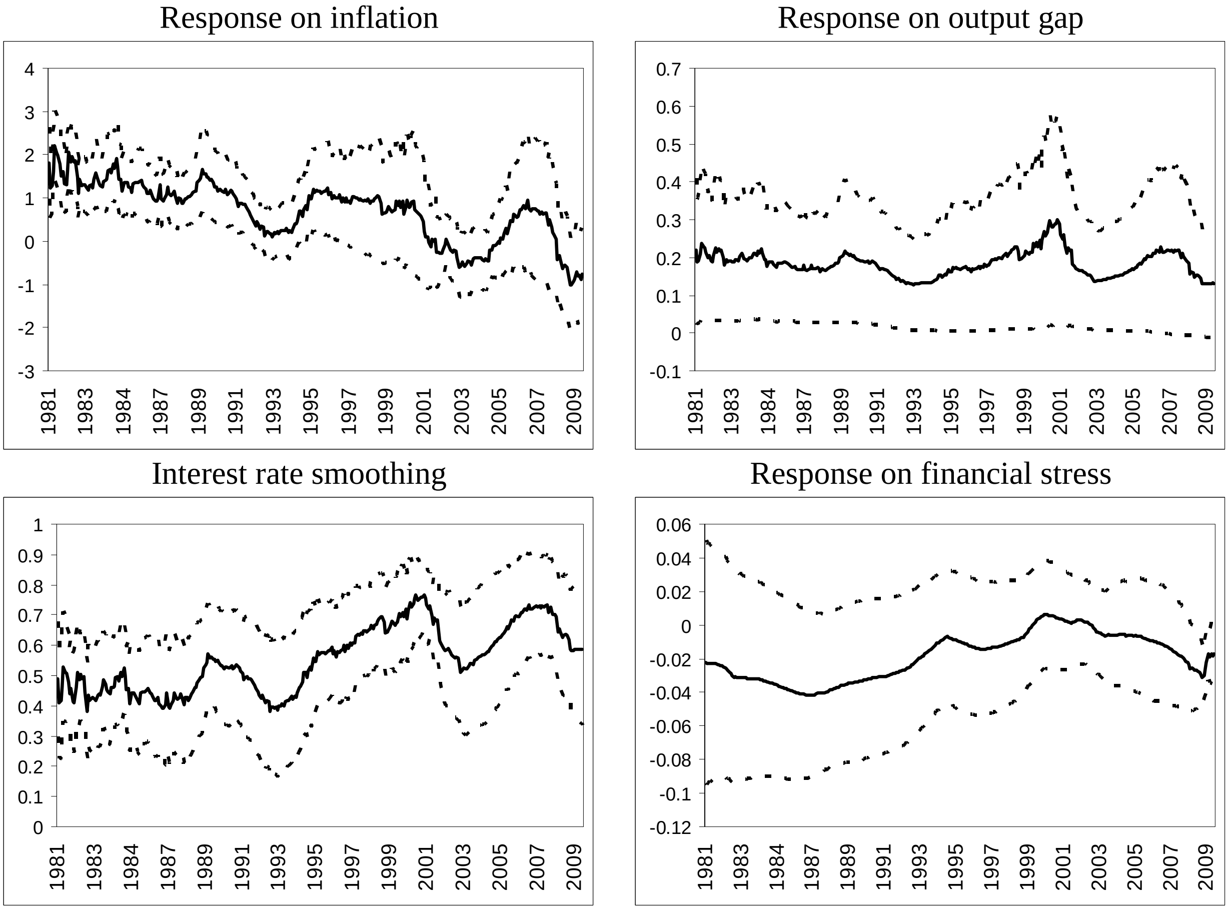
Appendix

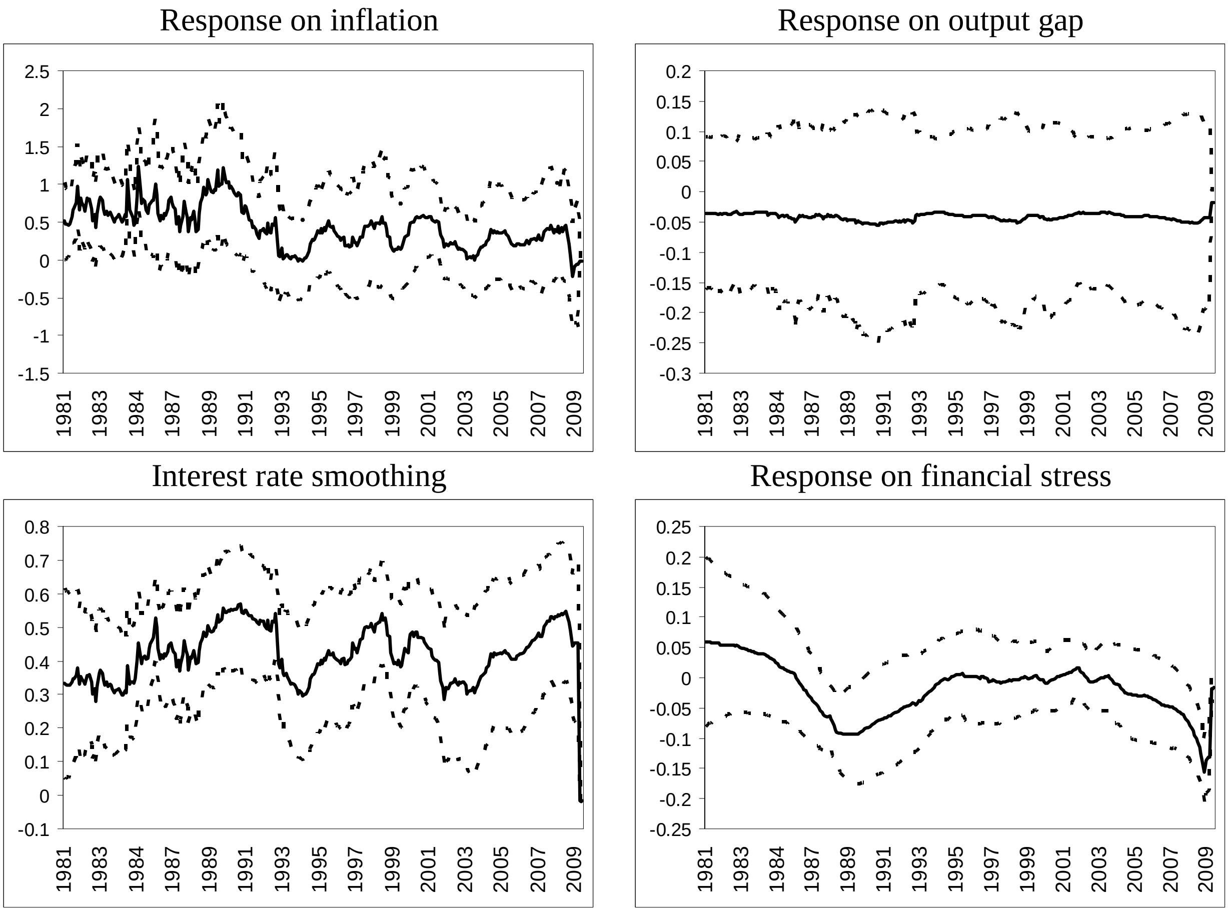
A.2.1 Estimates of other coefficients of time-varying monetary policy rule

Figure A.2.1.1: Time-Varying Monetary Policy Rules: US



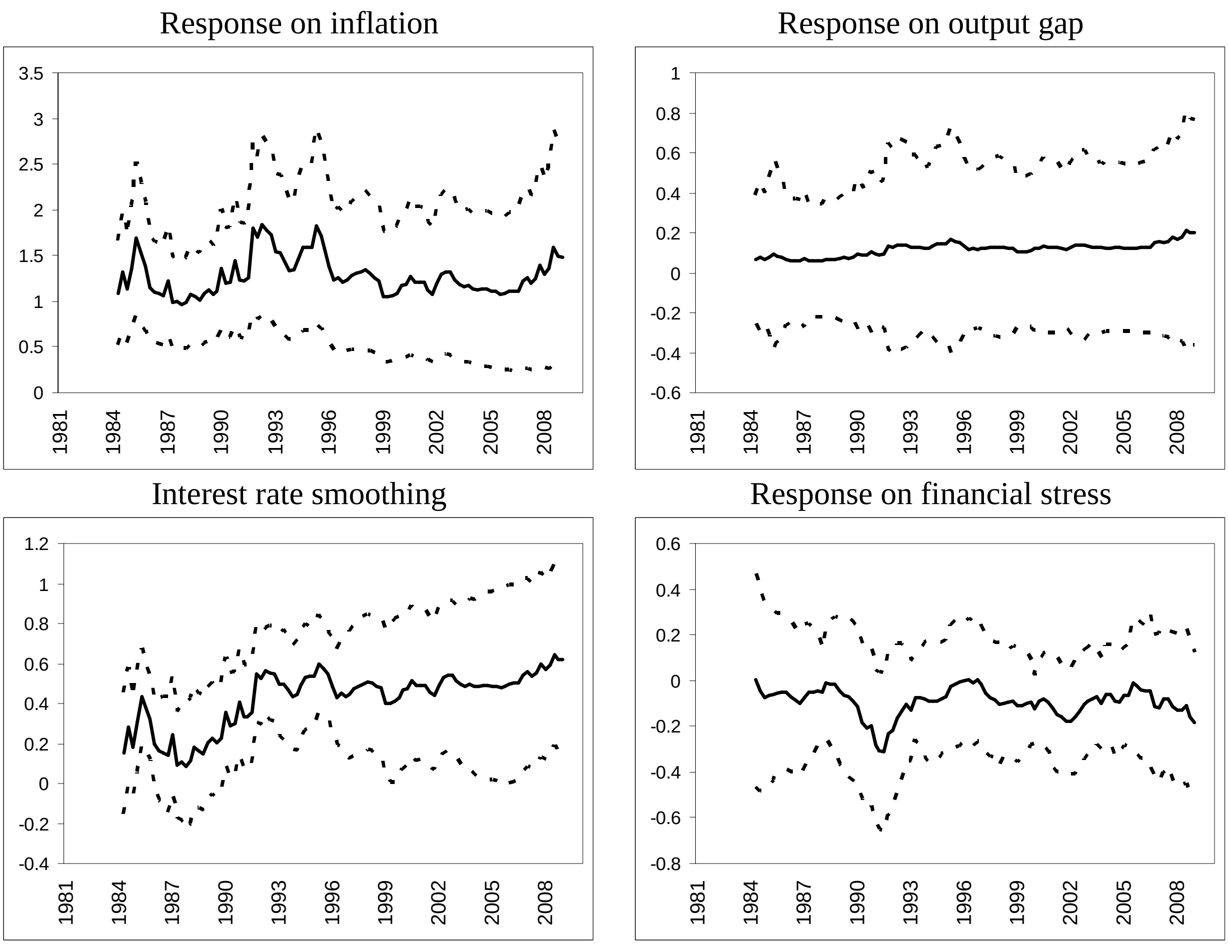
Note: The estimated coefficients of time-varying monetary policy rule are depicted with a 95% confidence interval.

Figure A.2.1.2: Time-Varying Monetary Policy Rules: UK



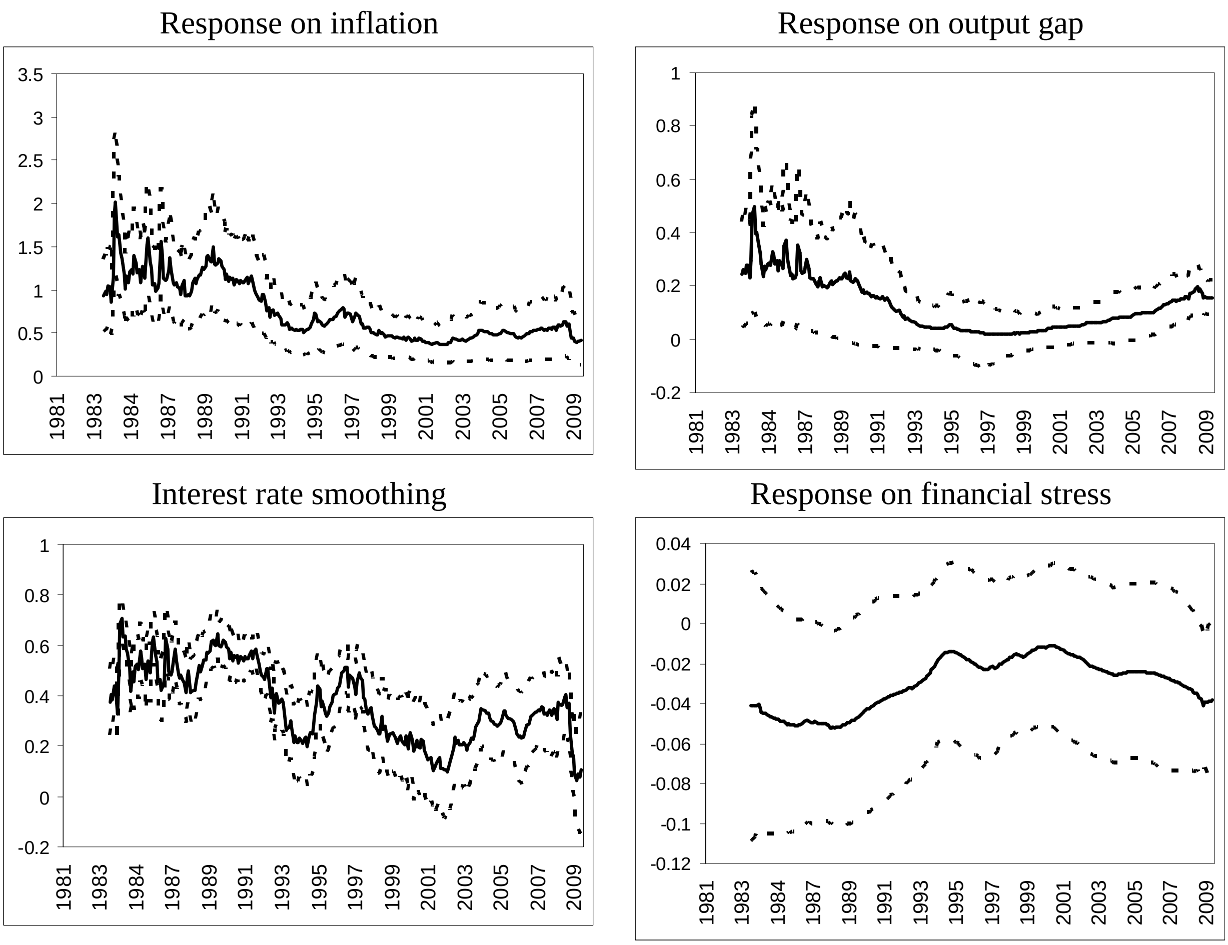
Note: The estimated coefficients of time-varying monetary policy rule are depicted with a 95% confidence interval.

Figure A.2.1.3: Time-Varying Monetary Policy Rules: Sweden



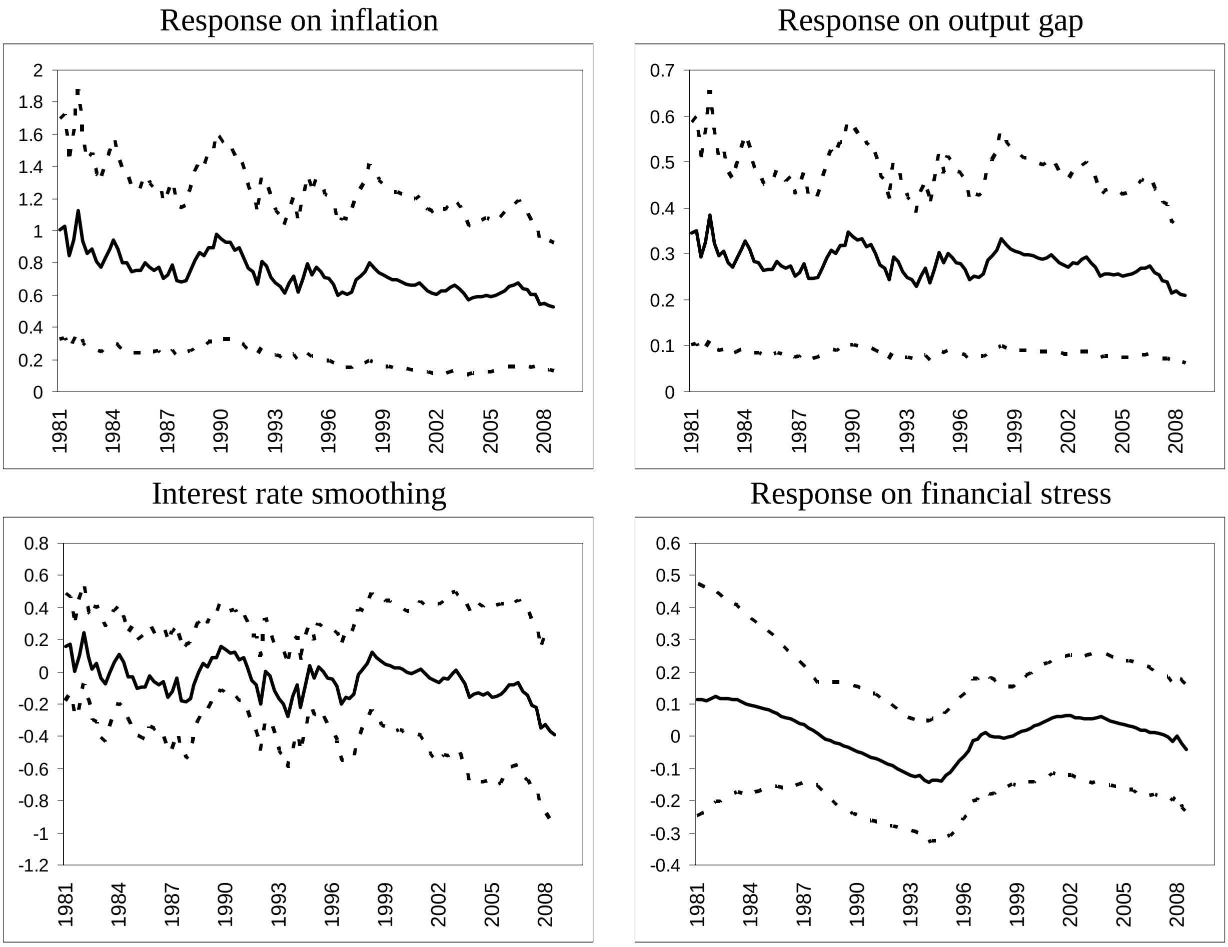
Note: The estimated coefficients of time-varying monetary policy rule are depicted with a 95% confidence interval.

Figure A.2.1.4: Time-Varying Monetary Policy Rules: Australia



Note: The estimated coefficients of time-varying monetary policy rule are depicted with a 95% confidence interval.

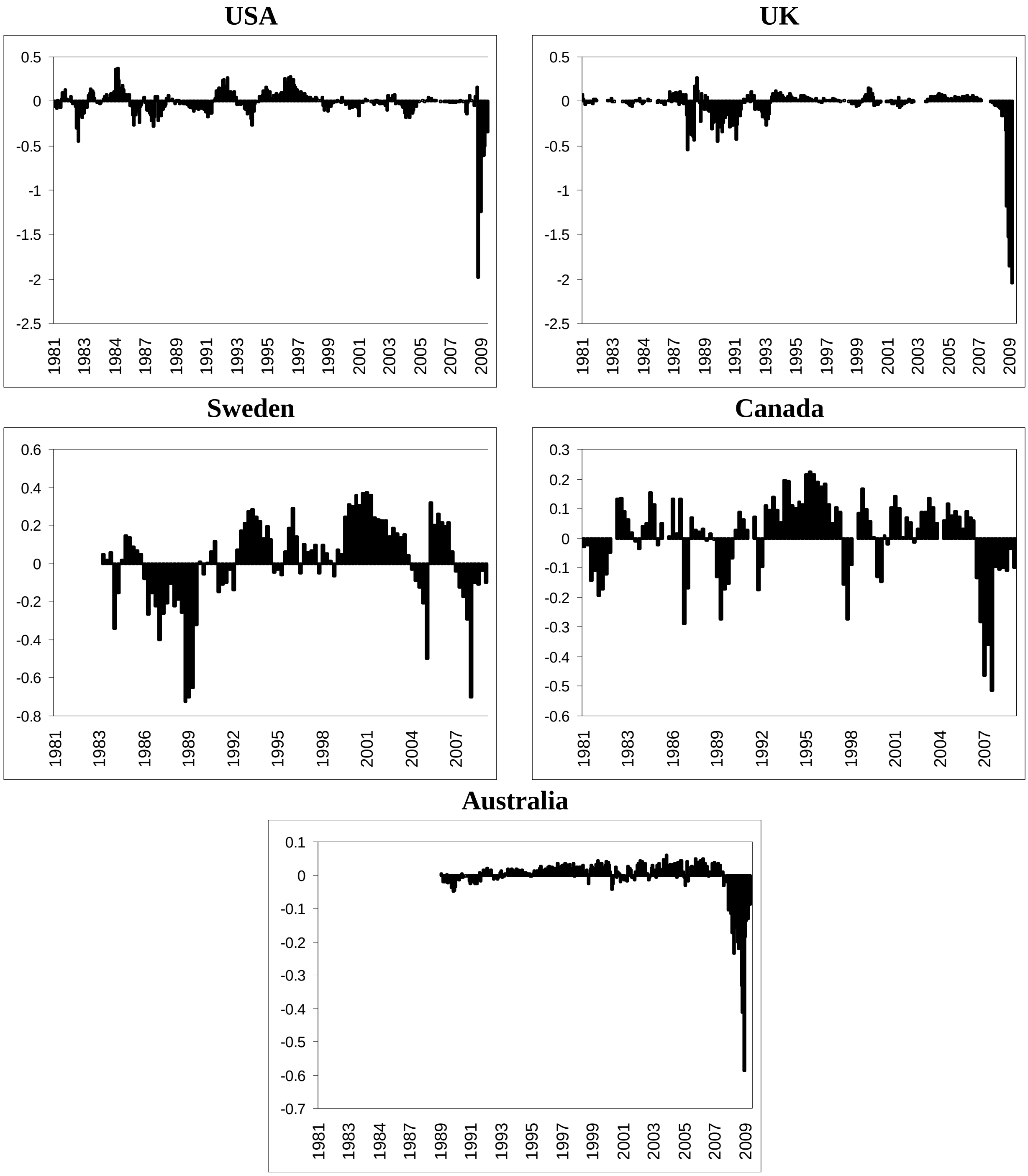
Figure A.2.1.5: Time-Varying Monetary Policy Rules: Canada



Note: The estimated coefficients of time-varying monetary policy rule are depicted with a 95% confidence interval.

