WORKING PAPER SERIES 5

2011

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WORKING PAPER SERIES

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5/2011

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Is Monetary Policy in the New EU Member States Asymmetric?

Bořek Vašíček*

Abstract

Estimated Taylor rules have become popular as a description of monetary policy conduct. There are numerous reasons why real monetary policy can be asymmetric and estimated Taylor rules nonlinear. This paper tests whether monetary policy can be described as asymmetric in three new European Union (EU) members (the Czech Republic, Hungary, and Poland), which apply an inflation targeting regime. Two different empirical frameworks are used: (i) Generalized Method of Moments (GMM) estimation of models that allow discrimination between sources of potential policy asymmetry but are conditioned by specific underlying relations, and (ii) a flexible framework of sample splitting where nonlinearity enters via a threshold variable and monetary policy is allowed to switch between regimes. We find generally little evidence for asymmetric policy driven by nonlinearities in economic systems, some evidence for asymmetric preferences, and some interesting evidence on policy switches driven by the intensity of financial distress in the economy.

JEL Codes: C32, E52, E58.

Keywords: Inflation targeting, monetary policy, nonlinear Taylor rules, threshold estimation.

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I am grateful to Martin Čihák, Branislav Saxa, Pierre Siklos, Tsvetomira Tsenova, two anonymous referees, and the participants of the 28th GdRE International Annual Symposium on Money, Banking and Finance, whose comments substantially improved the quality of my manuscript.

Nontechnical Summary

Estimated monetary policy rules have become a convenient way to approximate monetary policy conduct. Empirical studies originally estimated linear policy rules assuming that monetary policy responds symmetrically to economic developments. However, there are numerous reasons why real monetary policy can have asymmetric features. The sources are usually characterized as either exogenous or endogenous to monetary policy. The former case refers to different kinds of nonlinearities present in the economic system which oblige optimizing central bankers to behave asymmetrically. In the latter case, asymmetric monetary policy can be driven by genuinely asymmetric preferences of central bankers. For example, inflation targeting central banks may respond, for reputation reasons, more actively when expected inflation exceeds its target value. Empirical studies have confirmed the existence of asymmetric policy conduct in many major central banks.

This paper provides extensive testing for the existence of asymmetric monetary policy in three new EU member states that apply inflation targeting (the Czech Republic, Hungary, and Poland). Two alternative empirical frameworks are used: (i) estimation of empirical models that allow discrimination between sources of potential policy asymmetry but are conditioned by specific underlying relations, and (ii) a flexible threshold model where policy asymmetry enters by means of threshold effects splitting the policy rule into two regimes.

Our main results can be summarized as follows. First, we find generally little evidence for asymmetric policy driven by nonlinearities in economic systems. Second, there is some indication of asymmetric policy due to asymmetric preferences of central banks in terms of inflation and interest rate volatility. Third, when testing asymmetric policy by means of the threshold model with inflation, the output gap, and a financial stress indicator as competing threshold variables, we find the most consistent threshold effects with the degree of financial stress. This suggests that central banks handle monetary policy slightly differently at times of financial instability.

1. Introduction

The monetary policy setting in the Central and Eastern European countries (CEECs) evolved substantially during the economic transition. These countries experimented with diverse monetary policy and exchange rate frameworks until the late 1990s, when their policy regimes fell into line with the then influential bipolar view, i.e., that intermediate regimes between hard exchange rate pegs and free floating are not sustainable. Some countries (the Baltic States and Bulgaria) adopted hard pegs, which put a significant constraint on their monetary policy, while other economies decided to maintain an overall flexible exchange rate, allowing their central banks to pursue internal macroeconomic targets (the Central European countries and Romania). Ongoing nominal and real convergence coupled with EU membership and the obligation to meet the Maastricht criteria put another constraint on policy making in general and monetary policy in particular in the New Member States (NMS). Some countries have merely formalized their previous exchange rate pegs by means of participation in the Exchange Rate Mechanism II (ERM II) and consecutive euro adoption, while others have retained their monetary policy autonomy under the framework of inflation targeting (IT) to the present day. Given the relative success of the latter countries in achieving price stability with decent levels of economic growth, it is of interest to understand their monetary policy conduct in greater detail. In particular, it seems interesting to explore empirically interest rate setting behavior under the IT mandate as well as the subtle differences between these countries.

There is a vast amount of empirical research on the way central banks handle interest rate setting. Since Taylor (1993), researchers have been estimating Taylor rules, as they seem to characterize well the interest rate setting of central banks. 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. The theoretical underpinning of the linear policy rule is the linear-quadratic (LQ) representation of macroeconomic models, with the economic structure assumed to be linear and the policy objectives to be symmetric, as represented by a quadratic loss function (e.g., Clarida et al., 1999). However, when the assumptions of the LQ framework are relaxed, the optimal monetary policy can be asymmetric. Asymmetric monetary policy implies that the monetary policy rule, which is a schematization of the policy reaction function, is nonlinear. In reality, however, asymmetric monetary policy can arise even when the underlying relations are essentially linear but the policy responses (slope elasticities) are different for positive and negative shocks. Unfortunately, owing to difficulties with shock identification, most empirical research relates asymmetric policy only with departures from the LQ framework and, therefore, nonlinear underlying relations.

Departures from the LQ framework involve two different sources of policy asymmetry. The first source lies in nonlinearities in the economic system. A common example of such nonlinearity is a steeper inflation-output trade-off when the output gap is positive. Such 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 IT, 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). Looking at monetary policy decisions from the risk management perspective, it seems plausible that central banks would like to avoid tail risk, which implies a disproportional response to certain vulnerabilities bringing about asymmetric policy responses. For example, deflationary risks in the US around 2003 could be seen as a factor behind its policy rate hovering around 1% for a rather extended period. The CEECs may also be more vulnerable to certain risks, such as those stemming from other emerging countries, e.g., the 1998 Russian crisis. In general terms, real monetary policy conduct seems to be too complex to be described by a simple linear equation, and nonlinear representation of monetary policy may be more appropriate irrespective of its underlying sources.

Several empirical studies have provided evidence that the monetary policy setting of many central banks may really be characterized as asymmetric. An 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, the Bundesbank, and the Bank of France responded more actively to inflation during economic booms. Leu and Sheen (2006) and Karagedikli 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 contractions than expansions and that the level of the interest rate itself was a source of policy asymmetry. Surico (2007b) establishes similar evidence of the Fed's asymmetric response to the output gap in the pre-Vocker era and quantifies the inflation bias induced by such policy. Asymmetries due to convexity of the PC found in some European countries (Dolado et al., 2005) and the ECB (Surico, 2007a) were linked to wage rigidity in European countries.

A few studies (Maria-Dolores, 2005; Frömmel and Schobert, 2006; Mohanty and Klau, 2007; Paez-Farrell, 2007; Vašíček, 2010) provide some evidence of linear monetary policy rules of the CEECs. However, some narratives suggest that monetary policy may also be asymmetric in these countries. In particular, inflation targeters may show anti-inflationary bias and therefore asymmetric policy due to reasons of reputation. The case for anti-inflationary bias could arguably be even stronger in these countries, as the announced inflation targets, unlike in most developed countries, had a downward-sloping trend and the countries were de facto targeting disinflation. Under such circumstances, the central banks often declare multi-year targets (Mishkin and Schmidt-Hebbel, 2001) and can treat the bottom of the (short-term) inflation target band or the (short-term) point target undershooting more leniently in order to