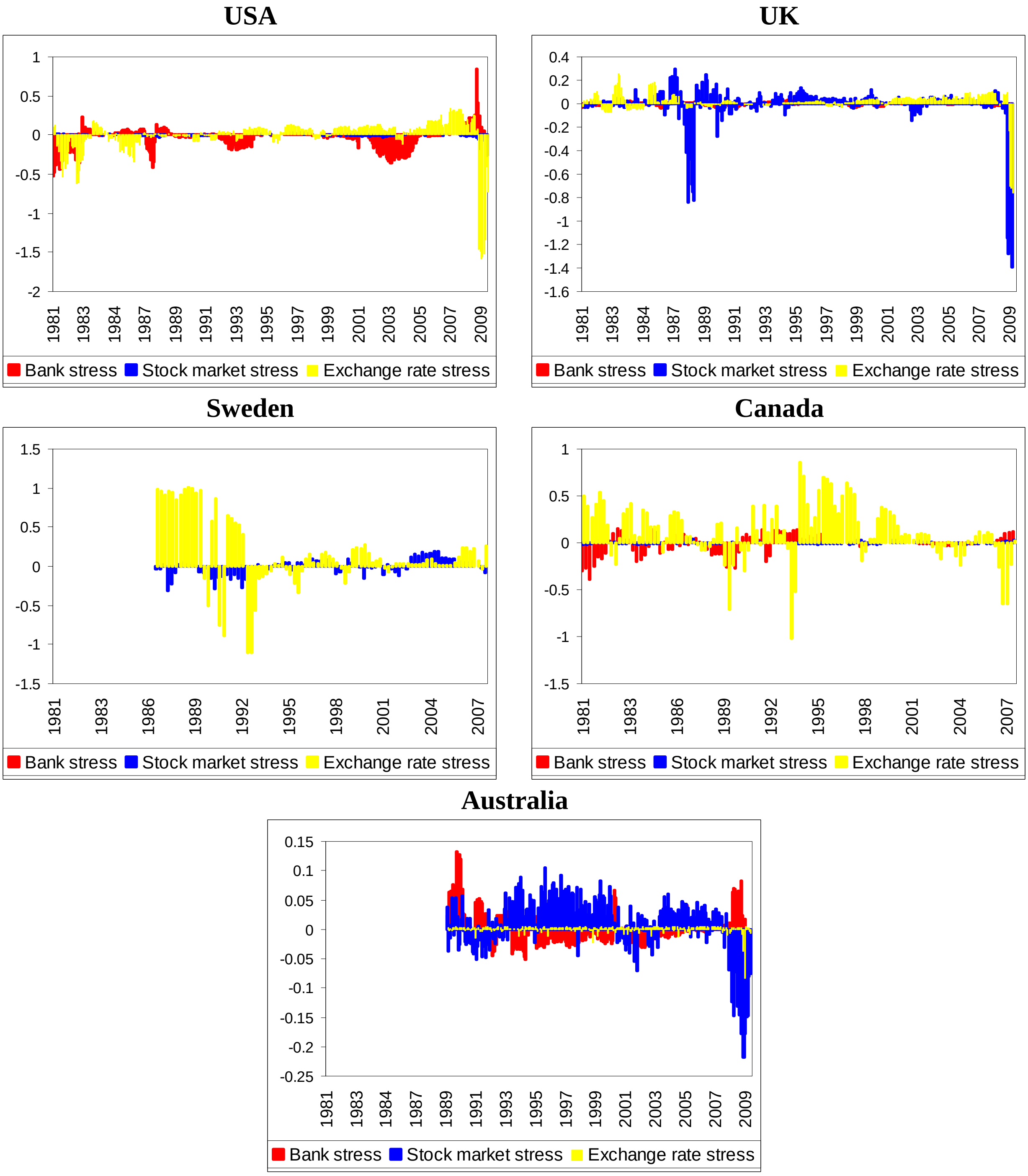
A.2.2 The results with the interbank rate as the dependent variable in the policy rule

Figure A2.2.1: The Effect of Financial Stress on Interest Rate Setting



Notes: The figure depicts the evolution of the financial stress effect. The stress effect (y-axis) is defined as the product of the estimated coefficient on the financial stress indicator in monetary policy rule and the value of the IMF financial stress indicator. The stress effect shows the magnitude of the interest rate reaction to financial stress in percentage points.

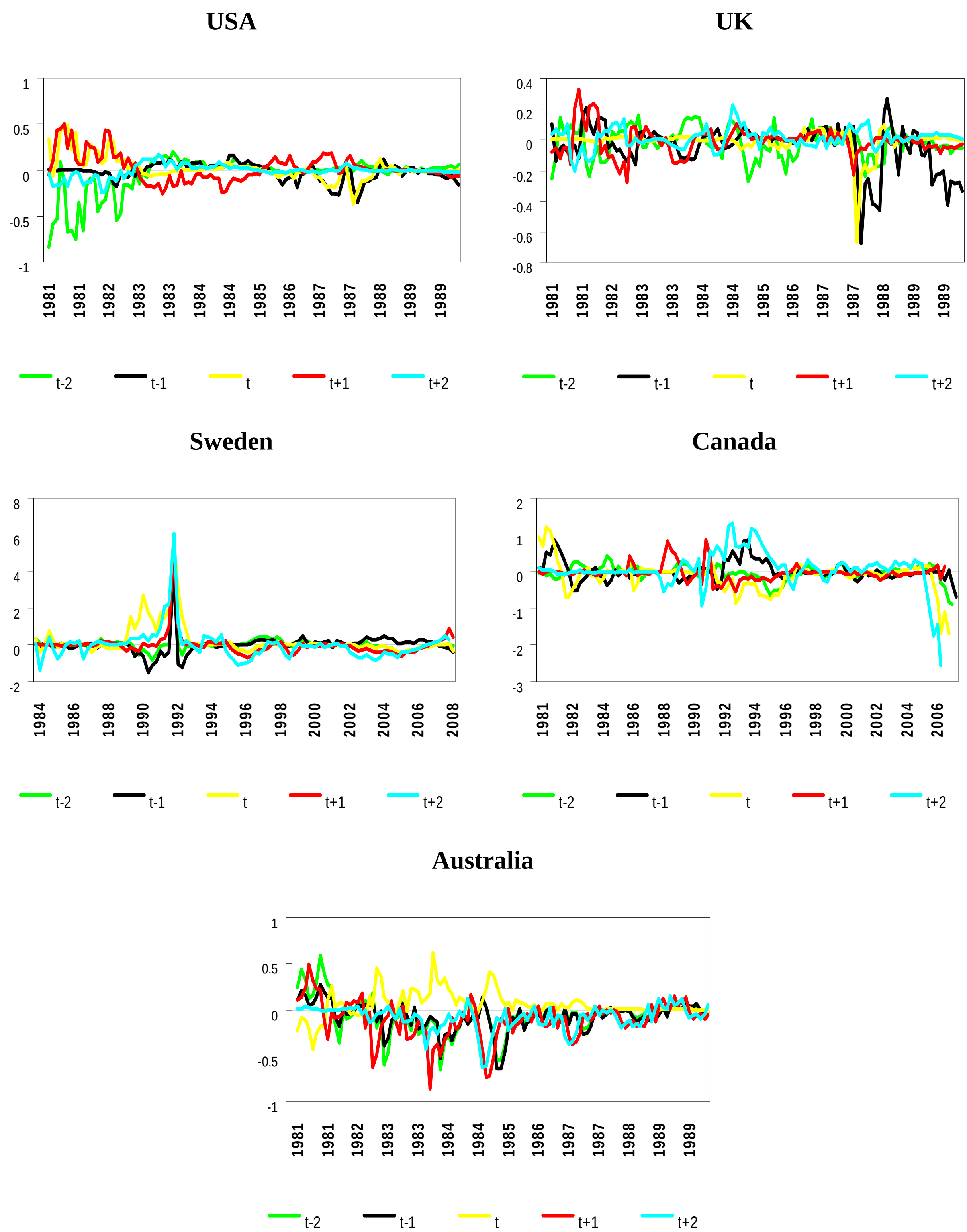
Figure A.2.2.2: The Effect of Financial Stress Components on Interest Rate Setting: Bank Stress, Exchange Rate Stress and Stock Market Stress



Notes: The figure depicts the evolution of the components of the financial stress effect, namely bank stress effect, exchange rate stress effect and stock market stress effect. The stress effect (y-axis) is defined as the product of the estimated coefficient on the given component of the financial stress indicator in monetary policy rule and the value of the corresponding component of the IMF financial stress indicator. The stress effect shows the magnitude of the interest rate reaction to financial stress in percentage points.

A.2.3 The results with different leads and lags of the FSI

Figure A.2.3.1: The Effect of Financial Stress (t-1 vs. t-2, t, t+1, t+2) on Interest Rate Setting



A.2.4 The Results with Individual Variables of Bank Stress and Stock Market Stress

Figure A.2.4.1: The Effect of Bank Stress on Interest Rate Setting

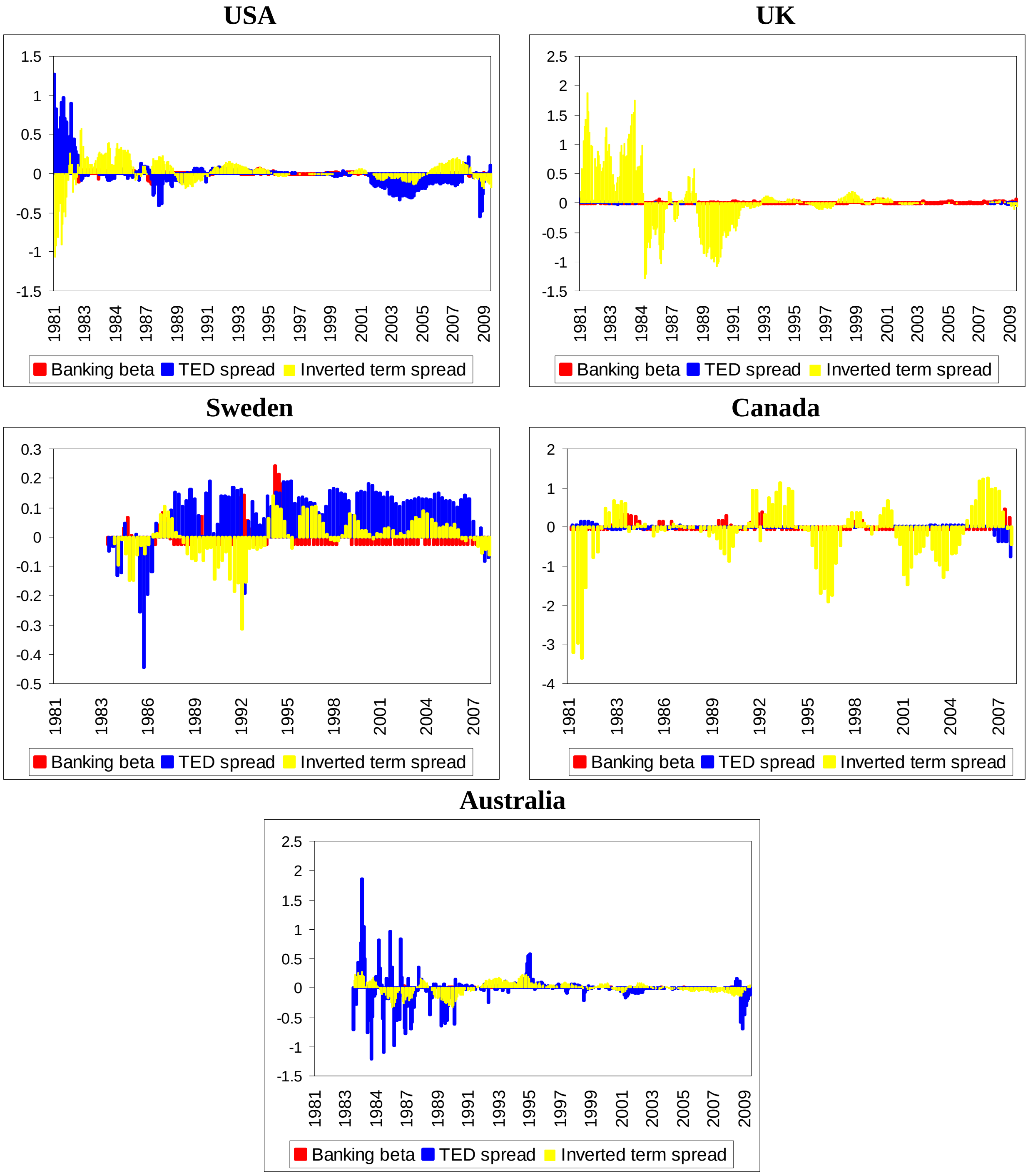
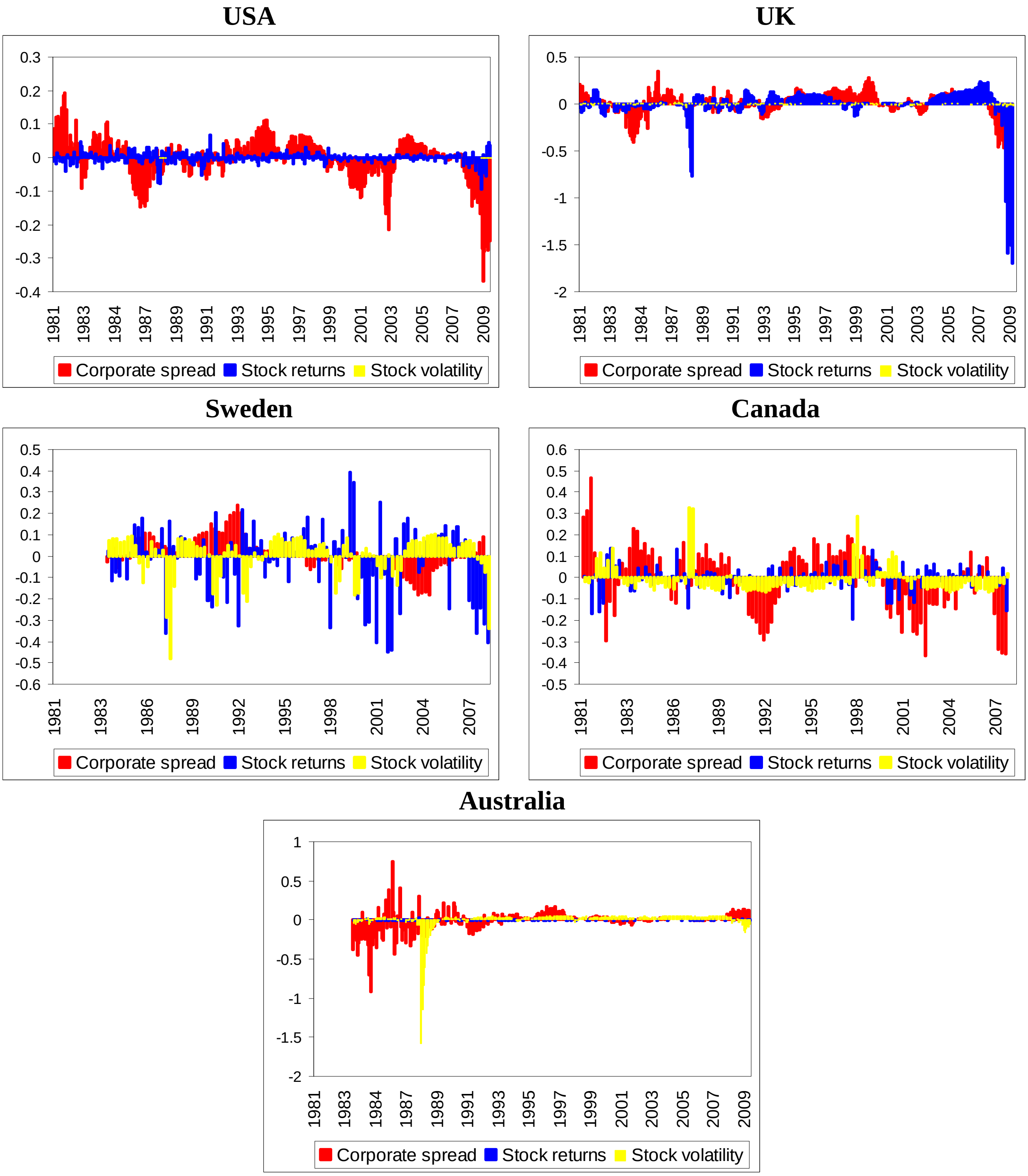


Figure A.2.4.2: The Effect of Stock Market Stress on Interest Rate Setting



Chapter 3.

Is Monetary Policy in New Members States Asymmetric?

3.1 Introduction

Since Taylor's (1993) influential paper there has been vast research about the way central banks handle interest rate setting. Clarida et al. (1998, 2000) propose that central bankers are proactive rather than reactive and set interest rates with respect to expected values of macroeconomic variables. Estimated monetary policy rules typically take a linear form assuming that monetary policy responds symmetrically to economic developments. A theoretical underpinning of a linear policy rule is the linear-quadratic (LQ) representation of macroeconomic models with an economic structure assumed to be linear and the policy objectives to be symmetric (the loss function is quadratic).

However, when the assumptions of the LQ framework are relaxed, the optimal monetary policy can be asymmetric, which can be represented by a nonlinear monetary policy rule.[1] The first source of policy asymmetry lies in the nonlinearities in the economic system. A common example of such nonlinearity is a steeper inflation-output trade-off when the output gap is positive. The convexity of the Phillips curve (PC) implies that the inflationary effects of excess demand are larger than the disinflationary effects of excess supply (e.g. Laxton et al., 1999). This can lead optimizing central bankers to behave asymmetrically (Dolado et al., 2005). However, asymmetric monetary policy can also be related to genuinely asymmetric preferences of central bankers. While central banks in the past were prone to inflation bias due to a preference for high employment or uncertainty about its natural level (Cukierman, 2000), reputation reasons can drive central banks, especially those pursuing inflation-targeting, to have an anti-inflation bias, which means that they respond more actively when inflation is high or exceeds its target value (Ruge-Murcia, 2004). In any case, it seems plausible that real monetary policy conduct is too complex to be described by a simple linear equation and nonlinear representation of the monetary policy can be more appropriate irrespective of its underlying sources.

[1] Asymmetric monetary policy implies that monetary policy rule or the Taylor rule, which is a schematization of policy reaction function, is nonlinear.

Numerous empirical studies have provided evidence that monetary policy setting of many central banks can really be characterized as asymmetric. The asymmetric loss function was found to affect the decisions of the Bank of England (Taylor and Davradakis, 2006) and the US Fed (Dolado, et al., 2004). Bec et al. (2002) confirm that the US Fed, Bundesbank and the Bank of France responded more actively to inflation during economic booms. Leu and Sheen (2006) and Karadelinkli and Lees (2007) detect an asymmetric response to the output gap by the Reserve Bank of Australia. Surico (2007a) claims that the European Central Bank ECB responded in its early years more strongly to output contraction than expansions and that the level of the interest rate itself was a source of policy asymmetry. Surico (2007b) establishes similar evidence on the FED's asymmetric response to the output gap in the pre-Vocker era and quantifies the inflation bias induced by such a policy. The asymmetries due to the convexity of the PC found in some European countries (Dolado et al., 2005) and the ECB (Surico, 2007a) were linked to wage rigidity in the European countries.

The monetary policy in the new member states (NMS) of the EU was subject to substantial changes along their economic transition. They were experimenting with diverse monetary policy and exchange rate regimes until the late 1990s when the policy regimes were settled in line with the bipolar view. Some countries adopted hard exchange rate pegs, which put a significant constraint on their monetary policy (the Baltic States, Bulgaria, Cyprus, Malta), while other economies decided to maintain an overall flexible exchange rate, allowing their central banks to pursue internal macroeconomic targets, in particular price stability (Central European Countries, Romania).

The empirical evidence on the monetary policy setting in the NMS is rather limited. A few studies (María-Dolores, 2005; Frömmel and Schobert, 2006; Mohanty and Klau, 2007; Vašíček, 2010a) provided some evidence on linear monetary policy rules. However, some narratives suggest that monetary policy in the NMS can be asymmetric. In particular, countries who adopted a regime of direct inflation targeting (DIT) can show asymmetric behavior due to reasons of reputation. Horváth (2008) finds some evidence of an asymmetric policy of the Czech National Bank after it adopted this regime. The reason was the need to gain credibility and to anchor inflation expectations. On the other hand, DIT is flexible enough to allow the policy makers not to contract demand when inflation is slightly above the target and the shocks are likely to be short-lived (Blinder, 1997). Similarly, it seems plausible that other concerns such as economic growth and financial stability can lead to the temporal dismissal of inflation targets.

In this paper, we test the hypothesis of asymmetric monetary policy in three Central European NMS, the Czech Republic, Hungary and Poland, who adopted the framework of DIT and maintain a flexible exchange rate. We employ two empirical frameworks to test the policy asymmetry: (i) a framework based on an underlying structural model, which allows discrimination between the sources of policy asymmetry but is conditioned by the specific model setting; and (ii) a flexible econometric framework, where monetary policy switches between two regimes, according to a threshold variable. Besides the common choice for the threshold variable, inflation deviation from the target and the stance of the business cycle, we use also a degree of financial stress in the economy to see whether inflation-targeting central banks behave differently when the economy is distressed.

The rest of the paper is organized as follows: the next section briefly reviews the main rationales for asymmetric monetary policy; in section 3, we present the empirical strategies that will be used to test the policy asymmetry and in section 4, our dataset. In section 5, we review the empirical results and the last section concludes.

3.2 Rationales for Asymmetric Monetary Policy

While linear monetary policy rules can be derived in the common LQ framework (Svensson, 1999; Clarida et al. 1999), the nonlinear policy arises when we allow for some departures from this setting. The structure of the economy is commonly described by two-equations, tracking the evolution of inflation and the output:

$$\pi_t = (1-\beta)\pi_{t-1} + \beta\theta E[\pi_{t+1}] + \lambda g\{y_t\} + \xi_t \tag{1}$$

$$y_t = (1-\mu)y_{t-1} + \mu E[y_{t+1}] - \varphi\{i_t - E[\pi_{t+1}]\} + \varsigma_t \tag{2}$$

where π_t is the inflation rate, y_t is the output gap, i_t is the nominal short-term interest rate and ξ_t and ς_t are supply and demand shock, respectively. Eq. (1) represents the AS schedule or the PC and Eq. (2) is the inter-temporal IS curve. While the traditional backward-looking model (Svensson, 1997) assumes $\beta = \mu = 0$, the New Keynesian model (Clarida et al., 1999) is forward-looking $\beta = \mu = 1$. The monetary authority is usually assumed to set the nominal interest rate so as to minimize the loss function:

$$L_t = f\left\{\left(\pi_{t+s} - \pi^*_{t+s}\right), y_t, x_t\right\} \tag{3}$$

where *f* represents general functional form, which can be quadratic if the preferences are symmetric, π^*_t is the inflation target and x_t are other policy objectives such as exchange rate stabilization or interest rate smoothing.

For a derivation of asymmetric monetary policy, which in practice is represented by a nonlinear reaction function, both functional forms *f* and *g* are important. While Dolado et al. (2005) assume a case where *g* is convex, Dolado et al. (2004) propose a more general setting with *g* that may not be linear and *f* may not be quadratic, though both papers use a backward-looking model ($\beta = \mu = 0$). Surico (2007a,b) employs a forward-looking setting ($\beta = \mu = 1$) with a linex form of the policy loss function, adding additional policy objectives x_t in Eq. (3), in particular that central banks wants to minimize the interest rate volatility around the implicit target as well as the deviation of the current interest rate from the past value. Therefore, different combination of functional forms (1)–(3) give rise to different versions of nonlinear policy rule that can be brought to the data. However, imposing a specific model structure can turn problematic given that many variables and their relations are not directly observable. In addition, the NMS are small open economies, where numerous external factors may affect the domestic inflation π_t and output y_t and the relations itself can be subject to structural change. Therefore, an alternative can be to use an empirical framework that tracks asymmetries in a monetary policy setting but does not rely on the specific structure of the model.

3.3 Empirical Testing of Asymmetric Monetary Policy

There are diverse empirical strategies to test monetary policy asymmetry. They typically consist of an estimation of a monetary policy rule that includes some nonlinear feature. We define, as a benchmark, a linear forward-looking monetary policy rule (Clarida et al., 1998, 2000), which can also be derived as optimal monetary policy in the New Keynesian model (Clarida et al., 1999):

$$i^*_t = \bar{i} + \beta\left(E\left[\left(\pi_{t+s} - \pi^*_{t+s}\right)\middle|\Omega_t\right] - \pi^*_{t+s}\right) + \gamma\left(E\left[y_{t+k}\middle|\Omega_t\right]\right) + \varepsilon_t \tag{4}$$

where all the variables have the previous meaning, i_t^* is the interest rate target, $\bar{i}$ is the nominal equilibrium interest rate, E is the expectation operator, Ω_t is the information available to the central bank at the time of policy decision and ε_t is the error term. Given that real-time data, underlying the policy decision (see Orphanides, 2001), is not available for the NMS, we need to use actual realizations of the variables as proxies of their expected values. In addition, we allow for interest rate smoothing. Therefore, the observed short-term interest rate is a combination of a rule-implied target i_t^* and the previous value of the interest rate i_{t-1}:

$$i_t = \rho i_{t-1} + (1-\rho)\Big(\alpha + \beta\left(\pi_{t+12} - \pi_{t+12}^*\right) + \gamma y_t\Big) + \upsilon_t \tag{5}$$

where all the variables have the previous meaning, α is the constant term, ρ is the smoothing coefficient, representing the strength of policy inertia, and υ_t is the new error term. The partial-adjustment behavior is typically justified by the fact that sudden changes in interest rate could have destabilizing effects on financial markets but its true intensity is still the subject of debate (Rudebusch, 2002, 2006). We set s = 12, which corresponds to common inflation targeting horizon, and k = 0, assuming that central banks respond to the current output gap. Given that the current value of the potential output is not observable, it must be also proxied by ex-post data, which makes it also potentially an endogenous regressor. The error term υ_t is a linear combination of forecast errors of the right-hand side variables and the original exogenous disturbance ε_t. Therefore, it shall be orthogonal to the present information set Ω_t. We will fit the Eq. (5) as a benchmark linear model using GMM with common Newey-West (1994) covariance estimator robust to heteroskedaticity and autocorrelation. The instruments are three lags of short-term interest rate, inflation rate, the output gap and interest rate in the euro area.[2]

[2] There is a certain controversy about additional variables that can affect the interest rate decisions. In particular, small open economies could adjust the interest rate e.g. to the exchange rate or international interest rates. However, the three NMS use the DIT, where the domestic price stability is the only official policy target. Moreover, there is no evidence that Hungarian and Polish central banks respond to any additional variable (Vašíček, 2010a). Although the interest rate of the euro area turns sometimes significant in the estimated policy rule of the Czech National Bank (Horváth, 2008, Vašíček, 2010a), it is puzzling whether this means a genuine aim to stabilize domestic interest rate vis-à-vis the euro area or it is only an effect of the euro area interest rate on the Czech inflation forecast, which the central bank responds to. That is why we include the euro area interest rate as instrument rather than regressor.

3.3.1 Nonlinearities in the economic system

Monetary policy asymmetry can be related to nonlinearities in the economic system. In particular, nominal price stickiness can cause a nonlinear trade-off between inflation and output. Dolado et al. (2005) derive nonlinear monetary policy rule when the PC is convex. They propose to augment the standard linear policy rule, such as Eq. (5), by an interaction term of the expected inflation and the output gap given that any inflationary pressures driven by the output gap are larger if the PC is convex, which calls for an additional interest rate increase whenever the output gap is positive.

To implement empirically this framework, we estimate in the first step a very simple backward-looking PC defined as:

$$\pi_t = \alpha + \beta\pi_{t-1} + \gamma y_{t-1} + \gamma\phi y_{t-1}^2 + u_t \tag{6}$$

where the present inflation rate π_t depends on its lagged value π_{t-1} and a lagged output gap y_{t-1}. The PC is nonlinear when the coefficient ϕ is significantly different from zero, in particular it is convex when ϕ >0 and concave when ϕ <0. Second, we estimate the corresponding nonlinear policy rule:

$$i_t = \rho i_{t-1} + (1-\rho)\left(\alpha + \beta\left(\pi_{t+12} - \pi_{t+12}^*\right) + \gamma y_t + \kappa\left(\pi_{t+12} - \pi_{t+12}^*\right) y_t\right) + \upsilon_t \tag{7}$$

where the positive and statistically significant value of the coefficient accompanying the interaction term of inflation and output κ is an evidence of rule asymmetry. In particular, the increase of the interest rate is more than proportional when the inflation is above the defined target or the output gap is positive.[3]

3.3.2 Asymmetric preferences of the central bank

Asymmetric preferences with respect to economic outcomes represent another rationale why the central banks can behave asymmetrically. They may disproportionally decrease the interest rate when the output is below its potential (to prevent further recession) or increase it when the inflation exceeds the specified target (for credibility reasons). Dolado et al. (2004) show that under asymmetric preference, the optimal policy rule is nonlinear, irrespective of the form of the AS schedule. In their model when the central bank assigns a higher weight to

[3] As we allow the inflation target to vary in time, we use an interaction term of inflation gap and the output gap rather than inflation rate and the output gap as Dolado et al. (2005).

positive inflation deviations from the target, the inflation volatility (conditional variance) becomes an additional argument in the monetary policy rule.

This claim can be empirically tested as follows. First, if the conditional inflation variance is time varying, the residuals of the PC (Eq. (6)) shall contain autoregressive conditional heteroscedasticity (ARCH) effects. The null hypothesis of conditional homoskedasticity can be tested by means of an ARCH LM test. If the null is rejected, Eq. (6) can be estimated more efficiently using an ARCH-type of model. We use the common GARCH (1,1) with the variance equation defined as:

$$\sigma_{\pi,t}^2 = \omega_\pi + \nu_1 \xi_{\pi,t-1}^2 + \nu_2 \sigma_{\pi,t-1}^2 \tag{8}$$

where the conditional inflation variance $\sigma_{\pi,t}^2$ (one-period ahead forecast variance) depends on the long-term variance (the constant term) ω_π, the ARCH term $\xi_{\pi,t-1}^2$ (the squared residuals from the last period), representing the impact of new information about volatility from the last period, and the GARCH term $\sigma_{\pi,t-1}^2$, representing the impact of forecast variance from the last period. We obtain the estimate of the conditional inflation variance $\sigma_{\pi,t}^2$, which is included as an additional regressor in an otherwise linear policy rule:

$$i_t = \rho i_{i-1} + (1-\rho)\left(\alpha + \beta\left(\pi_{t+12} - \pi_{t+12}^*\right) + \gamma y_t + \kappa\sigma_{\pi,t}^2\right) + \upsilon_t \tag{9}$$

If the coefficient κ is positive and significant, the monetary policy rule is nonlinear by virtue of an asymmetric loss function of the central bank.[4]

Surico (2007a, b) proposes a model with both asymmetric preferences and nonlinear PC, which leads to an exponential monetary policy rule. The way to bring such a nonlinear equation to the data is a linearization using a Taylor series approximation around points where the asymmetry-driving parameters are zero. This results in a policy rule:

$$\begin{aligned} i_t &= \rho i_{i-1} + (1-\rho)\alpha + \beta\left(\pi_{t+12} - \pi_{t+12}^*\right) + \gamma y_t \\ &+ \kappa_1\left(\pi_{t+12} - \pi_{t+12}^*\right)^2 + \kappa_2 y_t^2 + \kappa_3\left(\pi_{t+12} - \pi_{t+12}^*\right) y_t + \kappa_4\left(i_t - \alpha\right)^2 + \upsilon_t \end{aligned} \tag{10}$$

where the asymmetric preferences enter via squared terms of inflation and the output gap while the inflation-output interaction term controls, as in Dolado et al. (2005), for potential rule nonlinearity coming from nonlinearity in the PC. Moreover, the last term track potential

[4] To obtain consistent results of this estimation, it is necessary to assure that the previous ARCH model has not been misspecified and the estimated conditional variance is not noisy. The misspecification is tested by means of an LM test applied at standardized residuals from the GARCH model, which must not be serially correlated.

asymmetric preferences in terms of the deviation of actual interest rate i_t from the estimated equilibrium value α.

3.3.3 Policy regimes with a threshold effect

The previous frameworks derive the nonlinear monetary policy rule assuming specific functional forms and parameterizations of Eqs. (1), (2) and (3). Such a model-based approach allows linking of the estimated coefficients of the policy rule to the parameters describing the policy preferences and the structure of the economy. However, the results are greatly conditioned given that the underlying relations are not observable and can be more complex, especially in the case of small open economies. For instance, as far as the PC (Eq. (1)) is concerned, there is some evidence (Franta et al., 2008, Stavrev, 2009, Vašíček, 2010b) that inflation in the NMS holds both backward- and forward-looking components and is determined by diverse (external) factors above the output gap. At the same time, there is little empirical evidence about the shape of the AD schedule (Eq. (2)). High economic openness of these countries again suggests that domestic output can have external determinants. Finally, the loss function of monetary authorities (Eq. (3)) is not observable. Although all three countries apply officially the DIT aimed at price stability, other objectives are not discarded as long as they do not jeopardize the price stability.

Therefore, it may be preferable not to rely on a specific model and use statistical techniques that enable possible nonlinearities in monetary policy to be detected irrespective of their underlying sources. Kim et al. (2006) test the nonlinearities in FED policy rule using a flexible framework of Hamilton (2001) that takes into account the uncertainty about the function forms. Cukierman and Muscatelli (2008) employ smooth transition regression to test nonlinearities of the Taylor rules in the US and the UK. Florio (2006) augments their model with possibility of nonlinearities in the interest rate smoothing using the change in FED policy rate as a transition variable.

An alternative way is to model policy asymmetry by means of switches between regimes according to some threshold variable. This is an intuitive strategy considering the nature of monetary policy decisions. In particular, it seems more plausible that central banks modifying the policy stance in the face of information about (realized or expected) inflation than assuming that they consider the nature of the country's PC.

Using the benchmark forward-looking policy rule of Clarida et al. (1998, 2000), the simplest case occurs when the threshold variable and the threshold value are both known. In this case, the sample can be split and policy rule estimated in each regime (e.g. Bec et al., 2002):

$$i_t = \rho_1 i_{t-1} + (1-\rho_1)\left(\alpha_1 + \beta_1\left(\pi_{t+12} - \pi^*_{t+12}\right) + \gamma_1 y_t\right) + \upsilon_{1,t} \text{ if } q_i \geq Q \tag{11}$$

$$i_t = \rho_2 i_{t-1} + (1-\rho_2)\left(\alpha_2 + \beta_2\left(\pi_{t+12} - \pi^*_{t+12}\right) + \gamma_2 y_t\right) + \upsilon_{2,t} \text{ if } q_i < Q$$

where q_i is the threshold variable and Q is the threshold value. For example, we could assume different policy regimes depending on whether inflation is above and below the target or whether the output is above and below its potential (threshold value is assumed to be zero).

In reality, the threshold value may not be known. For example, the central bankers can turn very inflation-averse only when the inflation rate exceeds the target value very substantially. Taylor and Davradakis (2006) find such evidence for the UK using the current inflation rate as a threshold variable. Gredig (2007) estimates a threshold value of different variables (the inflation gap, the output gap and gross domestic product (GDP) growth) for the Central Bank of Chile (CBC) finding two different regimes according to the business cycle stance.[5] Moreover, the threshold variable may not be a direct argument in the monetary policy rule and no reasonable guess about the threshold value can be made. An intuitive example of such a variable is financial stress. While inflation is arguably the main concern of inflation-targeting central banks in normal times, it can be disregarded when the financial sector or local currency comes under significant pressure.

Threshold estimation (Hansen, 1996, 2000) uses statistic criteria to estimate consistently the threshold value (of continuous variable) that splits the sample into two regimes. Although his method requires that both regressors and the threshold variable are exogenous, Caner and Hansen (2004) suggested an extension for endogenous regressors.[6] We follow this framework given that we estimate a forward-looking policy rule from ex-post data. The model can be written as:

$$\begin{aligned} i_t = {} & \rho_1 i_{t-1} + (1-\rho_1)\left(\alpha_1 + \beta_1\left(\pi_{t+12} - \pi^*_{t+12}\right) + \gamma_1 y_t\right) f\left(q_i \geq Q\right) + \\ & \rho_2 i_{t-1} + (1-\rho_2)\left(\alpha_2 + \beta_2\left(\pi_{t+12} - \pi^*_{t+12}\right) + \gamma_2 y_t\right) f\left(q_i < Q\right) + \upsilon_t \end{aligned} \tag{12}$$

[5] Assenmacher-Wesche (2006) uses a Markov switching model for the US, the UK and Germany. She finds evidence in favor of low- and high-inflation regimes for all three countries.
[6] Taylor and Davradakis (2006) employ GMM estimation (of three-regime policy rule for the Bank of England) with a grid search of two threshold values (of inflation rate) that minimize the GMM criterion function.

where the function *f* indicates whether the threshold variable q_i takes the value above or below the threshold value *Q*. This method assumes sample splitting into two regimes and is suitable for random samples and weakly dependent time series.[7] The procedure is sequential. The first step consists of OLS estimation of endogenous variables (in our case inflation and output gaps) on a set of exogenous instruments:

$$\left(\pi_{t+12} - \pi^{*}_{t+12}\right) = \Pi_1 z_t + \zeta_{1,t} \tag{13}$$

$$y_t = \Pi_2 z_t + \zeta_{2,t}$$

where z_t are the instruments; in our case the lagged values of variables as in the regression such as in the linear case, Eq. (5). We obtain the predicted values of the endogenous regressors $\left(\hat{\pi}_{t+12} - \hat{\pi}^{*}_{t+12}\right)$ and $\hat{y}_t$ that are substituted in the original threshold regression (Eq. (12)):

$$\begin{aligned} i_t = \rho_1 i_{t-1} + (1-\rho_1)\left(\alpha_1 + \beta_1\left(\hat{\pi}_{t+12} - \hat{\pi}^{*}_{t+12}\right) + \gamma_1 \hat{y}_t\right) f\left(q_z \geq Q\right) + \\ \rho_2 i_{t-1} + (1-\rho_2)\left(\alpha_2 + \beta_2\left(\hat{\pi}_{t+12} - \hat{\pi}^{*}_{t+12}\right) + \gamma_2 \hat{y}_t\right) f\left(q_z < Q\right) + \upsilon_t \end{aligned} \tag{14}$$

Second, the threshold value Q is estimated in Eq. (14) sequentially according to criterion:

$$\hat{Q} = \underset{Q \in \mathbb{Q}}{\arg\min}\, S_n(Q) \tag{15}$$

where S_n is the squared residual of Eq. (14) and $\mathbb{Q}$ is the set of values of threshold variable q_z. S_n can be used to obtain inverted likelihood ratio (LR) statistics to test whether a particular value belongs to the threshold interval (Hansen, 2000):

$$LR_n(Q) = n\frac{S_n(Q) - S_n\left(\hat{Q}\right)}{S_n\left(\hat{Q}\right)} \tag{16}$$

At last, we estimate by GMM the monetary policy rule for sub-samples allowing for all the parameters switching between the two regimes. Unlike Caner and Hansen (2004), we use again the Newey and West (1994) heteroskedaticity and autocorrelation consistent (HAC) estimator given that the residuals of estimated Taylor rules are often serially correlated due to autocorrelated shocks or omitted variables. While a specific version of the Wald test can be employed to test the degree of dissimilarity of the coefficient in each regime and at the same time the nonlinearity of the monetary policy rule, we rely on a simple visual inspection of inverted likelihood ratio statistics (more details below).

[7] Caner and Hansen (2001) develop a threshold (autoregressive) model for variables with unit root but it has not been so far extended for the case for endogenous regressors.

We use three threshold variables: (i) inflation gap; (ii) the output gap; and (iii) the financial stress index (EM-FSI, more details below). While the FSI is a new variable not considered in our analysis yet, the use of inflation and output gaps is useful for testing whether their zero threshold value de facto assumed in nonlinear rules based on structural models (Dolado et al., 2004, 2005; Surico, 2007a,b) previous models is justified. Since the method requires the threshold variable to be exogenous, we always use the first lag of the respective variables as a threshold.

3.4 Data Description

Our dataset consists of monthly data ranging from 1998/1.M until 2010/3.M.[8] The principal data source is the Main Economic Indicator database of OECD and Eurostat.

The short-term interest rate is the three-month interbank interest rate for CZE and POL and overnight interbank interest rate for HUN given that the former is not available for the whole period of analysis.

The inflation rate is measured by year-on-year changes in the consumer price index (CPI). We assume a forecasting horizon of 12 months and use three measures of the inflation target (inflation gap is always a deviation of expected inflation from the target value): (i) the actual inflation target of each central bank;[9] (ii) the smoothed (HP) trend of inflation target;[10] and (iii) the smoothed (HP) trend of actual CPI inflation.[11] Figure A.3.1 compares the inflation gaps constructed by the three methods.

[8] We use monthly data to have a sufficient number of observations to apply the sample splitting techniques. Unfortunately, this comes at a cost. Some variables, such as inflation rates and interest rates, are highly persistent at monthly frequency. The persistency of the dependent variable in a model with partial adjustment drives the result that the coefficient of lagged dependent variable is very close to unity. This finding implies, in terms of the monetary policy rule estimates, an unfeasible conclusion that the response of interest rate to inflation rate is very limited in the short-term, while its long-term multiplier is very high.

[9] The construction of the inflation target series is not straightforward. First, the target definition varies across time (net inflation, headline inflation, CPI inflation). Moreover, it is often specified in terms of band, whose width changes over time as well. Therefore, we use always the official inflation target irrespective of its changing definition and when the target is defined by a band, we use its mean value. Second, the inflation targets are usually defined as the year-on-year inflation increase measured in the last month of each year. Therefore, we have assigned this value to all months of the respective year.

[10] The problem with the former method (see previous footnote) is that the inflation target changes abruptly between December and January. This is unfortunate because inflation expectations (forecast) of the central bank and economic agents do not follow this pattern. Therefore, it seems reasonable to smooth the series by HP filter to avoid such breaks.

[11] It can be argued that central banks aim rather at eliminating inflation that is significantly above its trend. This seems plausible for the NMS given that inflation targeting was introduced when inflation rates were still

The output gap is measured as the difference between the logarithm of the current value of the seasonally adjusted GDP (in millions of euros in 1995 prices) and the trend value obtained by Hodrick-Prescott (HP) filter (the smoothing parameter set to 14400). Given that the GDP is available only quarterly, we have disaggregated it to monthly frequency using a univariate statistical method of Fernandez (1981) that allows the information to be augmented with the related series. For this purpose we have used a monthly industrial production index, which is arguably the most related series to GDP available on monthly frequency.

The financial stress is measured by EM-FSI elaborated by the International Monetary Fund (IMF) (Balakrishnan et al., 2009). It is a composite index of five subcomponents: (i) 12-months rolling beta (from the capital asset pricing model – CAPM) of bank stock index; (ii) stock market returns (year-on-year change in stock market index multiplied by minus one, so that declines in stock prices implies index increase); (iii) stock market volatility (six-month rolling monthly squared stock returns); (iv) sovereign debt spread (10-year government bond yield minus 10-year US Treasury bill yield); and (v) exchange market pressure index (month-over-month percent changes in the exchange rate and total reserves minus gold). The EM-FSI is constructed as a simple sum of standardized subcomponents and is plotted for each country in Figure A.3.2.

3.5 Empirical Results

3.5.1 Linear monetary policy rules

The GMM estimates of the linear monetary policy rules (Eq. (5)) are presented in Table 3.1. As noted above, given a fundamental uncertainty of what is the best measure of inflation gap, we report for each country the results with inflation gaps derived from the three alternative measures of the inflation target: (i) the actual inflation target of each central bank; (ii) the smoothed (HP) trend of inflation target; and (iii) the smoothed (HP) trend of the CPI inflation. We can see that most of the coefficients have the expected sign. The expected inflation gap (coefficient β) enters significantly in the Czech Republic but not in Hungary and Poland (due to elevated standard errors). This finding is a bit puzzling but it can be an indication that the intensity of the interest rate response to the inflation gap is not linear. The significant response

relatively high. To anchor the inflation expectation, the central banks had to stick to targets that were lower to what monetary policy could immediately achieve. However, they indicated the intention of monetary authorities to stabilize the price level.

to the output gap (coefficient γ) found in Poland can be interpreted as a policy aimed at price stability as long as the output gap predicts future inflation pressures. At last, we can see that the degree of interest rate smoothing (coefficient ρ) is substantial. Nevertheless, a high smoothing coefficient can be interpreted in terms of true policy inertia only with a lot of caution (see Rudebusch, 2006) since the interest rates at monthly frequency are autocorrelated by construction.

Table 3.1: GMM estimates of the linear monetary policy rule (Eq. (5))

Country	α (const.)	β ($\pi_{t+12} - \pi^*_{t+12}$)	γ (y_t)	ρ (i_{t-1})	R^2	LB	J-stat.
CZE (infl. targ.)	3.23	1.24	0.53	0.93	0.99	0.00	0.85
	(0.52)***	(0.47)***	(0.41)	(0.01)***			
CZE (infl. targ. trend)	3.31	1.35	0.53	0.93	0.99	0.00	0.85
	(0.55)***	(0.51)***	(0.43)	(0.01)***			
CZE (infl. trend)	2.62	1.34	0.35	0.94	0.99	0.00	0.66
	(0.55)***	(0.60)**	(0.39)	(0.01)***			
HUN (infl. targ.)	4.37	2.32	3.14	0.97	0.93	0.00	0.68
	(6.24)***	(3.16)	(3.06)	(0.02)***			
HUN (infl. targ. trend)	-5.79	6.49	6.87	0.98	0.93	0.00	0.45
	(22.36)	(10.17)	(9.20)	(0.02)***			
HUN (infl. trend)	5.77	4.98	6.01	0.98	0.93	0.00	0.68
	(3.15)*	(4.36)	(4.19)	(0.01)***			
POL (infl. targ.)	5.05	2.43	3.59	0.96	0.99	0.00	0.72
	(1.02)***	(1.59)	(1.82)*	(0.01)***			
POL (infl. targ. trend)	5.41	1.35	2.41	0.95	0.99	0.00	0.80
	(0.69)***	(1.20)	(1.22)*	(0.01)***			
POL (infl. trend)	-6.84	25.54	21.28	0.99	0.99	0.00	0.38
	(11.42)*	(20.15)	(14.65)	(0.01)***			

Notes: HAC standard errors in parenthesis. *, **, *** denotes significance levels at 10, 5 and 1%. LB is p-value of Ljung-Box test for 1. order serial correlation. J-stat is p-value of Sargan overidentification test.

3.5.2 Nonlinear monetary policy rules due to nonlinearities in the economic system

The first potential driver of nonlinear monetary policy is a convex AS schedule implying that inflationary tendencies are stronger (due to capacity constrains) when the output gap is positive. Hence, as a first step we must test whether there is any evidence on the nonlinear relation between the inflation rate and the output gap.

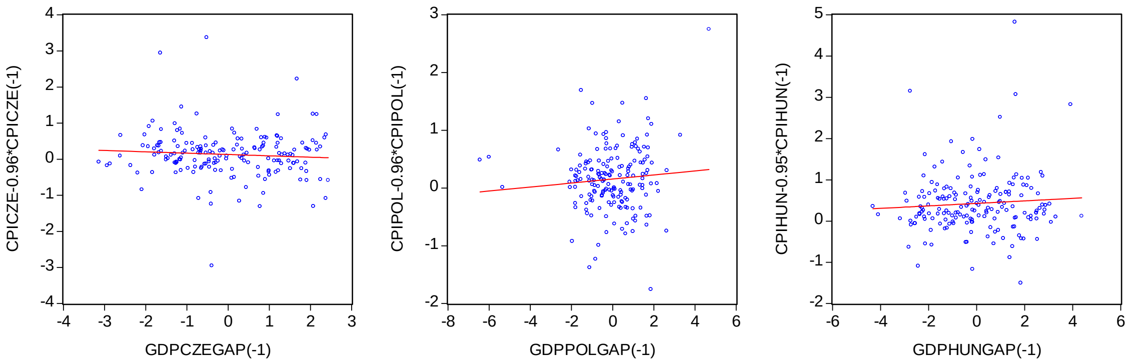
Estimates of the linear and nonlinear version of the simple backward-looking PC (Eq. (6)) appear in Table 3.2. Besides OLS we also use a GARCH(1,1) model to take into account the potential time-varying volatility of inflation. We are mainly interested in sign and statistical significance of the coefficient of the squared output gap $\gamma\varphi$. The PC is convex when this term is positive. The results show that there is little evidence of any (linear or nonlinear) relationship between inflation and the stance of business cycle in these three NMS. This is also evident by simple visual inspection of Figure 3.1, showing the scatter plots between the smoothed inflation rate $\left(\pi_t - \hat{\beta}\pi_{t-1}\right)$ and the output gap y_{t-1}. Although the results can be affected by noise in measuring the output gap, there is also some evidence showing that inflation rates in the NMS have significant external determinants (Stavrev, 2009; Vašíček, 2010b).

Table 3.2: OLS/GARCH estimates of simple linear/non-linear Phillips curves (Eq. (6))

Country	α (const.)	β (π_{t-1})	γ (y_{t-1})	$\gamma\varphi$ (y_{t-1}^2)	ω (const.)	ν_1 (ξ_{t-1}^2)	ν_2 (σ_{t-1}^2)	R^2	LB
CZE (OLS.)	0.10	0.96	0.06					0.95	0.00
	(0.07)	(0.01)***	(0.04)						
CZE (GARCH)	0.16	0.94	0.02		0.05	-0.04	0.87	0.95	0.00
	(0.11)	(0.04)***	(0.06)		(0.01)***	(0.02)**	(0.03)***		
CZE (OLS)	0.00	0.97	-0.07	0.04				0.95	0.00
	(0.11)	(0.02)***	(0.06)	(0.04)					
CZE (GARCH)	0.06	0.93	-0.02	0.07	0.32	0.95	0.05	0.95	0.00
	(0.13)	(0.04) ***	(0.07)	(0.04)	(0.11)***	(0.01)***	(0.03)		
HUN (OLS)	0.26	0.95	0.05					0.98	0.00
	(0.11)***	(0.01)***	(0.03)						
HUN (GARCH)	0.30	0.94	0.05		0.05	-0.02	0.83	0.97	0.00
	(0.12)***	(0.02)***	(0.03)		(0.06)	(0.02)	(0.23)***		
HUN (OLS)	0.32	0.95	0.05	-0.03				0.97	0.00
	(0.11)***	(0.01)***	(0.03)	(0.01)***					
HUN (GARCH)	0.36	0.95	0.05	-0.03	0.06	-0.03	0.80	0.98	0.00
	(0.12)***	(0.02) ***	(0.03)	(0.02)	(0.06)	(0.03)	(0.22)***		
POL (OLS)	0.10	0.96	0.06					0.98	0.00
	(0.07)	(0.01)***	(0.04)						
POL (GARCH)	0.19	0.93	0.09		0.00	-0.03	1.01	0.98	0.00
	(0.04)***	(0.00)***	(0.02)***		(0.00)	(0.00)***	(0.00)***		
POL (OLS)	0.07	0.96	0.06	0.03				0.98	0.00
	(0.08)	(0.01)***	(0.04)	(0.03)					
POL (GARCH)	0.14	0.93	0.09	0.04	0.06	-0.03	0.80	0.98	0.00
	(0.07)*	(0.02)***	(0.04)***	(0.03)	(0.06)	(0.03)	(0.22)***		

Notes: Standard errors in parenthesis. *, **, *** denotes significance levels at 10, 5 and 1%. LB is p-value of Ljung-Box test for . order serial correlation.

Figure 3.1: Scatter plots between smoothed inflation rate and output gap (the Phillips curve) and fitted linear trend



Although the previous results put into question the convexity of the AS schedule, we continue to estimate Eq. (7), where the inflation-output interaction term appears as an additional regressor. These results are reported in Table 3.3. As expected, this term is mostly insignificant and there is no indication of asymmetric central bank reaction driven by a nonlinear PC. In any case, it is important to keep in mind that the results are conditioned by the underlying model.[12]

Table 3.3: GMM estimates of the nonlinear monetary policy rule (Eq. (7))

Country	α (const.)	β $(\pi_{t+12} - \pi^*_{t+12})$	γ (y_t)	ρ (i_{t-1})	κ $((\pi_{t+12} - \pi^*_{t+12})y_t)$	R^2	LB	J-stat.
CZE (infl. targ.)	3.40	1.57	0.56	0.93	-0.48	0.99	0.00	0.72
	(0.49)***	(0.47)***	(0.32)*	(0.02)***	(0.38)			
CZE (infl. targ. trend)	3.70	1.80	0.57	0.95	-0.68	0.99	0.00	0.59
	(0.48)***	(0.45)***	(0.33)*	(0.01)***	(0.36)*			
CZE (infl. trend)	1.73	2.39	1.47	0.94	-1.06	0.99	0.00	0.43
	(0.83)**	(0.86)***	(0.59)**	(0.01)***	(0.59)*			
HUN (infl. targ.)	-0.86	3.64	8.41	0.98	-3.55	0.93	0.00	0.82
	(19.00)	(7.90)	(12.87)	(0.03)***	(5.92)			
HUN (infl. targ. trend)	1.82	2.64	5.61	0.96	-3.55	0.91	0.07	0.83
	(10.74)	(4.43)	(6.31)	(0.03)***	(5.92)			
HUN (infl. trend)	6.94	2.07	3.69	0.95	-1.71	0.93	0.02	0.71
	(1.76)***	(1.95)	(1.89)*	(0.02)***	(1.40)			
POL (infl. targ.)	5.03	2.46	3.61	0.96	0.03	0.99	0.00	0.63
	(1.36)***	(1.86)	(1.88)*	(0.02)***	(1.68)			
POL (infl. targ. trend)	7.24	-1.33	2.48	1.03	-3.65	0.98	0.00	0.26
	(1.44)***	(1.24)	(1.37)*	(0.03)***	(2.29)			
POL (infl. trend)	8.37	-8.08	0.30	1.07	-6.64	0.97	0.00	0.97
	(1.43)***	(4.16)**	(2.88)	(0.05)***	(3.90)			

Notes: HAC standard errors in parenthesis. *, **, *** denotes significance levels at 10, 5 and 1%. LB is p-value of Ljung-Box test for 1. order serial correlation. J-stat is p-value of Sargan overidentification test.

[12] Moreover, this framework implicitly assumes that the threshold value of inflation and output gaps driving policy asymmetry are each zero because the interaction term turns positive when inflation gap and the output gap are both positive or negative.

3.5.3 Nonlinear monetary policy rules due to asymmetric preferences

Central banks can respond in a nonlinear way to macroeconomic variables due to their genuine asymmetric preferences. These are usually represented by a non-quadratic loss function.

First, we explore whether the central banks of the three NMS applied nonlinear policy rule due to higher weight assigned to positive deviation of expected inflation from the target. Dolado et al. (2004) suggested tracking such nonlinearity by the inclusion of conditional inflation variance (Eq. (9)) to an otherwise linear policy rule. Therefore, first, we need to check whether the inflation volatility is truly time-varying to be used as an regressor in Eq. (9) . The inflation is again modeled by simple backward-looking PC (Eq. (6)) and the ARCH LM test is used to check the neglected ARCH in residuals. The test gives affirmative evidence for the Czech Republic and Poland but cannot reject the null of no conditional heteroskedasticity for Hungary. Conditioned on these results, we re-estimate the PC using GARCH (1,1). The results of the corresponding mean and variance equation both for linear and quadratic speciation of the PC appear in Table 3.2. We can see that the conditional variance of inflation is a rather persistent process in the three countries as the coefficient of the GARCH term ν_2 is significant and close to unity. We obtain the estimated series of conditional inflation variance and use it as a regressor (Eq. (9). The results appear in Table 3.4.

The short-term interest rate responds significantly to the conditional inflation variance in the Czech Republic, which suggests that the Czech National Bank handles the inflation in an asymmetric manner, in particular that it weights more positive deviations from the target than negative ones. On the contrary, the conditional inflation variance enters with a counter-intuitive negative sign for Hungary, which is likely related to noisiness (the residuals of PC for Hungary does not contain ARCH effects) and very low variance of this series (standard deviation is 0.02 as compared 0.43 for the Czech Republic and 0.16 in Poland). In any case, the results are again conditioned by the specification of PC that was used to derive the conditional inflation variance.

Table 3.4: GMM estimates of the nonlinear monetary policy rule (Eq. (9))

Country	α (const.)	β $(\pi_{t+12} - \pi^*_{t+12})$	γ (y_t)	ρ (i_{t-1})	κ (σ_t^2)	R^2	LB	J-stat.
CZE (infl. targ.)	2.63	1.35	1.19	0.92	4.72	0.99	0.00	0.59
	(0.80)***	(0.38)***	(0.34)***	(0.01)***	(1.97)**			
CZE (infl. targ. trend)	2.74	1.70	1.58	0.91	5.93	0.99	0.00	0.93
	(0.77)***	(0.48)***	(0.41)*	(0.01)***	(1.57)***			
CZE (infl. trend)	0.13	-0.53	3.85	0.99	11.31	0.99	0.00	0.77
	(6.89)	(2.62)	(6.55)**	(0.02)***	(18.82)			
HUN (infl. targ.)	27.54	0.75	0.72	0.94	-71.01	0.93	0.00	0.54
	(8.39)***	(1.06)	(0.85)	(0.02)***	(33.66)**			
HUN (infl. targ. trend)	40.07	1.88	1.56	0.95	-125.81	0.93	0.00	0.37
	(14.15)	(1.86)	(0.29)	(0.02)***	(60.94)**			
HUN (infl. trend)	6.94	2.07	3.69	0.94	-1.71	0.93	0.02	0.33
	(1.76)***	(1.95)	(1.89)*	(0.01)***	(1.40)			
POL (infl. targ.)	4.10	1.53	1.76	0.95	8.44	0.99	0.00	0.80
	(1.54)***	(0.47)	(0.84)**	(0.02)***	(15.41)			
POL (infl. targ. trend)	5.05	1.17	1.72	0.95	2.30	0.99	0.00	0.84
	(1.67)***	(1.35)	(0.80)**	(0.02)***	(16.53)			
POL (infl. trend)	4.26	1.08	2.68	0.94	9.16	0.99	0.01	0.68
	(1.26)***	(1.57)	(0.85)***	(0.02)***	(11.10)			

Notes: HAC standard errors in parenthesis. *, **, *** denotes significance levels at 10, 5 and 1%. LB is p-value of Ljung-Box test for 1. order serial correlation. J-stat is p-value of Sargan overidentification test.

An alternative way to test whether monetary policy rule is nonlinear due to asymmetric preferences is suggested by Surico (2007a,b). His approach does not require estimation of conditional inflation variance to test asymmetric response to inflation. In addition, it allows the testing of whether the central bank has asymmetric preferences with respect to the output gap and the interest rate gap, the latter is defined as a deviation of the current interest rate from its long-term equilibrium value. The asymmetric preferences enter the policy rule by square components for inflation, output and interest rate gaps (Eq. (10)). We adjust the nonlinear rule derived in Surico (2007a,b) to be more plausible for the inflation-targeting NMS. In particular, we replace the response to contemporaneous inflation by response to expected inflation gap given that inflation-targeting central banks are forward-looking and the inflation target is not constant. The estimates of such nonlinear policy rule appear in Table 3.5. The columns with estimates of κ_1, κ_2 and κ_4 refer to nonlinearities related to asymmetric preferences for inflation, output and interest rate gaps, respectively, and κ_3 captures a response to nonlinearities in economic structure. First, the only country where we find some evidence of asymmetric response to the inflation gap is Hungary, though the sign of coefficient κ_1 is negative, implying a stronger response when inflation is below its target. This contra-intuitive finding is in fact consistent with the evidence from Eq. (9) where we find a negative response to conditional inflation variance. On the other hand, we do not confirm the previous finding that the Czech National Bank treated the positive inflation deviation asymmetrically from its target. Second, the coefficient κ_2 of the squared output gap is insignificant for the three countries and if their central banks considered the stance of business cycle (see Tables 3.2–3.4), they did it in a symmetric manner. Third, for all countries we reveal a preference to limit the volatility of the current interest rate from its equilibrium value (proxied by the intercept α). The positive value of κ_4, found in the Czech Republic and Hungary, reflects a distaste for actual interest rates exceeding the equilibrium value. The negative value found for Poland can be a sign that the Polish National Bank was resistant to keeping the interest rates too low. In fact, the preference for higher interest rates (κ_4 negative) can also be an indication of a preference for price stability, while the opposite (κ_4 positive) can also indicate a preference to avoid contraction. As compared to the benchmark linear case (Eq. (5)), the interest rate smoothing has substantially decreased to more plausible levels (Rudebusch, 2002, 2006) as compared with the linear case (Eq. (5)). Finally, the inflation response coefficient β is not altered for the Czech Republic and Poland but it turns significant and higher than unity for Hungary, indicating a stabilizing nature of monetary policy conduct when the nonlinear nature of monetary policy is taken into account. These findings are promising as compared to Surico (2007a), who obtains for the ECB less plausible results such as a negative and insignificant response to inflation rate.[13] Due to reasons of space, we do not report the autocorrelation and over-identification tests but they provide a very similar picture as in previous tables.

[13] A bit surprisingly, he interprets these results as evidence that the ECB follows a nonlinear policy rule.

Table 3.5: GMM estimates of the nonlinear monetary policy rule (Eq.(10))

Country	α (const.)	β $(\pi_{t+12} - \pi^*_{t+12})$	γ (y_t)	ρ (i_{t-1})	κ_1 $(\pi_{t+12} - \pi^*_{t+12})^2$	κ_2 (y_t^2)	κ_3 $(\pi_{t+12} - \pi^*_{t+12})y_t$	κ_4 $(i_t - \alpha)^2$	R^2
CZE (infl. targ.)	2.26	1.05	0.03	0.85	0.03	0.63	-0.32	0.07	0.99
	(1.02)**	(0.47)**	(0.52)	(0.01)***	(0.18)	(0.53)	(0.32)	(0.02)***	
CZE (infl. targ. trend)	2.29	1.54	0.21	0.89	-0.12	1.17	-0.39	0.06	0.99
	(1.43)	(0.72)**	(0.91)	(0.05)***	(0.37)	(0.90)	(0.54)	(0.03)*	
CZE (infl. trend)	2.11	0.71	0.12	0.82	0.44	0.00	-0.63	0.04	0.99
	(0.88)*	(0.46)	(0.42)	(0.04)***	(0.29)	(0.49)	(0.31)**	(0.02)*	
HUN (infl. targ.)	6.59	2.09	-0.13	0.58	-0.55	-0.14	-0.15	0.14	0.88
	(0.43)***	(0.71)***	(0.37)	(0.16)***	(0.23)**	(0.16)	(0.33)	(0.02)***	
HUN (infl. targ. trend)	6.31	1.72	0.03	0.59	-0.43	-0.07	0.12	0.13	0.90
	(0.36)***	(0.64)***	(0.27)	(0.13)***	(0.21)**	(0.11)	(0.27)	(0.01)***	
HUN (infl. trend)	8.14	2.43	0.81	0.91	-1.49	0.52	-0.79	0.10	0.93
	(2.31)***	(2.96)	(1.24)	(0.06)***	(2.19)	(0.91)	(1.14)	(0.06)	
POL (infl. targ.)	19.84	0.58	-0.46	0.72	-0.37	1.38	0.80	-0.08	0.98
	(6.59)***	(0.73)	(1.18)	(0.41)***	(0.35)	(2.64)	(1.78)	(0.04)*	
POL (infl. targ. trend)	20.26	0.36	-0.49	0.51	-0.35	0.60	0.63	-0.07	0.97
	(5.97)***	(0.45)	(1.05)	(0.63)	(0.23)	(1.35)	(1.36)	(0.03)**	
POL (infl. trend)	4.57	0.28	0.17	0.83	0.78	0.23	2.14	0.06	0.99
	(2.16)**	(1.15)	(0.75)	(0.19)***	(0.74)	(1.40)	(2.34)	(0.03)*	

Notes: HAC standard errors in parenthesis. *, **, *** denotes significance levels at 10, 5 and 1%. LB is p-value of Ljung-Box test for 1. order serial correlation. J-stat is p-value of Sargan overidentification test.

3.3.4 Nonlinear monetary policy rules via threshold effects

As argued before, the previous methods of inference on policy asymmetry rests on the specific assumption about the structure of the economy and the central bank's loss function. In what follows, we use the empirical forward-looking policy rule proposed by Clarida et al. (1998, 2000) and allow the response coefficients to switch between two regimes according to the evolution of threshold variable. Given that the method of threshold estimation (Hansen, 2000; Caner and Hansen, 2004) requires the threshold to be exogenous, we use as a threshold observed (rather than expected) values.[14] Using inflation and the output gap as thresholds, we want to see whether the interest rate setting differs in high and low inflation regimes and in recession and expansions. Moreover, we include a new variable that can arguably give some insight on asymmetries in a monetary policy setting, the financial stress index (EM-FSI). In this case, we try to uncover whether central banks alter their consideration of common policy targets in the face of financial instability and whether they directly adjust policy rates according to the degree of financial stress in the economy.

[14] The econometric procedure is not suitable if the variables have a unit root. We apply common tests of unit roots, rejecting it (at conventional significance levels) for all the time series used for estimation.

The inference on monetary policy asymmetry has so far been carried out by means of conventional t-tests of statistical significance of additional nonlinear terms (inflation – output interaction term, conditional inflation variance or squared terms of inflation, output and interest rate gaps). With the current method, the policy asymmetry is tested by means of threshold effects. Unfortunately, a standard Wald test comparing the point estimates in each regime cannot be used because the method provides a sample split even in the absence of true threshold effects, which makes estimates inconsistent.[15] Given that the threshold estimation is based on the minimization of the squared residual of Eq. (14), we can draw the inverted LR statistics (Eq.(16)) for the entire set of possible threshold values $\mathbb{Q}$ to evaluate the precision of the estimated threshold (see Figures A.3.3 – A.3.5). $LR_n(Q)$ reaches its minimum, zero, at the estimated threshold $\hat{Q}$. The horizontal line represents the confidence interval and the values of Q whose $\mathrm{LR_n}(Q)$ are below this line are within the confidence interval. The shape of $\mathrm{LR_n}(Q)$ indicates the strength of the threshold effect. If the sequence of $\mathrm{LR_n}(Q)$ is peaked with a clearly defined minimum (of form V), it is also an indication of a significant threshold effect, which justifies sample splitting and separate estimation for each subsample. On the contrary, irregular shape where $\mathrm{LR_n}(Q)$ crosses the confidence interval more than once and the minimum is less evident, is an indication that the sample may be split more than once or that there is not threshold effect at all.

In Figures A.3.3 – A.3.5, we report the LR sequence using the inflation gap, the output gap and the EM-FSI as alternative threshold variables. As noted above, we always use the first lag of the respective variable as the threshold variable must be exogenous. For each threshold variable and country, we report three figures corresponding to a model with each measure of the inflation gap. As we can see in Figure A.3.3, the threshold effect of the inflation gap is not evident and depends on the measure of the inflation target. Although the LR sequences features a usually well-defined minimum, it leads to a very asymmetric sample split, leaving one regime only with a minimum number of observations permitted (when the inflation rate

[15] The method splits the sample at the value of threshold variable that minimizes the residuals of Eq. (14). When the splits imply that one regime contains only a minimum possible number of observations (10% of the total sample), while the other the remaining majority, it is an indication of no well-defined threshold. The Wald test comparing the slope estimates in each regime cannot be used as the slope coefficients in the smaller sub-sample are estimated very imprecisely. In addition, with no well-defined threshold, the estimation method encounters computation problems due to matrix singularity. As the threshold is not identified under the null hypothesis of no threshold effect, Hansen (1996) provides a bootstrapping procedure to test the presence of the threshold. However, given uncertainty about the threshold variable, the threshold value as well as the number of policy regimes, we assess the presence of threshold effect intuitively by the graphical inspection of LR statistics described below.

exceeds very substantially the target value for the Czech Republic and Poland and when it is significantly below it for Hungary). This disqualifies the reasonability of the sample splitting and asymmetric monetary policy along the value of the inflation gap. The only exceptions apply to the Czech Republic, when measuring the inflation gap by means of inflation deviation from its HP trend (right-most figure, estimated threshold is 1.25), and to Poland, when using inflation deviation from the target HP trend (middle figure, estimated threshold is 0.08). However, the estimated coefficients are mostly insignificant in both countries and regimes. To save space, we do not report the slope estimates.

Figure A.3.4 plots the respective LR sequences when the output gap is used as the threshold variable. We discard again the threshold model for Hungary as the LR reaches its minimum only at very high values of the inflation gap, making the sample split unfeasible. For the Czech Republic, we find a well-defined threshold only with the inflation deviation from the trend (the right-most panel). In this model, when the output gap exceeds the threshold value (estimated at 0.73), its coefficient γ is 2.43 versus 1.64 when it is below the target (in both cases this is highly significant). This suggests that the Czech National Bank handles monetary policy in an asymmetric way along the business cycle. In particular, it is ready to increase the interest rate by a larger amount during periods of economic expansion.[16] In the first two panels, we can see that the LR crosses the horizontal line more than once. However, the sample size does not allow another split. For Poland, we find a precise threshold in the first two models (with an inflation gap derived from the actual inflation target and from the HP trend of the target). The threshold value is estimated at -0.05 in both cases. While the corresponding response coefficient γ is insignificant in the regime below the threshold (i.e. when the output is below its potential), it turns significant and reaches a value of 14 when the threshold is breached. This finding is interesting in view of the linear model estimates (Table 3.1) showing that the National Bank of Poland (NBP) responds rather to the output gap than inflation. The results of the threshold model suggest that Polish monetary policy is asymmetric along the business cycle. However, this evidence cannot be directly interpreted that the NBP, as a long-term inflation targeter, aims at business cycle stabilization instead of

[16] These findings must be interpreted with caution given that the linear monetary policy rules (Table 3.1) features a substantially smaller and statistically insignificant response to the output gap. Similarly, the estimates of nonlinear policy rules in line with Surico (2007 a,b) reported in Table 3.5, do not indicate statistical significance of the squared term for the output gap.

the inflation target. It might mean that the output gap affects the NBP's inflation forecast that is the driver of interest rate setting.[17]

At last, we use the financial stress indicator (EM-FSI). The evolution of this variable (normalized to have zero mean) is depicted in Figure A.3.2. It is notable that all three countries experienced a degree of financial stress during the recent global turmoil unseen in the previous decade but that the stress was also high as a consequence of the Russian crises in late 1998. On the other hand, unlike many developed countries, the NMS did not suffer an increase in financial stress on the eve of the new millennium following the NASDAQ crash (2000), the terrorist attack on the US (2001) or the US corporate scandals (2002). EM-FSI allows, unlike binary crisis variables (Leaven and Valencia, 2008), the intensity of financial stress to be measured and can be used for the threshold estimation. Nevertheless, it is not evident whether EM-FSI should enter directly in the estimated policy rule as regressor or "stay behind" as threshold variable driving the regime switches. In other words, it is puzzling whether the central bank may directly respond to some stress measure or only to modify its consideration of other objectives. Consequently, we estimate the threshold model with and without the financial stress as an additional regressor. Figure A.3.5 depicts the LR evolution when EM-FSI is included as a regressor, which is almost identical with EM-FSI dropped. We can see that the threshold value is clearly delimitated in all three figures for the Czech Republic and first two figures for Poland. For Hungary, the LR sequence reaches its minimum at very high values of the stress but there are still 28 observations in the upper regime. We split all the samples and pursue GMM estimation for each regime. These results are reported in Table 3.6.[18]

In all but one case, the upper regime has substantially less observations than the lower one. The coefficient of EM-FSI is mostly significant suggesting that central bankers adjust policy rates when they are faced by financial stress. Since the central bankers might respond to increasing financial stress by monetary easing, the expected sign of the coefficient is negative. Yet, EM-FSI also includes a sub-component representing the exchange rate pressures, in

[17] GARCH estimates of the Polish PC reported in Table 3.2 indicate that the output gap has a significant effect on the inflation rate.
[18] We report results with EM-FSI included as an regressor given that this specification has a higher fit and the accompanying coefficient of EM-FSI is mostly statistically significant.

particular domestic currency depreciation,[19] whose prevalence in the overall index can drive an interest rate increase in an attempt to support domestic currency. For the Czech Republic and Poland, we find that the coefficient κ, accompanying EM-FSI is mostly negative and significant when the financial stress exceeds the estimated threshold. This suggests that both central banks decrease policy rates when the economy suffers high financial stress. On the contrary, the response is mostly insignificant when the stress falls below the threshold value. Hungary seems to be the opposite case; the interest rate response to financial stress is significantly positive and does not differ substantially between the two regimes. This could be related to the forint depreciation pressures that were a significant driver of the Hungarian overall EM-FSI and Hungarian monetary policy faced them by means of interest rate increase.[20]

As far as the other coefficients are concerned, their size usually differs between the regimes with the exception of the smoothing parameter ρ. Its estimated size still suggests a substantial degree of "policy inertia" even when we account for possible policy asymmetry via the threshold effects.[21] On the other hand, the serial correlation is much less pronounced in the split samples than in the models (linear, nonlinear) based on all observations. The inflation coefficient β does not have any clear pattern. While two specifications suggest that the Czech National Bank is a stricter inflation targeter when financial stress is high, the other points to the contrary. For Poland, in two specifications, there is no response to inflation when the stress is high and a positive response when it is low. The third specification that suggests the opposite pattern is in fact dubious because it cuts off a few observations when financial stress is very low. For Hungary, we still cannot determine the pattern of its inflation targeting because the response to the inflation gap is mostly insignificant. The coefficient of the output gap γ suggests that the real economy raises concerns only when the inflation stress is low (the Czech Republic and Poland) if at all (Hungary).

[19] This subcomponent is not present in the financial stress index proposed by the IMF for advanced economies (Cardarelli, et al. 2009).

[20] Baxa et al. (2010) study the response of main central banks (the US, the UK, Australia, Canada and Sweden) to financial stress using a time-varying parameter model that does not impose two policy regimes but allows a unique response in each period. Their results also suggest that the central banks are ready to decrease policy rates when the financial stress is high. Nevertheless, the size of the response varies substantially across countries and time not excluding periods when financial stress implied an interest rate increase. Unfortunately, due to the limited length of time series available, we cannot apply such a framework for the NMS.

[21] Although we have rejected the presence of unit roots in the short interest rates, they are still very persistent at monthly frequency. This seems to be the main reason for elevated policy inertia found across this study.

Table 3.6: 2SLS estimates of the FSI threshold value and GMM estimates of the monetary policy rule in each regime (Eq. (12))

Country	α (const.)	β (π_{t+12} - π^*_{t+12})	γ (y_{t-1})	ρ (i_{t-1})	κ (fsi_{t-1})	Q (threshold)	Observ.	R^2	LB
CZE (infl. targ.)	4.32	1.05	1.00	0.97	0.02	< 1.12	100	1.00	0.16
	(0.63)***	(0.55)*	(0.28)***	(0.01)***	(0.01)**				
	5.93	0.31	-0.89	0.90	-0.05	> 1.12	33	0.98	0.32
	(0.37)***	(0.13)**	(0.32)***	(0.01)***	(0.01)				
CZE (infl. targ. trend)	4.79	0.53	0.81	0.97	0.02	< 1.48	105	1.00	0.01
	(0.510***	(0.47)	(0.31)***	(0.01)***	(0.00)***				
	4.66	2.63	3.97	0.94	-0.01	> 1.48	27	0.98	0.78
	(0.68)***	(0.00)***	(0.97)***	(0.01)***	(0.01)				
CZE (infl. trend)	5.61	0.13	1.20	0.96	0.10	< 1.54	107	0.99	0.11
	(0.70)***	(0.35)	(0.35)***	(0.01)***	(0.03)***				
	9.73	2.92	-1.94	0.96	-0.10	> 1.54	28	0.98	0.25
	(1.98)***	(1.45)*	(1.43)	(0.02)***	(0.03)*				
HUN (infl. targ.)	10.51	-0.26	0.98	0.92	0.07	< 1.50	104	0.93	0.01
	(1.98)***	(0.79)	(0.87)	(0.02)***	(0.09)				
	10.80	-2.00	-6.46	0.95	0.13	> 1.50	28	0.88	0.30
	(3.05)***	(1.13)*	(2.77)**	(0.02)***	(0.01)***				
HUN (infl. targ. trend)	11.63	-0.54	0.77	0.92	0.17	< 1.50	104	0.94	0.02
	(0.91)***	(0.44)	(0.72)	(0.02)***	(0.03)***				
	6.23	-1.37	-4.65	0.95	0.15	> 1.50	28	0.90	0.49
	(2.72)***	(0.46)***	(1.42)***	(0.02)***	(0.01)***				
HUN (infl. trend)	11.06	-0.73	0.58	0.93	0.18	< 1.50	107	0.94	0.02
	(0.68)***	(0.58)	(0.74)	(0.02)***	(0.03)***				
	6.19	0.75	-1.21	0.90	0.15	> 1.50	28	0.91	0.70
	(0.46)***	(0.33)***	(0.47)***	(0.01)***	(0.01)***				
POL (infl. targ.)	5.39	1.95	0.69	0.95	0.01	< 0.14	96	0.99	0.01
	(0.01)***	(0.44)***	(0.39)*	(0.01)***	(0.02)				
	25.50	17.56	36.83	0.99	-0.20	> 0.14	36	0.98	0.24
	(26.34)	(32.27)	(61.39)	(0.02)***	(0.06)***				
POL (infl. targ. trend)	-2.05	2.06	0.63	0.96	-0.14	< 0.14	96	0.99	0.03
	(9.48)	(1.05)*	(1.01)	(0.01)***	(0.16)				
	32.69	13.10	61.53	0.99	-0.51	> 0.14	36	0.96	0.29
	(60.55)	(35.07)	(138.51)	(0.02)***	(0.09)***				
POL (infl. trend)	22.47	-1.97	3.02	0.97	0.13	< -2.44	26	0.99	0.89
	(3.88)***	(0.32)***	(0.39)***	(0.00)***	(0.03)***				
	4.36	1.10	1.27	0.95	-0.09	> 2.44	108	0.99	0.00
	(0.85)***	(1.73)***	(0.56)***	(0.1)***	(0.02)***				

Notes: HAC standard errors in parenthesis. *, **, *** denotes significance levels at 10, 5 and 1%. LB is p-value of Ljung-Box test for 1. order serial correlation.

There are, of course, several caveats of the threshold estimation. First, the method is purely statistical and can lead to sample split, which is contra-intuitive and slope estimates inconsistent with economic logic. Second, the present framework (Hansen, 2000; Caner and Hansen, 2004) allows only for two regimes. Therefore, the results are not reliable if there were more than two regimes or if the monetary policy was shaped by various threshold variables. For instance, under DIT, inflation is arguably the policy main concern but once the inflation target is reached; there can be other sub-regimes according to other variables such as the output gap, the exchange rate or the financial stress.

3.6 Conclusions

Numerous empirical studies try to describe the monetary policy decisions by means of estimated Taylor rules. There are different reasons why the monetary policy can be in fact asymmetric in the sense that the intensity of the central bank response varies according to economic developments. Our empirical analysis tries to reveal whether the monetary policy can be described as asymmetric in three NMS that apply a regime of inflation targeting (the Czech Republic, Hungary and Poland). We find that the overall evidence is mixed. When we use a GMM estimation of nonlinear policy rules derived from specific underlying models (Dolado et al., 2004, 2005; Surico, 2007a,b) we do not find rationales for asymmetric policy in terms of nonlinear economic relations. On the other hand, there is some indication of asymmetric preferences in inflation; in particular, the Czech National Bank seems to weight positive inflation deviations more severely from the target, while the opposite holds for Hungary. Interestingly, for all three countries we reveal their preferences to limit the volatility of the current interest rate from its equilibrium value. While for the Czech Republic and Hungary, we detect a distaste for actual interest rates exceeding the equilibrium value and for Poland we find that too low interest rates were of concern. In addition, the preference for lower rather than higher interest rates can also be an indication of a preference to avoid contraction, while the opposite points to a preference for price stability.

The previous results rely on specific nonlinear form because they are derived from specific parametric models. Although such an approach allows for discriminating between different sources of policy asymmetry, it can turn problematic when the underlying relations are not observable. Consequently, we use as an alternative a method of sample splitting where nonlinearities enter via a threshold variable and monetary policy is allowed to switch between two regimes (Hansen, 2000; Caner and Hansen, 2004). Besides the inflation and output gaps, we used a financial stress index as competing threshold variables. The threshold effects are most evident with financial stress index. While the Czech and Polish National Banks seem to face the financial stress by decreasing their policy rates, the opposite pattern is found Hungary.

There are different avenues of future research. First, it could be interesting to compare the behavior of central banks in the NMS and other emerging countries who use DIT but have faced very different economic challenges such as South Africa, Mexico or Chile. Second, the models that were used for the derivation of nonlinear policy rules (Dolado et al., 2004, 2005; Surico, 2007a,b) could be extended for small open economies to derive model-based nonlinear policy rules that are more suitable for the NMS. Third, with respect to the threshold model, the assumption of an exogenous threshold variable can be too restrictive given the forward-looking nature of inflation targeting. Recently, Kourtellos et al. (2009) extended the model of Caner and Hansen (2004) for an endogenous threshold variable. Finally, more complex econometric techniques such as Markov switching models (Assenmacher-Wesche, 2006) or state space models (Kim and Nelson, 2006) could be employed to take into account both the possibility that monetary policy is asymmetric but also that it evolves in time.

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Appendix

Figure A.3.1: Proxies of inflation gap (inflation deviation from 1. the inflation HP trend, 2. the official target, 3. the official target HP trend)

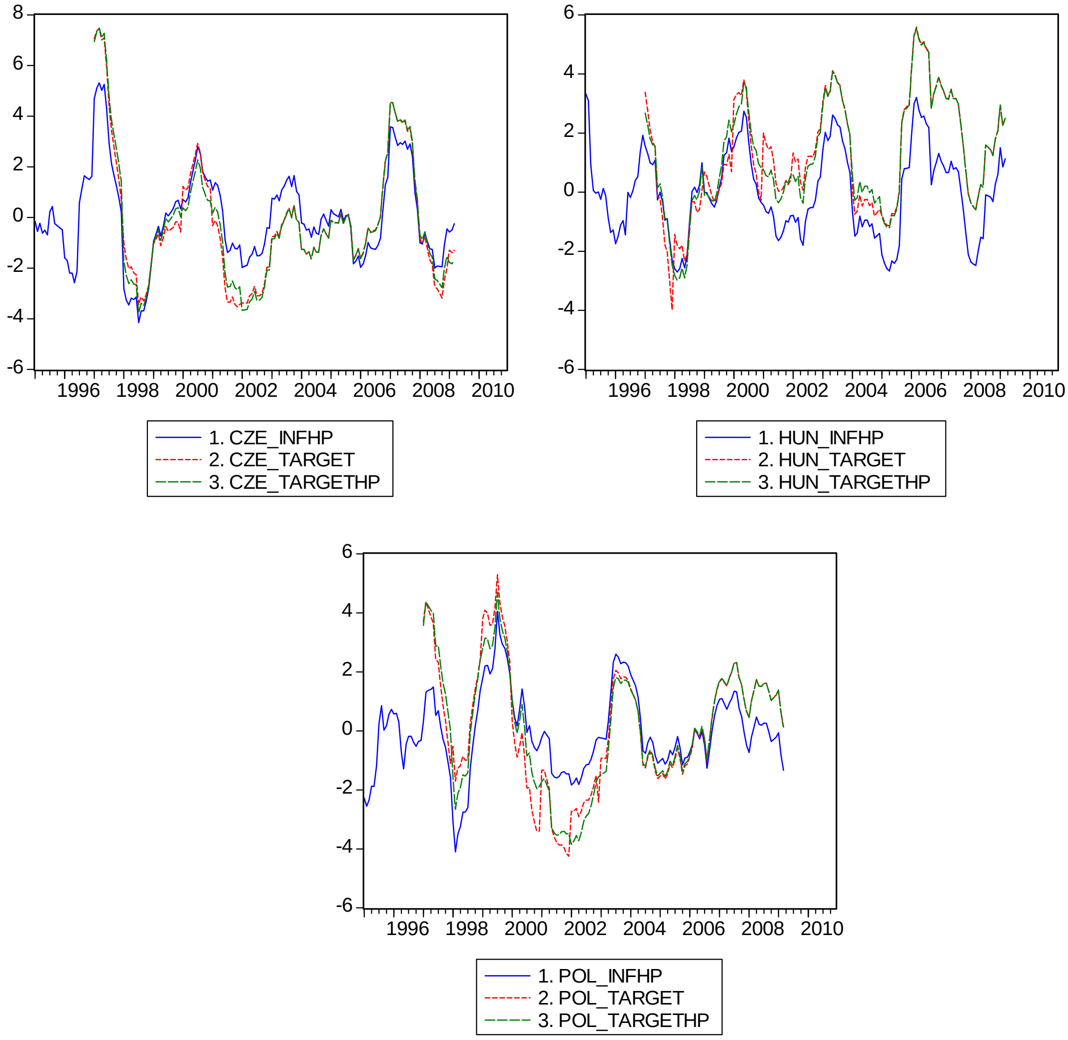


Figure A.3.2: IMF's Emerging-Markets Financial Stress Index (EM-FSI)

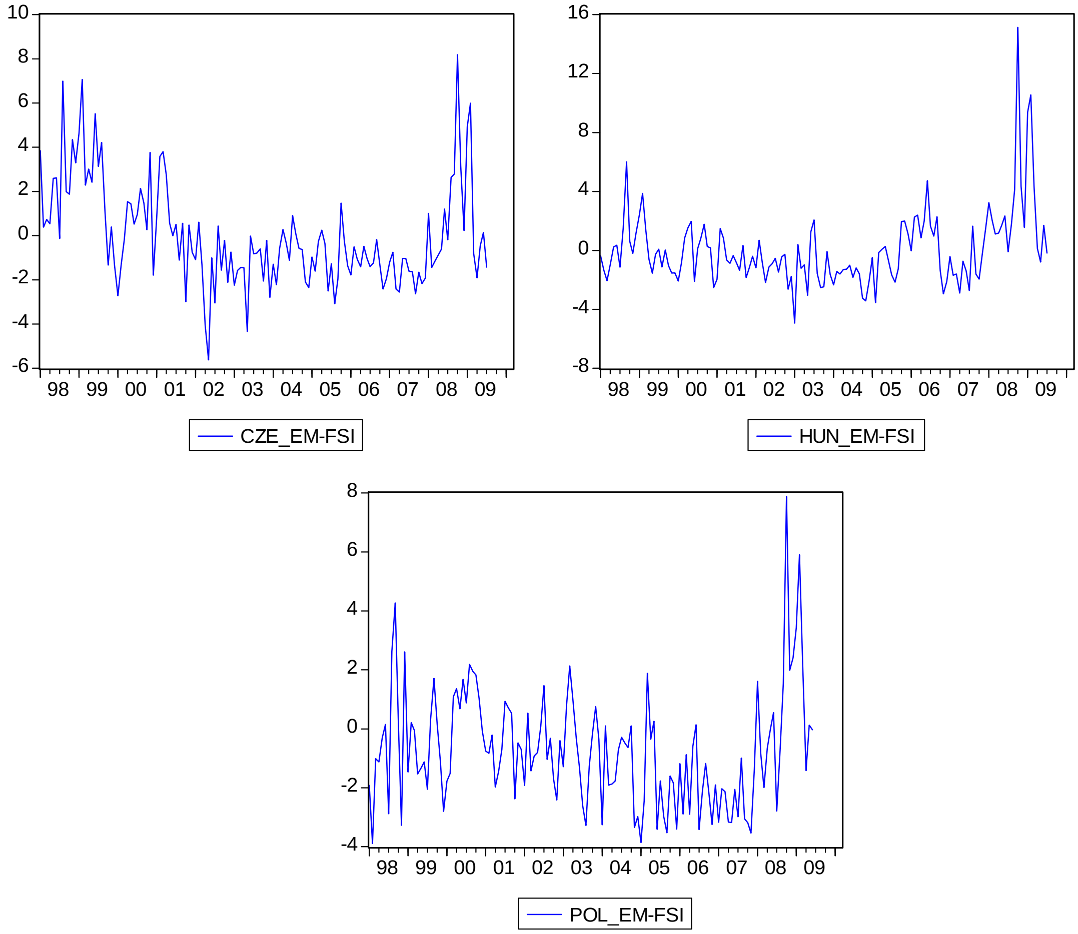
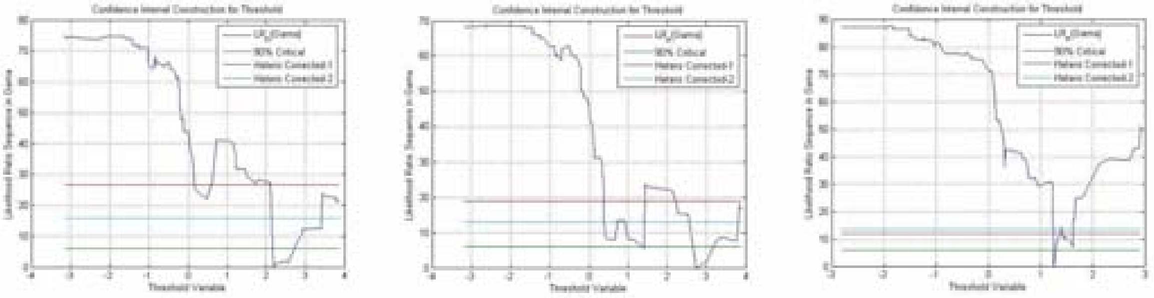
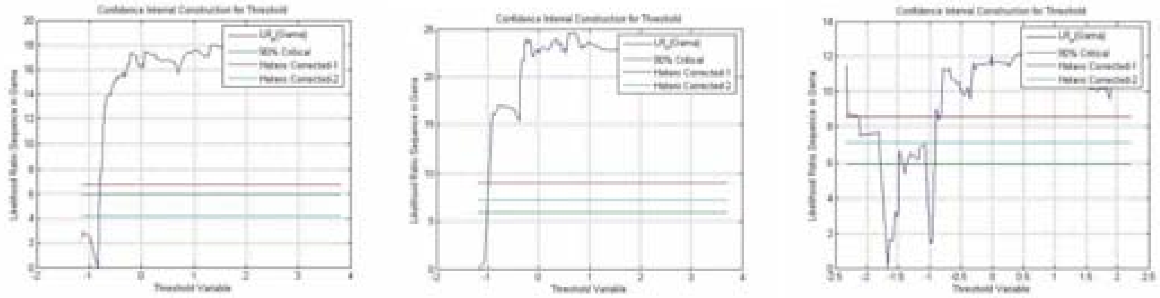


Figure A.3.3: Likelihood ratio sequence for different values of the threshold variable (the inflation gap)

The Czech Republic



Hungary



Poland

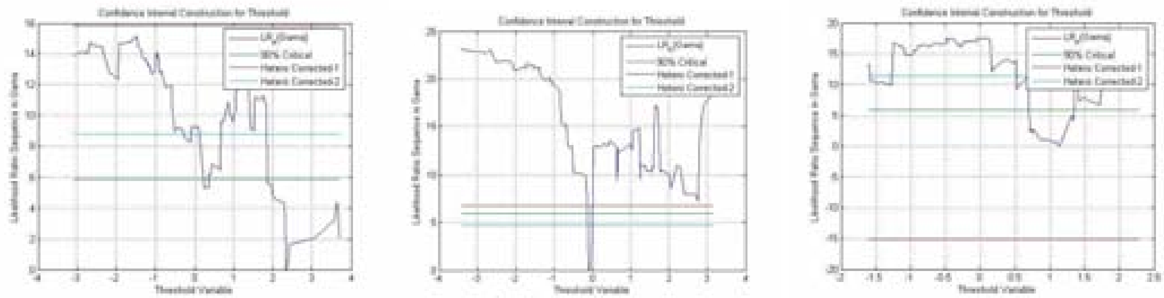
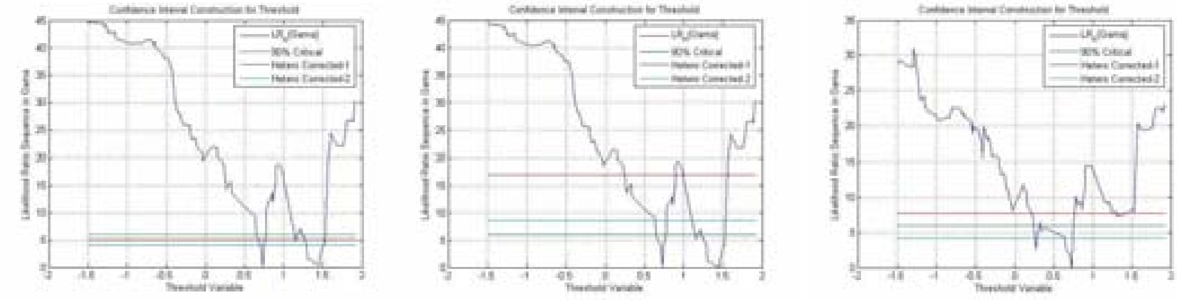
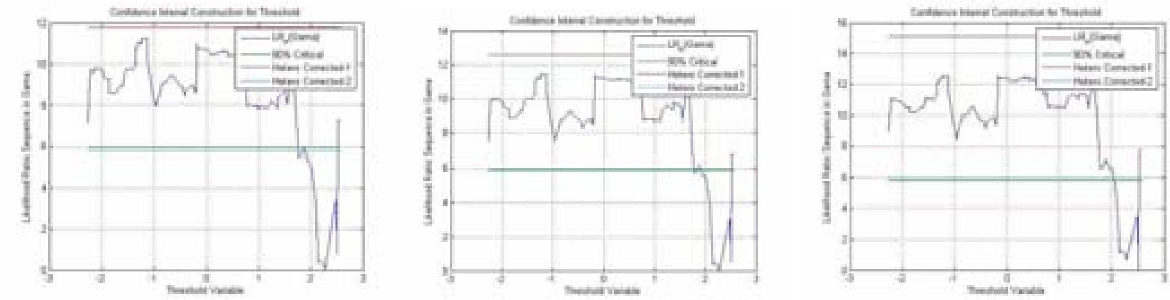


Figure A.3.4: Likelihood ratio sequence for different values of the threshold variable (the output gap)

The Czech Republic



Hungary



Poland

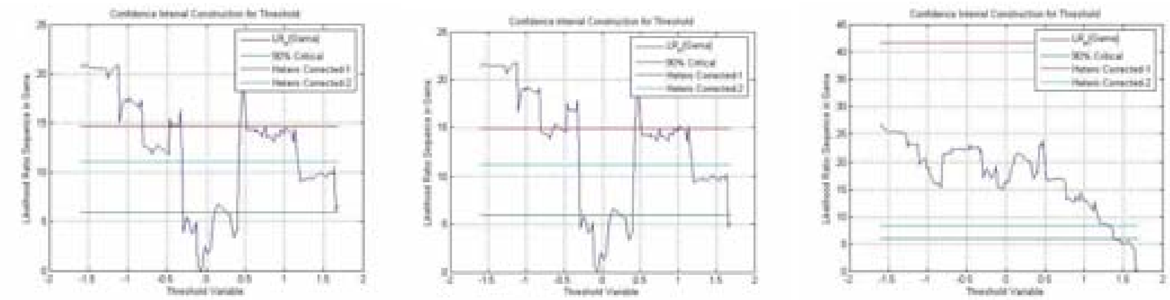
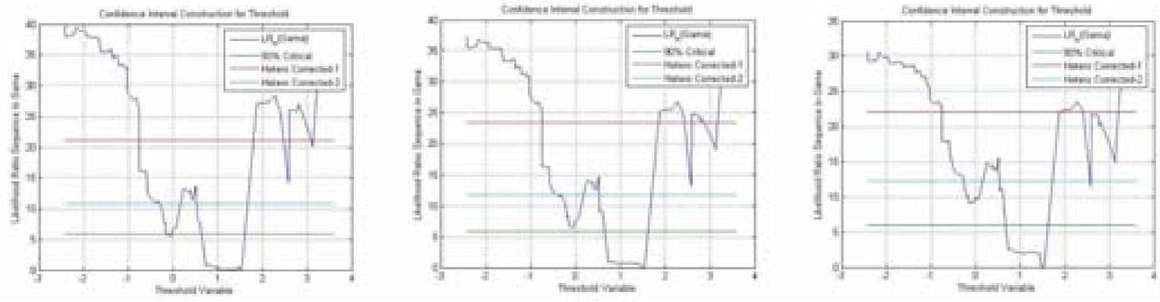
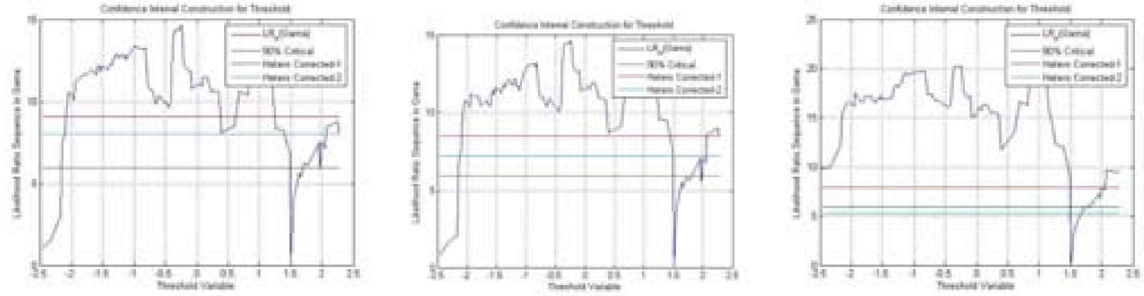


Figure A.3.5: Likelihood ratio sequence for different values of the threshold variable (the EM-FSI)

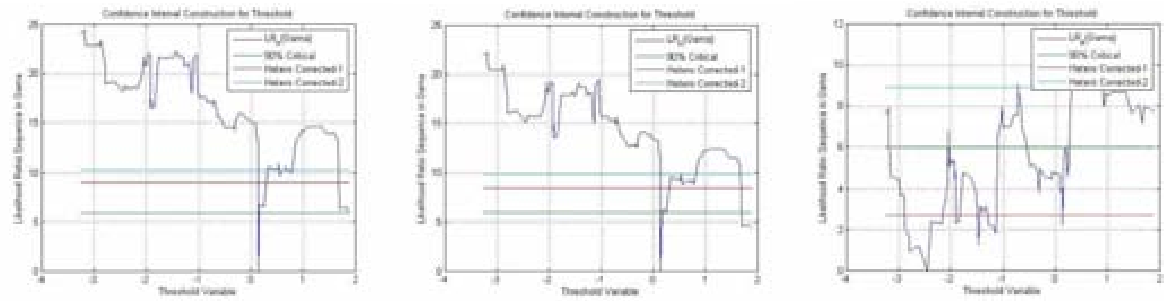
The Czech Republic



Hungary



Poland



Chapter 4.

Inflation Dynamics and the New Keynesian Phillips Curve in Four Central European Countries

4.1 Introduction

Understanding the nature of short-term inflation dynamics is very important for the implementation of monetary policy. The traditional (wage) Phillips curve (PC) suggested that there is a stable trade-off between (wage) inflation and economic activity, measured for example by the unemployment rate. At the same time, inflation was deemed very persistent. Although the traditional PC became subject to criticism in the 1970s, the persistence became a generally accepted feature of the inflation process. The New Keynesian Phillips curve (NKPC) that appeared during the 1990s possesses, unlike the empirical PC, elaborated microeconomic foundations. The NKPC links the current price inflation to expectations of the future inflation of economic agents. This NKPC arises in a framework of monopolistic competition and price rigidities. The proponents of the NKPC criticize the backward-looking nature of the traditional PC as being non-stable across policy regimes (the Lucas critique), inconsistent with rational expectations and overpredicting inflation in developed countries in the last decades. At the same time, traditional cyclical measures of real economic activity such as the output gap or the unemployment rate are disregarded as relevant determinants of inflation in favor of the aggregated marginal cost. Therefore, two issues related to inflation dynamics and the NKPC are subject to ongoing discussion: (i) whether inflation is a backward-looking or forward-looking process and (ii) what the main inflation-forcing variable is in the short term. The nature of inflation dynamics has important consequences for monetary policy. In particular, if inflation is a predominantly forward-looking phenomenon and its dependence on the past (intrinsic persistence) is limited, a credible monetary policy can achieve disinflation at no cost (in terms of real output loss).

The empirical evidence on the NKPC is vast, especially for major economies, but its results are ambiguous. The capacity of the NKPC to explain the inflation dynamics of small open economies is subject to controversy as well, and the (limited) availability and quality of data are the reason why studies on emerging market economies (EMEs) are very scant. The new

EU member states (NMSs) have numerous specific features not only in comparison with developed economies but also with respect to other EMEs. Therefore, it is of interest to analyze the inflation dynamics of these countries and evaluate whether the NKPC (currently the most influential microfounded model of inflation dynamics) can shed some light on this issue.[1] To this end, we test the baseline NKPC as well as the NKPC augmented by open-economy variables for the Central-European NMSs (the Czech Republic, Hungary, Poland and Slovakia, CECs hereafter). We address some concerns that were raised with respect to the generalized method of moment (GMM) estimation of a forward-looking model. Besides assuming rational expectation, as is common, we use additionally inflation survey data as a proxy for inflation expectations.

The inflation process in the four CECs is distinguished by some peculiarities, which can have an effect on the NKPC estimates. (i) The transition process in general and the price liberalization in particular were decisive determinants of the price inflation during most of the 1990s (Barlow, 2010). In fact, the inflation rates observed during the transition are likely to be subject to upward bias due to uncaptured quality improvements of the products on the market (Filer and Hanousek, 2003). Therefore, the period in which one could reasonably link the inflation developments to the price-setting behavior of firms (consistently with the NKPC), is relatively recent and short. (ii) The effect the transitional experience and systematically higher inflation rates had on this price-setting behavior is not obvious. One possibility is that higher inflation rates induced more frequent price reviews. It is also possible that local firms have not yet learned to use all the available information or face a higher cost for gathering it (Mankiw and Reis, 2002). (iii) The countries subject to our analysis, with the exception of Slovakia, consistently applied a regime of inflation targeting that anchored the inflation rates (Holub and Hurník, 2008) but that is also believed to drive down inflation persistence (Benati, 2008), though the impact of a prudent fiscal policy should not be disregarded either (Mikek,

[1] Besides the NK approach to analyzing the inflation dynamics either by estimating a single equation (the NKPC) or by applying Bayesian methods to Dynamic Stochastic General Equilibrium (DSGE) models, there are numerous alternative empirical approaches to studying the inflation process, some of which are also suitable for forecasting purposes. The statistical methods include simple autoregressive (AR) models or autoregressive fractionally integrated moving average (ARFIMA) models applied to study inflation persistence or generalized autoregressive conditional heteroscedasticity (GARCH) models to study inflation volatility. Vector autoregressive (VAR) models have been used to detect inflation determinants. Among the alternative theoretical models of inflation that have been brought to the data are for example the monetarist P-star model based on the equation of exchange or the mark-up model, where inflation is determined by changes in production costs and mark-ups. The Balassa–Samuelson effect, which refers to productivity differences in the tradable and non-tradable sectors, has been extensively studied for transition economies as it is deemed to be one of the main inflation drivers in these countries.

2008). (iv) All four CECs are small and very open economies.[2] Therefore, their domestic prices and inflation rates can be affected by external sources that were practically ignored in the original empirical studies on the NKPC but have been recognized recently (Batini et al., 2005; Rumler, 2007; Mihailov et al., 2009).

The principal results of our analysis are the following. (i) While the forward-looking NKPC is at odds with the data of the four countries, the hybrid specification has a very reasonable fit. (ii) Even though the price inflation in the four CECs holds a significant forward-looking component, it is also rather persistent. (iii) The evidence in favor of the marginal cost as the main inflation-forcing variable is fragile and the inflation dynamics seems to be driven by external factors.

The paper is structured as follows. The next section reviews both theoretic and empiric issues related to the NKPC. In section 3, we present our estimation framework, and section 4 discusses our data set and resumes the results of basic time series testing. Section 5 presents the estimation results of different versions of the NKPC. The last section concludes and points to possible extensions.

4.2 The Theory and Empirics of the NKPC

The NKPC is based on models of staggered price (or wage) setting by forward-looking monopolistic firms (Taylor, 1980; Calvo, 1983). Such firms set prices as a mark-up over their marginal costs subject to constraints that may temporally impede them from doing so. In particular, Calvo’s (1983) model assumes that in each period a firm faces a given probability θ that it may not be able to reset the price, which leads to optimal price p_t^f :

$$p_t^f = \mu + (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^k E_t\left\{mc_{t+k}^n\right\} \tag{1}$$

where μ is the (log) mark-up, β is the subjective discount factor, θ represents the frequency of price adjustments and mc_{t+k}^n is the nominal marginal cost. The aggregation across firms gives rise to the dynamic inflation equation where the current inflation rate depends on its expected value and the average real marginal cost:

[2] The shares of imports of GDP are the following: the Czech Republic – 72.7%, Hungary – 77.3%, Poland – 41.7% and Slovakia – 88.2% (source: European Commission).

$$\pi_t = \beta E_t\left[\pi_{t+1}\right] + \lambda_{NK} mc_t \quad (2)$$

where $\lambda_{NK} \equiv (1-\theta)(1-\beta\theta)/\theta$. Therefore, any increase in price rigidity θ makes inflation less sensitive to the real marginal cost mc_t. However, this forward-looking model did not allow the inflation persistence present in the data of most developed countries to be explained (inflation tends to be more persistent than marginal costs). Therefore, a substantial effort was made to give inflation persistence some structural basis.[3] Although there are different ways to hardwire the intrinsic inflation persistence, most empirical studies stem from a model that allows some firms to be backward-looking (Galí and Gertler, 1999, GG hereafter; and Galí, Gertler and López-Salido, 2001, GGL hereafter).[4] While forward-looking firms set prices optimally, i.e. with respect to the discounted value of the future marginal cost, the backward-looking firms follow a simple rule of thumb:

$$p_t^b = p_{t-1}^* + \pi_{t-1} \quad (3)$$

This means that they set prices in each period p_t^b with respect to the newly set prices in the previous period p_{t-1}^* and correct them for observed inflation. The index of newly set prices in period t can be written as $p_t^* = \omega p_t^b + (1-\omega)\, p_t^f$. Therefore, if the share of the forward-looking firms $(1-\omega)$ is large, they dominate the price index p_t and the price set by backward-looking firms is close to the forward-looking price. The inflation equation with both kinds of firms is the hybrid NKPC:

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f \beta E_t\left[\pi_{t+1}\right] + \lambda_H mc_t \quad (4)$$

where the parameters depend on the underlying structural parameters: $\lambda_H = (1-\omega)(1-\theta)(1-\beta\theta)/\left\{\theta + \omega[1-\theta](1-\omega)\right\}$, $\gamma_f = \beta\theta/\left\{\theta + \omega\left[1-\theta(1-\omega)\right]\right\}$ and $\gamma_b = \omega/\left\{\theta + \omega\left[1-\theta(1-\omega)\right]\right\}$. While older studies on the hybrid NKPC such as that of

[3] Several recent studies such as that of Benati (2008) argue that the inflation persistence found in the data is not structural as it varies significantly across monetary policy regimes. Cogley and Sbordone (2008) claim that inflation persistence arises due to variation in the long-run trend of inflation. Zhang et al. (2008) document that forward-looking behavior is substantially weaker when the inflation rate is high.

[4] There are alternative ways to obtain the hybrid NKPC. Futher and Moore (1995) propose a contracting model where the relative wages of successive cohorts of workers are linked. Christiano et al. (2005) introduce a price indexation to past inflation.

Fuhrer and Moore (1995) continued to use the output gap as the main driving variable, GG suggested using the real marginal cost, in particular the labor income share.[5]

Although the papers of GG and GGL became a benchmark for most empirical studies, there are numerous issues related to their estimation of the NKPC. (i) The first issue concerns the measure of expected inflation. Many empirical studies in line with GG assume rational expectation and proxy the expected values of inflation by the realized future inflation. The GMM estimator, which employs forecast errors to define the orthogonality conditions, is used to address the endogeneity. A few recent studies (Henzel and Wollmershäuser, 2008, Zhang et al., 2008, 2009) do not impose rational expectations and use the inflation survey data directly. Such a focus questions the dominance of the forward-looking term. (ii) The use of limited information methods such as GMM is subject to criticism as well. Bårdsen et al. (2004, 2005) demonstrate that the significance of the marginal cost in GG depends on specific choices in the GMM estimation and is not robust. Rudd and Whelan (2005, 2007) claim that the sensitivity to instruments can be an indication that some of them should be used directly as regressors, which applies especially to additional inflation lags. Mavroeidis (2004, 2005) points out that weak identification can induce a bias in favor of the forward-looking specification. Lindé (2005) advocates system estimation by full information maximum likelihood (FIML) as it provides more efficient parameters than limited information methods. Rudd and Whelan (2005) claim that GG's results are inconsistent as the reduced-form estimates are substantially different from the reduced-form parameters derived from the structural estimates, which is rebutted by Galí et al. (2005). (iii) Other issues are related to the suitability of the empirical model for revealing the parameters of interest. Del Negro and Schorfheide (2008) insist that the degree of backward-lookingness cannot be identified in the estimated NKPC if the mark-up shocks (the random disturbance to the NKPC) are serially correlated. Fuhrer (2006) claims that even when the estimated coefficient of the marginal cost is significant, it is typically too small to explain the persistence present in the inflation data. He argues that it is the intrinsic persistence (from the disturbances of the estimated NKPC) that explains most of the persistence of inflation. (iv) A few recent studies consider the effects of changes in the economic system and monetary policy on inflation dynamics. Hondroyiannis et al. (2007, 2008) shows (with the US and the EU data) that once the NKPC

[5] The advantage of the real marginal cost over the output gap is that it includes the impact of both productivity and wages on inflation. Moreover, statistical filters extract the potential output as a smooth trend whereas it may in reality be rather humpy as it is affected by shocks (e.g. technology shocks).

parameters are allowed to vary, the lagged inflation term becomes insignificant and inflation is a purely forward-looking phenomenon. Similar findings are reported by Cogley and Sbordone (2008), who claim that inflation persistence was detected in the data because of the omission of the drift in inflation trend. Benati (2008) and Zhang et al. (2008, 2009) support the claim that the parameters of the NKPC can vary across policy regimes and inflation persistence is in fact not structural. This paper takes the contribution of GG and GGL as a starting point but tries to reflect the criticism pointed out above.

Most empirical studies on the NKPC relate to the US or other major economies. However, since the NKPC was proposed as a general theory of inflation dynamics, it is of interest to test whether it is supported by the data of other countries. This poses an additional problem given that many economies are small and open. Therefore, firms can choose between domestic and foreign intermediate inputs, which affect the marginal cost. Besides, there are additional inflation channels that need not be related to price setting and expectations formations. For example, the import prices of intermediate and final products and the exchange rate volatility and its pass-through to domestic prices are among the factors that have an unquestionable effect on domestic inflation. Galí and Monacelli (2005) derive a version of the NKPC for CPI inflation of small open economies, which includes the terms of trade as an additional forcing variable. Mihailov et al. (2009) find some evidence that the terms of trade really affect the inflation of small open economies. Batini et al. (2005) propose an open economy NKPC where the marginal cost is affected by import prices and external competition. They present affirmative empirical evidence with the UK data. Rumler (2007) extends the marginal cost definition with the costs of intermediate inputs (both domestic and imported). He finds that such a model has a better fit for EU countries, claiming that the presence of imported inputs (with more volatile prices) presses the domestic firms to adjust their prices more frequently. Other empirical studies question the validity of the NKPC for open economies, e.g. Balakrishnan and Lopez-Salido (2002) for the UK, Sondergart (2003) for Germany, France and Spain, Dufour et al. (2006) for Canada and Genberg and Pauwels (2005) for Hong Kong.

Empirical studies aimed at transitional countries identify the Balassa–Samuelson effect as the main source of inflation developments during the transition (Égert, 2002, Backé et al., 2003). There are a few studies that aimed to test the NKPC for the NMSs (e.g. Masso and Staehr, 2005, Debusinskas and Kulikov, 2007 for the Baltic countries). Arlt et al. (2005) reject the validity of the NKPC for the Czech Republic using cointegration methods. Ledvai (2005) claims that the hybrid NKPC (augmented by imported goods) provides a reasonable account of the Hungarian inflation dynamics. However, convincing evidence of the significance of the marginal cost is not provided. Franta et al. (2007) conclude that inflation in three CECs (CZE, POL, SVK) is more persistent and the NKPC in terms of GG does not adjust to the data of any analyzed country. Hondroyiannis et al. (2008) use panel data of the seven NMSs, claiming that once the NKPC is estimated by means of a time-varying model, inflation becomes a purely forward-looking phenomenon. Lastly, Mihailov et al. (2010) test a small economy NKPC derived by Galí and Monacelli (2005) using data from twelve NMSs. Although they find ambiguous evidence in favor of this version of the NKPC, where external factors affect domestic inflation via changes in terms of trade, the fit of this model is better for the NMSs than for the developed OECD economies (Mihailov et al., 2009).

Previous studies of the NMSs use data starting in the mid-1990s, when the monetary policy acted in a very discretional way and diverse administrative measures such as price liberalizations were affecting the inflation rates. Moreover, as the NKPC is an equilibrium relation departing from a model log-linearized at a zero inflation steady state, it does not seem plausible to be estimated when the steady-state inflation rates were high and due to changes in the policy framework arguably even time-varying. These issues became subject to empirical research only recently. Cogley and Sbordone (2005) derive the NKPC with positive steady-state inflation. Zhang et al. (2008) document that the NKPC is not structurally stable across regimes with substantially different inflation levels. Benati (2008) and Cogley and Sbordone (2008) claim that estimation of the NKPC across different policy regimes and across periods with a varying inflation trend leads to overestimation of the backward-looking component.

Our contribution is to provide comprehensive empirical evidence on the inflation dynamics and the NKPC of the CECs using a recent data set from the post-transitional period in which the monetary policy regimes were settled and the inflation rate remained at a one-digit level.[6] Departing from the framework of GG and GGL, we apply additional checks to make the

[6] Although the inflation target, which can be considered a proxy of steady-state inflation, was steadily decreasing even during the last decade, it was well announced to the public and it is reasonable to assume that the inflation expectations were adjusted accordingly. Similarly, even though the inflation rates have a decreasing trend over the sample, for the sake of comparability with previous studies and consistency with the original model we use raw data rather than implementing ad-hoc measures such as inflation detrending. The data span clearly does not allow the use of a time-varying model. However, as the NKPC is a structural model, its parameters should be structurally stable at least with the same policy regime.

GMM estimation reliable. We test the NKPC using alternative measures of both inflation rate and different forcing variables, including external variables, arguably relevant for small open economies. In addition, we provide some evidence based on survey data.

4.3 Econometric Approach

We use the hybrid version of the NKPC as developed in GG and GGL as a benchmark. However, we extend the empirical framework to reveal some additional evidence on the nature of inflation dynamics in the NMSs. To this end, we employ alternative model specifications, variables, measures and auxiliary diagnostic tests. The empirical NKPC in closed form can be written:[7]

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E\left[\pi_{t+1}\right] + \lambda mc_t + \varepsilon_t \tag{5}$$

where all the variables keep their meaning and ε_t represents the residual. Note that the residual ε_t can be autocorrelated for different reasons: (i) inflation shocks (e.g. cost-push shocks) are autocorrelated, (ii) the prediction error η_{t+1} and the random disturbance ζ_t that form the error term $\varepsilon_t = \zeta_t - \gamma_t \eta_{t+1}$ are correlated and ε_t is autocorrelated (up to the first order) by construction (Mavroeidis, 2005), (iii) there are omitted variables (e.g. various external variables for small open economies). The marginal cost enters the empirical model in deviation from its steady state. We use the sample mean and the HP trend for its approximation. However, given the uncertainty about the true value of the steady state and the noise that can be introduced by demeaning or detrending, we use also the original series of the marginal cost.[8]

A known shortcoming of the GMM estimator is its sensitivity to instruments. If the instrument set includes inflation-driving variables that are not included as regressors, the

[7] There are several reasons why the estimation of the underlying structural parameters is not very reasonable in our case. (i) The structural model requires the estimation of several additional parameters, which is unfeasible with the available data sample. (ii) The estimates of structural parameters are sensitive to normalization of the orthogonality conditions (GG, GGL). (iii) Some of the parameters cannot be directly estimated and must be calibrated by plausible values that are not available for the NMS. (iv) The estimation of overall slope parameters in the reduced form is equivalent to their calculation from the estimated structural parameters (GGL, 2005). (v) When the empirical model is augmented by additional (e.g. open-economy) variables, the structural relations do not hold anymore and there is no simple procedure mapping the overall slope coefficients with structural parameters of the benchmark NKPC.

[8] Consistent with GG, we proxy the real marginal cost with the unit labor cost. How to estimate the steady-state level is rather puzzling since we have a rather short data span (10 years vs. 30 years in GG). Therefore, the deviation neither from the sample mean nor from the HP trend has to provide a reasonable approximation.

estimate of the forward-looking term is upwardly biased (Rudd and Whelan, 2005, 2007). Therefore, we test for an alternative domestic forcing variable that GG suggested only as instruments:

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E\left[\pi_{t+1}\right] + \lambda_{in} in_t + \varepsilon_t \tag{6}$$

where in_t stands for the output gap and the nominal wage inflation. The output gap can be understood either as a proxy for the marginal cost given that both variables are proportional under Cobb–Douglass technology, or it can be considered, in line with traditional PC, a proper inflation-driving force. The nominal wage inflation is a measure of wage pressures (again rather in line with traditional PC).

We have suggested that the inflation dynamics of small open economies can be affected by additional factors. These factors may or may not be related to the price setting of domestic firms and consequently may affect the marginal cost or not.[9] Therefore, rather than modifying the definition of the marginal cost (Rumler, 2007), we follow Galí and Monacelli (2005), who propose an NKPC for small open economies where inflation is measured by CPI, the marginal cost is defined in the standard way and the effects of external factors enter separately via a change in terms of trade.[10] Consequently, we test individually the domestic and external factors but instead of terms of trade, we include several external variables that have been found to be relevant inflation determinants in empirical studies.[11] [12] This allows us to compare the results of the benchmark model of GG and GGL and to evaluate the differential impacts of domestic and external factors. In particular, we test the effect of four variables that partially affect the terms of trade: oil prices, import prices, the foreign inflation rate and the exchange rate. Consequently, we augment the hybrid NKPC with an open economy as follows:

[9] This claim seems especially plausible for CPI inflation. CPI inflation is very relevant in small open economies, which import a substantial part of their consumption basket.

[10] Their model is based on some restrictive assumptions, in particular a complete exchange-rate pass-through (ERPT) and producers' currency pricing are assumed whereas most empirical studies suggest that the ERPT is incomplete (e.g. Gagnon and Ihrig, 2004), which is consistent with consumer currency pricing.

[11] We believe that this strategy has some empirical advantages over using the change in terms of trade (compare with Mihailov et al., 2009, 2010): (i) the terms of trade are an aggregate that only reflects the changes in underlying factors such as the exchange rate or foreign prices; (ii) without introducing the factors explicitly, one cannot evaluate their differential effect; (iii) while the proposed relation links the inflation to the expected change in the terms of trade relative to the observed change in the terms of trade (de-facto second difference of the terms of trade), it does not contemplate the possibility that inflation has feedback from the foreign sector entirely unrelated to agents' expectations, which can occur in economies with a very high degree of openness such as the CECs; (iv) the terms of trade can be endogenously affected by the domestic inflation level.

[12] The inflation equations used in empirical studies, e.g. studying ERPT (Gagnon and Ihrig, 2004), routinely include external variables but entirely ignore the role of inflation expectation. On the other hand, the studies that estimate the NKPC usually do not contemplate the possibility that external variables can affect domestic inflation via channels unrelated to domestic price setting.

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E[\pi_{t+1}] + \lambda_{in} in_t + \lambda_{ex} ex_t + \varepsilon_t \tag{7}$$

where in_t stands for the domestic forcing variables and ex_t stands for the external forcing variables.[13] As the external variables can be correlated (e.g. oil prices and import prices when oil represents a substantial part of the import basket), we include them one by one. On the other hand, they will not be correlated with the domestic variables.[14] All the estimations are performed by means of the GMM.[15] In closed economy specification we use similar instruments to GGL (2005): two lags of the inflation rate, the marginal cost, the output gap, the nominal wage inflation and the (log) unemployment rate. The external variables are treated as potentially endogenous like the domestic ones and two lags of each are added as instruments.

The use of the GMM estimator controls for potential reversed causality from the inflation rate to explicative variables (e.g. the exchange rate). However, Mavroeidis (2004, 2005) shows that the application of the GMM framework to the model with rational expectations has many limitations, the most serious being the problem of identification. In particular, if the structural model (the NKPC) has a forward-looking solution, a model for forcing variables must be specified. It turns out that that such a model, which links the forcing variable with its lagged values and the lagged values of inflation, is suitable for testing the identification in the structural model. Accordingly, we estimate an auxiliary regression for alternative domestic inflation-forcing variables (the marginal cost, the output gap and the wage inflation):

$$in_t = \sum_{i=1}^{p} \rho_i in_{t-i} + \sum_{j=1}^{q} \varphi_j \pi_{t-j} \tag{8}$$

where p=q=4 and the under-identification can not be rejected when $\rho_i = \varphi_j$ =0 for all i>1 and all j>2 (Mavroeidis, 2005, Theorem 3.1).

[13] Note that this equation is the NKPC augmented with open-economy variables rather than an open economy NKPC. While the latter would have to be derived from a structural model, the former empirically motivated specification is sufficient to demonstrate our main points (see also footnote 12).

[14] As we proxy the marginal cost in the standard way with the unit labor cost, it does not include the impact of imported inputs and is not correlated with any of the foreign variables. The same applies to the output gap. On the contrary, there is positive albeit weak correlation between some external factors. Lastly, given that the unit labor cost is only weakly correlated with the output gap, we include both variables at the same time to test their relative impact when external factors are controlled for.

[15] We use a Newey–West (1994) heteroscedasticity and autocorrelation (HAC) consistent covariance matrix estimator. The correlation in moment conditions is soaked up by means of the previous VAR(1) estimation for IV (pre-whitening). The Bartlett kernel is used to weight the covariances with the Newey–West fixed bandwidth. For the sake of comparability with previous studies, all the results were obtained by means of a common N-step (iterative) GMM estimator. However, the main findings also hold with a continuously updated GMM estimator (CUE) that is superior in small samples.

4.4 Data and Time Series Analysis

4.4.1 Data description

Our data set consists of quarterly data ranging from 1998/1.Q to 2007/3.Q but some series are slightly shorter. The principal source of the data is the OECD (Main Economic Indicators) and Eurostat. Some series were obtained from additional sources.

The inflation rate is measured by the harmonized consumer price index (HCPI, Eurostat, 2005=100, quarterly data were obtained as averages of the corresponding seasonally adjusted monthly data). We use a CPI-based inflation measure since (i) both monetary policy and inflation expectations of private subjects are usually defined in terms of consumption-based indices, (ii) it is the most relevant measure of inflation in small open economies (Galí and Monacelli, 2005 derive an open-economy NKPC in terms of CPI inflation), (iii) the inflation rate derived from the implicit GDP deflator, which is consistent with GG and GGL, presents some unusual developments that disqualify it for econometric analysis (see Figure A.1).[16] We use yearly (year-on-year) and quarterly (quarter-on-quarter) inflation rates. The results using the former are reported while with the latter are only commented on, as these seem to be subject to seasonality noise that cannot be fully removed by statistical techniques.[17] The CPI inflation *survey data* (available only for the Czech Republic and Poland) are the inflation expectation of the financial market participants (yearly, four quarters ahead, quarterly frequency calculated as the simple mean of the monthly data); the data come from the Czech and Polish National Banks. *The foreign inflation rate* is measured by the inflation rate in the Euro area. *The wage inflation* is approximated by two measures: (i) the unit labor cost index (OECD, 2000=100, total economy) that is calculated as the ratio of the (nominal) total labor cost and the real output (both in the national currency); (ii) the labor cost index reported by Eurostat (2000=100), which includes the total labor costs, main components of wages and salaries (e.g. bonuses) and non-wage costs (e.g. social contributions). The wage inflation is calculated as the yearly/quarterly change in each index. *The real marginal* cost is proxied by the real unit labor cost. We use two measures of the real unit labor cost: (i) the (log) unit labor

[16] See the quarterly change in the GDP deflator of Hungary in 2002 and Poland in 1999. The series of yearly inflation rates from the GDP deflator and HCPI are strongly correlated.

[17] Assenmacher-Wesche and Gerlach (2008) argue that inflation-forcing variables can vary according to inflation frequency. In particular, inflation at a high frequency (quarterly) is linked to the output gap while the low-frequency fluctuations (horizon of several years) are driven by monetary factors. In our case, the quarterly inflation rate may be influenced by price shocks while the yearly inflation rates are affected by the real economic activity (the marginal cost, the output gap).

cost index (OECD 2000=100, total economy) deflated by the price index (HCPI) and (ii) the (log) ratio of the nominal total compensation to employees and the nominal GDP (Eurostat, both series in euros and disaggregated from the annual frequency by Ecotrim software by Eurostat). The deviation from the steady state is defined as a deviation from the sample mean/HP trend. *The output gap* is: (i) the difference between the logarithm of the current value of the seasonally adjusted GDP (in millions of euros in 1995 prices, Eurostat) and the trend value obtained by the HP filter (the smoothing parameter set to 1600) and (ii) provided by the OECD by means of the production function approach. *The unemployment rate* is the standardized unemployment rate (Eurostat, calculated from the monthly frequency, seasonally adjusted). *The import prices* are proxied by (i) the ratio of import value and import volume indices (Eurostat, 2000=100, disaggregated from the annual frequency by Ecotrim) and (ii) the ratio of imports in current prices to imports in constant prices (OECD, in the national currency, quarterly data calculated as the average of the seasonally adjusted monthly series). The yearly/quarterly changes in each index are used for estimation. *The oil price* is measured by the average quarterly world price of a barrel of crude in USD (quarterly prices are calculated as the average from weekly values); the data come from the US Energy Information Administration. The yearly/quarterly change is used for estimation. *The exchange rate* is the nominal effective exchange rate index (Eurostat, against 12 main trading partners; 1999 is the base year). The yearly/quarterly change in the index is used for estimation.[18]

4.4.2 Identification in forward-looking model

Table 4.1 provides the results of the F-test for the joint coefficient restriction that $\rho_i = \varphi_j = 0$ for all *i*>1 and all *j*>2 in Eq. (8). The rejection of this restriction is a necessary condition for the identification in the GMM framework. We report the test using year-on-year HCPI inflation rates and quarter-on-quarter inflation rates from both HCPI and GDP deflators

[18] Although most empirical studies on the NKPC do not address the properties of the time series used, we remit all the variables to throughout unity root analysis. The conventional tests point to stationarity of all the variables except the inflation rate, the wage inflation and (log of) the marginal cost where the results are ambiguous. It is of particular interest to compare the degree of persistence of the inflation and the marginal cost. We find that the marginal cost is more persistent than inflation (just the opposite of what was found for developed countries). Consequently, we test the cointegration between these two variables because its long-term comovement is the main precondition for the reasonability of the NKPC (GGL show for several OECD countries that GDP inflation and the (log) unit labor cost closely comove in the long term). While cointegration is found between the quarterly inflation rates and the (log) marginal cost for all the countries, it is rejected when yearly inflation rates are used for the Czech Republic and Poland.

against the main domestic forcing variables (the marginal cost, the output gap, the wage inflation) to see that the choice of inflation measure does not alter the results.

The results confirm that, under the model for the forcing variable (8), the second measure of the real marginal cost (the ratio of nominal total compensation to employees to the nominal GDP) meets the necessary condition for identification in all the countries, no matter which measure of inflation is used. The reported results are for the untransformed series of the (log) marginal cost. The results using its deviation from the sample mean and the HP trend are very similar. As for the output gap series, the gap derived by the HP filter is preferable to the OECD gap in all the countries but Slovakia (the OECD gap for Slovakia is only available from 2001). Finally, both measures of wage inflation have problems meeting the necessary condition for identification.

Table 4.1: F-test of the necessary condition for identification in the forward-looking model (p–values)

	CZE			**HUN**			**POL**			**SVK**			**Panel**		
	t/t-1	t/t-1*	t/t-4	t/t-1	t/t-1*	t/t-4	t/t-1	t/t-1*	t/t-4	t/t-1	t/t-1*	t/t-4	t/t-1	t/t-1*	t/t-4
RULC1	0.00	0.03	0.01	0.01	0.02	0.01	0.00	0.00	**0.20**	**0.50**	0.17	**0.45**	0.00	0.00	0.00
RULC2	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
WAGE1	**0.20**	**0.67**	0.04	**0.07**	0.00	**0.44**	**0.10**	0.03	**0.11**	0.00	0.00	0.12	0.00	0.00	0.00
WAGE2	**0.08**	0.00	0.00	0.00	0.00	**0.19**	**0.19**	0.00	**0.19**	0.00	0.00	0.00	0.00	0.00	0.00
GAP1	0.00	0.00	0.00	0.02	0.00	0.00	0.01	0.20	**0.07**	**0.11**	**0.19**	0.00	0.00	0.00	0.00
GAP2	0.00	0.00	0.00	**0.23**	0.03	**0.30**	**0.05**	**0.79**	**0.12**	0.00	0.00	0.00	0.00	0.00	0.00

Note: t/t-1 – quarterly inflation rate (HCPI), t/t-1* – quarterly inflation rate (GDP deflator), t/t-4 – yearly inflation rate (HCPI), RULC1 – the (log) unit labor cost index deflated by HCPI (OECD), RULC2 - the ratio of the nominal total compensation to employees and the nominal GDP (Eurostat), WAGE1 - the ratio of (nominal) total labor cost and the real output (OECD), WAGE2 - the labor cost index (the total labor costs, main components wages and salaries and non-wage costs (Eurostat), GAP1 – GDP gap by HP filter (OECD), GAP2 – GDP gap by production function (OECD)

4.4.3 Cross correlations

Another important issue that must be considered in the analysis of inflation dynamics is the temporal effect of the variables. While traditional PC assumes that the output gap (or other cyclical measure) leads inflation, the opposite pattern is consistent with the NKPC (see Eq. (8) in GG). The dynamic cross correlations of the (log) real unit labor cost and the output gap with the HCPI inflation rates are reported in Table 4.2.[19] The negative correlation for lagged

[19] We report only the result for the second measure of the real unit labor cost (RULC2) and the HP output gap (GAP1) because they performed better in the identification test and are consequently used for the NKPC estimates reported in the paper. The results for quarterly inflation rates are rather similar.

values of inflation and positive for leading values implies that the variable leads inflation. This pattern, found for the output gap for the Czech Republic and less significantly for Slovakia, is consistent with the traditional PC but not the NKPC. On the other hand, the output gap has a very strong contemporaneous correlation with inflation in Poland, consistent with the NKPC. Finally, for Hungary the correlations of the output gap with both lags and leads of inflation are negative. As for the unit labor cost, it is very strongly contemporaneously correlated with inflation in Poland, consistent with the findings of GG with the US data. For all the other countries, both lags and leads of inflation are negatively correlated with inflation. The results using the deviation of the marginal cost from its mean or trend are not reported. The marginal cost deviation from its mean becomes highly contemporaneously correlated with inflation for Hungary and Slovakia and the marginal cost deviation from the HP mean is negatively correlated with inflation for all the countries.

Table 4.2: Dynamic cross-correlations of the (log) real unit labor cost (RULC2) and the output gap (GAP1) with lags and leads of yearly HCPI inflation

	CZE				**HUN**				**POL**				**SVK**			
	RULC2		GAP1		RULC2		GAP1		RULC2		GAP1		RULC2		GAP1	
	lag	lead	lag	lead	lag	lead	lag	lead	lag	lead	lag	lead	lag	lead	lag	lead
0	-0.45	-0.45	0.05	0.05	-0.56	-0.56	-0.19	-0.19	0.70	0.70	0.62	0.62	-0.15	-0.15	-0.30	-0.30
1	-0.54	-0.37	-0.06	0.15	-0.58	-0.54	-0.27	-0.15	0.73	0.59	0.47	0.68	0.00	-0.25	-0.33	-0.17
2	-0.56	-0.29	-0.16	0.20	-0.62	-0.52	-0.32	-0.14	0.76	0.51	0.31	0.65	-0.01	-0.33	-0.38	-0.02
3	-0.51	-0.22	-0.21	0.19	-0.68	-0.46	-0.28	-0.13	0.77	0.44	0.18	0.56	-0.01	-0.33	-0.37	0.12
4	-0.42	-0.19	-0.21	0.13	-0.75	-0.39	-0.18	-0.11	0.74	0.38	0.18	0.45	0.03	-0.41	-0.38	0.32
5	-0.33	-0.16	-0.17	0.05	-0.74	-0.31	-0.05	-0.10	0.69	0.34	0.18	0.35	0.06	-0.45	-0.38	0.49

Note: see Note of Table 4.1

4.5 Empirical Results

4.5.1 Czech Republic

The estimates of the marginal-cost-based hybrid NKPC provide rather mixed evidence. The mean deviation and trend deviation of the marginal cost are each statistically significant but hold a “wrong” negative sign. On the contrary, the untransformed marginal cost (the log real unit labor cost) is significant and positive. However, we note that this correlation can be spurious, as the cointegration test and dynamic cross correlations indicated. The fit of the model substantially decreases in the case of a pure forward-looking NKPC (fourth line of

Table 4.3).[20] The residual autocorrelation appears in several specifications, though it does not automatically disqualify the model. The coefficient of the output gap is significant and positive. Additional lags of inflation in the spirit of the traditional PC do not seem to provide any information.[21] More interestingly, the additional inflation lags do not affect the size and the significance of the forward-looking term. The coefficient of wage inflation is also significant. The fit of the model is similar for all the domestic forcing variables.

The middle panel shows the results of using survey data on inflation expectations instead of explicitly assuming rational expectations. If the survey data are truly collected before the present time *t*, endogeneity should not arise and the OLS should provide consistent estimates. The Durbin–Hausman–Wu test rejects the consistency of OLS estimates only marginally (at the significance level of 10%). The main feature of these OLS estimates is that the sum of the backward- and forward-looking terms is significantly above unity and the fit of the model is substantially lower. The GMM estimates, used to deal with potential endogeneity, have different magnitudes but retain their significance. Figure A.4.2 suggests that endogeneity seems to arise due to the forecast error. It is notable that the realized inflation deviates significantly from the expectations of the financial market's participants. Certainly, this issue should be studied more carefully but goes beyond the scope of this paper.

The lower panel reports the results of the hybrid NKPC when we retain the (log) marginal cost and the output gap (these variables do not have a very strong correlation in any country, so we can evaluate their relative importance without introducing multicollinearity) and stepwise add the open-economy variables. Here we find something very interesting. The coefficients of all the external variables are significant and have the expected sign. Yet, we do not interpret the size of their coefficients as model (7) is empirically motivated rather than being structural. The oil prices have a positive (but limited) effect on yearly inflation and the same applies to import prices. The effect of the inflation in the euro area seems to be

[20] We report the adjusted R squared to compare the relative goodness-of-fit of different specifications given that alternative forcing variables are used (the unit labor cost vs. the output gap) and some variables are excluded (forward-looking term) or added (four open economy variables). However, the appropriate way to assess the fit of the NKPC requires comparison of the actual inflation with the fundamental inflation that arises as a solution to the estimated forward-looking model (see GG). Unfortunately, our sample size does not allow us to obtain a sufficiently long stream of future real marginal costs needed for calculation of the fundamental inflation. Moreover, our empirical models use alternative forcing variables that cannot enter in the same way as the marginal cost does, i.e. as the discounted stream of future values.

[21] The reported estimate of the backward-looking term is the sum of the estimates of four lagged values of inflation. The reported standard error is a mean of the four estimated standard errors.

especially relevant. Finally, the domestic currency depreciation (decrease of NEER) leads to an increase in CPI inflation (the size of the coefficient points to a very incomplete exchange rate pass-through). While the output gap is significant and correctly signed in three cases, the real marginal cost is only in one. Lastly, when the potential external effect is accounted for, the effect of the forward-looking term is generally reduced, which is a finding that should be analyzed in detail.

Table 4.3: GMM estimates of different versions of the closed economy PC and the PC augmented by external variables – the Czech Republic

Forcing variable	γ_b (π_{t-1})	γ_f (π_{t+1})	λ ($rulc_t$)	λ_{in} (y_{t-1}, w_t^a)	λ_{ex} (x_{t-1}, x_t^b)	R^2	IVWD	J-stat
RULC (dev. mean)	**0.43**	**0.55**	**-0.07**			0.90	0.01	0.35
	(0.02)***	(0.03)***	(0.02)***				0.00	
RULC (dev. HP trend)	**0.43**	**0.54**	**-0.08**			0.90	0.00	0.26
	(0.02)***	(0.03)***	(0.02)***				0.00	
RULC	**0.42**	**0.61**	**0.21**			0.90	0.00	0.33
	(0.02)***	(0.03)***	(0.07)***				0.00	
RULC		**0.98**	0.14			0.58	0.12	0.78
		(0.10)***	(0.36)				0.00	
GAP	**0.56**	**0.37**		**0.23**		0.89	0.00	0.43
	(0.05)***	(0.03)***		(0.04)***			0.00	
GAP (4 lags of infl.)	**0.59**	**0.46**		**-0.17**		0.79	0.18	0.20
	(0.08)***	(0.03)***		(0.03)***			0.00	
WAGE	**0.43**	**0.47**		**0.05** [a]		0.88	0.00	0.33
	(0.02)***	(0.04)***		(0.01)***			0.00	
RULC + infl.exp.(OLS)		**1.25**	**2.37**			0.42	0.00[c]	
		(0.26)***	(1.07)**				0.00[c]	
RULC + infl.exp. (GMM)		**1.90**	**4.75**			0.44	0.00	0.94
		(0.28)***	(1.20)***				0.00	
RULC + infl.exp.(OLS)	**0.70**	**0.57**	1.42			0.73	0.03[c]	
	(0.11)***	(0.21)***	(0.74)*				0.17[c]	
RULC + infl.exp. (GMM)	**0.53**	**0.99**	**2.70**			0.76	0.14	0.68
	(0.08)***	(0.18)***	(0.71)***				0.00	
RULC + GAP	**0.52**	**0.41**	-0.19	**0.39**		0.90	0.00	0.31
	(0.02)***	(0.06)***	(0.14)	(0.10)***			0.00	
RULC + GAP + OIL	**0.69**	0.21	0.23	**0.52**	**0.01**[b]	0.85	0.00	0.68
	(0.08)***	(0.14)	(0.16)	(0.20)**	(0.00)***		0.00	
RULC + GAP + IMP2	**0.56**	0.13	-0.98	**0.43**	**0.12**[b]	0.81	0.00	0.72
	(0.06)***	(0.12)	(0.40)**	(0.18)**	(0.03)***		0.00	
RULC + GAP + π_{EMU}	**0.56**	**0.35**	**1.62**	**-0.13**	**0.80**[b]	0.89	0.00	0.57
	(0.04)***	(0.06)***	(0.38)***	(0.05)***	(0.16)***		0.00	
RULC + GAP + NEER	**0.53**	**0.20**	**-1.54**	**0.40**	**-0.13**	0.87	0.00	0.49
	(0.05)***	(0.14)	(0.49)***	(0.13)***	(0.03)***		0.00	

Note: Standard errors in parenthesis. *, **, *** denotes significance levels at 10, 5 and 1%, respectivelly. Yearly change of HCPI is always the dependent variable. In the first column, there are the forcing variables in each specification. Domestic forcing variables: RULC – log of the ratio of the nominal total compensation to employees and the nominal GDP (Eurostat), GAP (y_{t-1}) - GDP gap by HP filter (OECD), WAGE (w_t) – yearly change of the labor cost index (Eurostat). Foreign forcing variables (x_t): OIL – yearly change of crude oil price, IMP – yearly change of import price index, π_{EMU} – yearly change of HCPI in the Euro area, NEER – yearly change of nominal effective exchange rate. IVWD is p-value of Godfrey´s (1994) Wald test test for 1. and 4. order serial correlation for dynamic model estimated by instrumental variables ([c].indicates p-value of Breusch-Godfrey serial correlation LM test)

These results indicate that the marginal-cost-based NKPC might not be suitable for describing the Czech inflation dynamics and that the output gap could be a better proxy for domestic inflation pressures. The Czech inflation also seems to be determined by external factors and it is quite persistent. The results using quarterly inflation rates are very similar with the exception that no domestic forcing variable becomes significant.

4.5.2 Hungary

The identification of domestic forcing variables in the NKPC for Hungary is puzzling. Neither the marginal cost measure nor the wage inflation become significant. The output gap coefficient is significant albeit negative while the additional inflation lags are not significant.

Table 4.4: GMM estimates of different versions of the closed economy PC and the PC augmented by external variables – Hungary

Forcing variable	γ_b (π_{t-1})	γ_f (π_{t+1})	λ ($rulc_t$)	λ_{in} (y_{t-1}, w_t^a)	λ_{ex} (x_{t-1}, x_t^b)	R^2	**IVWD**	**J-stat**
RULC (dev. mean)	**0.60**	**0.40**	0.02			0.98	0.01	0.39
	(0.06)***	(0.08)***	(0.03)				0.62	
RULC (dev. HP trend)	**0.63**	**0.37**	-0.09			0.98	0.01	0.46
	(0.06)***	(0.06)***	(0.07)				0.00	
RULC	**0.59**	**0.43**	0.28			0.98	0.47	0.44
	(0.04)***	(0.06)***	(0.14)*				0.00	
RULC		**0.99**	-0.20			0.92	0.00	0.81
		(0.06)***	(0.70)				0.00	
GAP	**0.53**	**0.47**		**-0.50**		0.98	0.00	0.58
	(0.04)***	(0.04)***		(0.17)***			0.12	
GAP (4 lags of infl.)	**0.60**	**0.40**		**-0.14**		0.97	0.00	0.64
	(0.08)***	(0.02)***		(0.07)*			0.00	
WAGE	**0.62**	**0.42**		-0.03[a]		0.98	0.85	0.42
	(0.04)***	(0.04)***		(0.02)				
RULC + GAP	**0.58**	**0.43**	0.26	-0.04		0.98	0.70	0.41
	(0.04)***	(0.06)***	(0.14)*	(0.07)			0.11	
RULC + GAP + OIL	**0.55**	**0.48**	**0.45**	-0.47	**0.01**[b]	0.95	0.65	0.59
	(0.03)***	(0.04)***	(0.16)***	(0.19)**	(0.00)***		0.43	
RULC + GAP + IMP2	**0.46**	**0.66**	**0.72**	**0.55**	**-0.06**[b]	0.98	0.00	0.65
	(0.03)***	(0.04)***	(0.15)***	(0.10)***	(0.01)***		0.00	
RULC + GAP + π_{EMU}	**0.32**	**0.79**	**2.85**	0.55	**0.80**[b]	0.97	0.06	0.63
	(0.12)**	(0.15)***	(1.21)**	(0.24)**	(0.31)**		0.00	
RULC + GAP + NEER	**0.46**	**0.55**	-0.09	0.04	-0.03	0.98	0.91	0.59
	(0.04)***	(0.08)***	(0.44)	(0.23)	(0.03)		0.00	

Note: see Table 4.3

The lower panel of Table 4.4 reports the estimates of the NKPC with the open-economy variables. The oil prices as well as the inflation in the euro area seem to have a positive and significant impact. On the contrary, the coefficient of the import prices is contra-intuitively found to be significant and negative, while the exchange rate does not have a significant impact. Although the (log) real unit labor cost is statistically significant in three specifications augmented with open-economy variables, we shall recall that the cross correlations with inflation rates were in fact negative. Unlike in the Czech Republic, the relative size of backward- vs. forward-looking terms is not affected when the external variables augment the model. We can again conclude that the inflation in Hungary has a significant forward-looking component but it is also quite persistent and driven by external rather than domestic factors. The results using quarterly inflation rates confirm as in the Czech case the predominance of foreign inflation factors while neither the marginal cost nor the output gap is significant.

4.5.3 Poland

Poland is the only country in the sample where both the real unit labor cost (untransformed series in logs and the deviation from the mean) and the output gap are strongly correlated with lags and leads of inflation. However, the AEG test still suggests that the correlation between the inflation rate and the (log) marginal cost might be spurious.

The NKPC estimates for Poland are reported in Table 4.5. The estimates of the hybrid model resemble the findings for large closed economies such as the US. That is, the marginal cost is significant (both in the logs and in the perceptual deviation from the sample mean) and the forward-looking term is dominant. While the coefficient of the output gap is significant, that of wage inflation is not and the additional inflation lags do not seem to carry any additional information either.

The middle panel shows the estimates with the inflation survey data. In Figure A.4.2, we can appreciate that the inflation expectations in Poland move very closely with the actual inflation rates. This can be a sign either of a very good forecast realized by the financial markets or, more likely, of problematic construction of this series. This feature is evident if we compare the series with the expectations in the Czech Republic, which more logically show a pattern of forecast error.[22] The strong correlation of the inflation expectation with the actual inflation seems to drive the estimates provided in the middle panel. The coefficient of the forward-looking term is always significant but the coefficient of the marginal cost has a “wrong”

[22] The time series for both countries are plotted together with the actual HCPI inflation. Note that the series of inflation expectations are moved forward and in each quarter we can compare the actual inflation against its expected value (one year ago).

negative sign. The Durbin–Hausman–Wu test points to the presence of endogeneity (at the 10% significance level) in the hybrid model but not in the forward-looking model.

Table 4.5: GMM estimates of different versions of the closed economy PC and the PC augmented by external variables – Poland

Forcing variable	γ_b (π_{t-1})	γ_f (π_{t+1})	λ ($rulc_t$)	λ_{in} (y_{t-1}, w_t^a)	λ_{ex} (x_{t-1}, x_t^b)	R^2	LB	J-stat
RULC (dev. mean)	**0.42**	**0.62**	**0.01**			0.98	0.07	0.70
	(0.03)***	(0.03)***	(0.01)*				0.00	
RULC (dev. HP trend)	**0.47**	**0.56**	**-0.03**			0.98	0.02	0.40
	(0.02)***	(0.03)***	(0.03)				0.00	
RULC	**0.45**	**0.60**	**0.16**			0.98	0.00	0.70
	(0.02)***	(0.02)***	(0.05)***				0.00	
RULC		**0.97**	0.08			0.93	0.00	0.77
		(0.02)***	(0.18)				0.00	
GAP	**0.46**	**0.56**		**0.09**		0.98	0.00	0.68
	(0.02)***	(0.03)***		(0.03)***			0.00	
GAP (4 lags of infl.)	**0.56**	**0.44**		**0.06**		0.97	0.00	0.54
	(0.05)***	(0.02)***		(0.02)***			0.00	
WAGE	**0.47**	**0.61**		-0.02^a		0.98	0.00	0.50
	(0.02)***	(0.02)***		(0.01)*			0.00	
RULC + infl.exp.(OLS)		**0.87**	**-0.38**			0.98	0.50^c	
		(0.02)***	(0.15)**				0.21^c	
RULC + infl.exp. (GMM)		**0.84**	**-0.26**			0.97	0.16	0.80
		(0.01)***	(0.08)***				0.00	
RULC + infl.exp.(OLS)	-0.17	**1.02**	**-0.37**			0.98	0.61^c	
	(0.13)	(0.13)***	(0.15)				0.23^c	
RULC + infl.exp. (GMM)	**-0.53**	**1.39**	**-0.54**			0.97	0.79	0.82
	(0.11)***	(0.10)***	(0.12)***				0.00	
RULC + GAP	**0.45**	**0.61**	**0.1/**	0.00		0.98	0.00	0.57
	(0.02)***	(0.03)***	(0.08)**	(0.05)			0.00	
RULC + GAP + OIL	**0.45**	**0.67**	**0.42**	-0.05	0.00^b	0.96	0.00	0.80
	(0.06)***	(0.10)	(0.14)**	(0.07)	(0.00)*		0.00	
RULC + GAP + IMP2	**0.39**	**0.68**	0.12	**-0.01**	**-0.02**^{b}	0.97	0.02	0.65
	(0.04)***	(0.06)***	(0.10)	(0.07)	(0.01)**		0.00	
RULC + GAP + π_{EMU}	**0.50**	**0.55**	**2.67**	0.15	**1.22**^{b}	0.97	0.00	0.83
	(0.04)***	(0.05)***	(0.46)***	(0.07)**	(0.21)**		0.00	
RULC + GAP + NEER	**0.41**	**0.67**	0.24	**-0.04**	**-0.01**	0.97	0.00	0.69
	(0.02)***	(0.05)***	(0.11)**	(0.07)	(0.01)*		0.00	

Note: see Table 4.3

In the lower panel, we evaluate the effect of the open-economy variables. While the oil prices do not have a significant impact, the coefficients of the import prices have a wrong negative sign. The effect of the foreign inflation is the most substantial. The results based on the quarter-on-quarter inflation rate are similar to those for the previous two countries. Neither the marginal cost nor the output gap has a statistically significant effect and all the foreign variables are significant with the expected sign. The results indicate that the NKPC is relatively successful in explaining the Polish inflation dynamics. Moreover, the external variables still seem relevant, though the Polish economy is relatively larger in size than the other three countries.

4.5.4 Slovakia

The HCPI inflation series for Slovakia has some distinct features compared with the other countries. It is less similar to inflation from the GDP deflator and it is more volatile and spiky (see Figure A.4.1). A possible explanation could be that the Slovak consumer prices are more affected by the changes in imported consumer goods as Slovakia has the highest share of imports of the GDP.

The NKPC estimates are reported in Table 4.6. Each transformation of the unit labor cost series provides a very different result. The deviation from the mean is insignificant, the deviation from the HP trend is significant and negative and the (log) marginal cost is significant and positive. The pure forward-looking model again has a substantially lower fit. The output gap enters significantly albeit with a negative sign. Neither the additional inflation lags not the wage inflation seem to play any role.

In the lower part of the table, we can see that three of the four variables that underpin the external factors are significant. The results using quarter-on-quarter HCPI series (not reported here) feature a very low fit of the model.

Table 4.6: GMM estimates of different versions of the closed economy PC and the PC augmented by external variables – Slovakia

Forcing variable	γ_b (π_{t-1})	γ_f (π_{t+1})	λ ($rulc_t$)	λ_{in} (y_{t-1}, w_t^a)	λ_{ex} (x_{t-1}, x_t^b)	R^2	LB	J-stat
RULC (dev. mean)	**0.55**	**0.46**	-0.01			0.85	0.81	0.53
	(0.03)***	(0.03)***	(0.01)				0.00	
RULC (dev. HP trend)	**0.48**	**0.52**	**-0.15**			0.85	0.42	0.73
	(0.02)***	(0.03)***	(0.06)***				0.00	
RULC	**0.55**	**0.51**	**0.31**			0.85	0.00	0.61
	(0.02)***	(0.03)***	(0.15)*				0.00	
RULC		**1.26**	0.72			0.45	0.09	0.90
		(0.08)***	(0.55)				0.00	
GAP	**0.53**	**0.48**		**-0.11**		0.85	0.33	0.69
	(0.02)***	(0.03)***		(0.02)***			0.00	
GAP (4 lags of infl.)	**0.65**	**0.38**		**-0.11**		0.82	0.00	0.65
	(0.05)***	(0.04)***		(0.03)***			0.00	
WAGE	**0.49**	**0.59**		**-0.05**[a]		0.85	0.00	0.81
	(0.02)***	(0.04)***		(0.01)***			0.00	
RULC + GAP	**0.55**	**0.48**	0.14	**-0.06**		0.84	0.00	0.57
	(0.03)***	(0.03)***	(0.19)	(0.05)			0.00	
RULC + GAP + OIL	**0.59**	**0.45**	0.34	-0.06	**0.01**[b]	0.83	0.95	0.69
	(0.03)***	(0.03)***	(0.17)*	(0.04)	(0.00)***		0.00	
RULC + GAP + IMP2	**0.46**	**0.56**	0.16	-0.05	**0.04**[b]	0.88	0.00	0.58
	(0.06)***	(0.05)***	(0.14)	(0.03)	(0.03)		0.00	
RULC + GAP + π_{EMU}	**0.65**	**0.33**	1.30	-0.03	**0.61**[b]	0.81	0.41	0.65
	(0.03)***	(0.03)***	(0.52)**	(0.05)	(0.17)***		0.00	
RULC + GAP + NEER	**0.54**	**0.59**	0.38	-0.15	0.15	0.77	0.00	0.48
	(0.05)***	(0.05)***	(0.20)*	(0.06)**	(0.06)**		0.00	

Note: see Table 4.3

4.5.5 Discussion

The previous country-specific results allow us to draw some general conclusions. There is some evidence in favor of the NKPC in four CECs. In particular, the inflation rates in the four analyzed countries hold significant forward-looking components and therefore the current inflation is (at least partially) determined by its future expected value. However, the backward-looking term is also significant and often has a higher magnitude than the forward-looking term. This is consistent with the result of studies using both micro (e.g. Konieczny and Skrzypac, 2005, Babetski et al., 2007) and macro data (Lendvai, 2005, Franta et al., 2007, Mihailov et al., 2010), showing that inflation in the NMSs is more persistent. Nevertheless, the backward-looking coefficient can be upward-biased because (i) the quarterly inflation data are constructed from monthly data, which include some degree of autocorrelation, and (ii) our empirical model does not take into account the slightly decreasing inflation trend (see Cogley and Sbordone, 2008, Kim and Kim, 2008).

The estimates accompanying the forcing variable are more ambiguous. The marginal cost is significant only when we use its absolute value (in logs) rather than its deviation from the steady state. Yet, as we pointed out, neither the mean nor the HP trend have to be a good proxy for the steady state. In any case, the unit labor cost is a good proxy for the aggregate marginal cost only under several restrictive assumptions that in reality might not hold (e.g. that the production technology is consistent with the simple Cobb–Douglas production function).

The potential superiority of the output gap over the marginal cost found in the Czech Republic and Poland has two different interpretations. The first one is that the cyclical evolution of the output has a stronger effect on inflation than suggested by the NKPC. However, once we also include the additional lags of inflation, in the spirit of the traditional PC, they turn insignificant and their sum is never close to unity. Second, given that the forward-looking term is mostly significant, another interpretation of the result can be that the NKPC holds and the output gap is a better proxy for the unobservable marginal cost. The unit labor cost may not be truly sufficiently representative as a measure of firms' costs when a large share of their inputs is imported. The relevance of the output gap is also reported by Genberg and Pauwels (2005), Jondeaua and Le Bihanb (2005), Henzel and Wollmershäuser (2008) and Zhang et al. (2009).

The claim that the inflation dynamics of small open economies is driven by external impulses is supported by estimates of the NKPC augmented by open-economy variables that are usually significant and have the correct sign. This is consistent with Stavrev (2009), who decomposes inflation in the 10 NMSs into common and country-specific components by a generalized dynamic facto model, and Mihailov et al. (2010), who estimate the NKPC augmented by the change in terms of trade. [23]

[23] Vizek and Broz (2009) report similar findings (higher inflation inertia and importance of external factors) for Croatia as a current EU candidate country.

4.6 Concluding Remarks

In this paper, we have explored the inflation dynamics of four CECs (the Czech Republic, Hungary, Poland and Slovakia) by means of econometric estimation of the NKPC. The NMSs are very specific among open EMEs, given their previous transitional experience, their high degree of economic openness and their convergence process with the EU. This is why it is very interesting to learn about the nature of the inflation dynamics of these economies. We extend the existing, though rather scant, empirical evidence for the NMSs using a new data set and a more comprehensive analysis. In particular, we accompany the empirical framework of GG and GGL by additional tests, use alternative definitions of inflation, employ both a rational-expectations-based framework and the survey data and test the effect of alternative forcing variables including open-economy variables. Given the recent contributions pointing to the instability of inflation dynamics across policy regimes (Benati, 2008, Cogley and Sbordone, 2008), we focus on the post-transitional period (1998–2007), which was free of major changes in monetary policy regimes, and where the inflation series were not subject to a structural break.

Our results are in general terms favorable for the NKPC. In particular, we have found strong evidence that inflation is determined by future inflation expectations. However, the CECs exhibit a higher degree of inflation persistence than was found for developed economies. An intuitive explanation could be that many firms in the CECs still employ simple backward-looking price setting, which is consistent with adaptive rather than rational expectations. This could be caused by a lack of credible monetary policy or a missing nominal anchor in some countries. When the monetary policy is unable to anchor the agents' inflation expectations, they logically prefer to use the past information. Orphanides and Williams (2004) suggest that expectations formation can be conditioned by the learning process about monetary policy. This sheds some doubt on the suitability of the NK framework, where backward-looking price setting is clearly suboptimal and the welfare loss increases with the rising share of the backward-looking firm. On the other hand, some recent contributions question the link between the backward-looking inflation term and genuine inflation persistence.[24]

[24] Sheedy (2010) shows that the intrinsic inflation persistence appears in an environment where newer prices are stickier than older ones. In his model, inflation persistence can arise even if the price setting is purely forward-looking. Blanchard and Galí (2007) argue that inflation persistence is consistent with a standard NK model where real wage rigidities are taken into account. Cogley and Sbordone (2008) and Kim and Kim (2008) find for

The identification of the inflation-forcing variable remains the biggest puzzle. We have not found unambiguous evidence that the average real marginal cost (proxied by the real unit labor cost) plays such a role. The output gap performs only marginally better. Although the statistical insignificance of variables such as the unit labor cost or the output gap is problematic for the NKPC, it does not automatically disqualify it given that these variables are only noisy proxies of the model variables. In addition, we obtain some indications that the short-term inflation impulses in the CECs can be rather external. This is not surprising given the degree of economic openness of these countries. Nevertheless, most of the current inflation seems to be intrinsic, i.e. it is linked either to the past inflation or to its future expectations. Due to the fragile statistical significance of some estimates, we limit our attention to the reduced-form estimates rather than the structural parameters of the NKPC (e.g. Calvo's price staggering parameter θ).[25]

Our findings have several potential implications for monetary policy. First, the fact that inflation in the NMSs has a significant backward-looking component suggests that inflation is a persistent process, its current value depends strongly on the past ones and it is affected by monetary policy only with substantial time lags. In this case, inflation stabilization is believed to go along with an output loss. On the contrary, the standard forward-looking NK framework implies no trade-off between inflation and output gap stabilization, though, as pointed out by Blanchard and Galí (2007), this *divine coincidence* disappears once real imperfections, e.g. real wage rigidities, are introduced to the model. Second, the structural model behind the NKPC suggests that the effect of monetary policy on inflation occurs via the marginal cost. However, if the marginal cost is not an inflation-driving variable, the monetary policy can influence inflation only via its credibility and its effect on inflation expectations. This claim seems plausible for the inflation targeters. Finally, if the inflation dynamics of the current EMU members is consistent with the NKPC (as reported by GGL), it is of a slightly different nature from the inflation dynamics of the four NMSs. While the former is marginal cost driven and forward-looking, the latter has an important backward-looking component (similar to that reported for the US by GG) and is affected by external factors. Nevertheless, given that prices in the NMSs seem to adjust to prices in the euro area and that the central banks of the

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the US that once the structural breaks (e.g. changes in the steady-state inflation rate) are accounted for, the backward-looking inflation component dissipates.

[25] Chari et al. (2009) claim that a structural model of inflation dynamics must be consistent with micro evidence on the price setting while the NKPC is not.

NMSs have gained substantial credibility over the last decade, it is likely that the inflation dynamics of the NMSs will converge with the current EMU members as well. Therefore, an interesting agenda for future research can be to evaluate the inflation dynamics of the NMSs in a time-varying framework. This could allow the testing of whether the forward-looking component truly strengthens (and inflation persistence decreases) as a result of a more credible monetary policy or whether the external inflation factors gradually displace the domestic ones (as the globalization hypothesis suggests). On the theoretical side, additional open economy extensions of the NKPC could be explored to take into account both domestic and external inflation factors, in the latter case both those related and those unrelated to price setting behavior.

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Appendix

Figure A.4.1: Yearly inflation rates (left) and annualized quarterly inflation rates (right) derived from HCPI (solid line) and GDP deflator (dotted line)

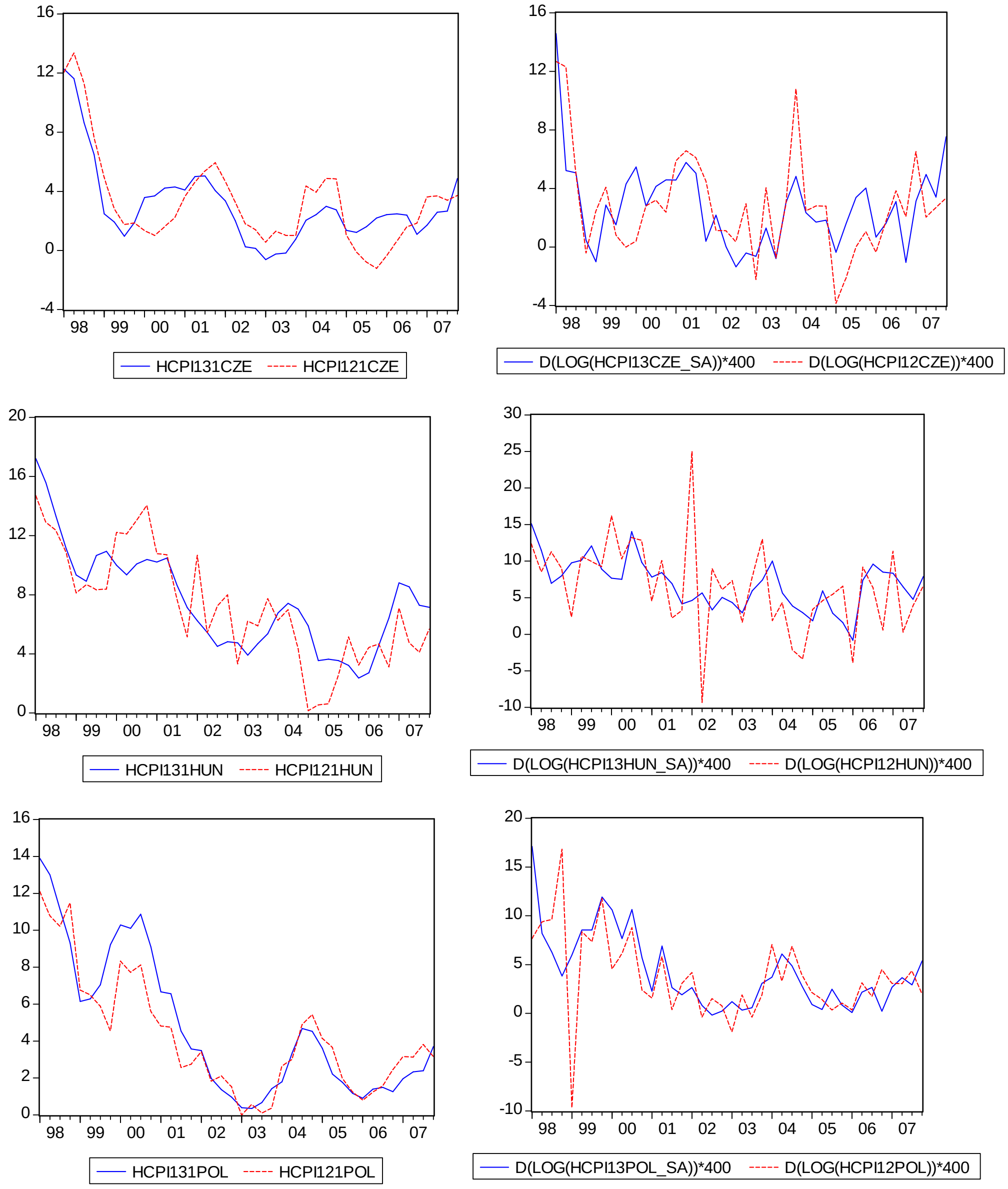

Figure A.4.2: Yearly rates HCPI inflation (solid line) and inflation expectations 12 months ahead by financial markets (dotted line) – the Czech Republic (left) and Poland (right)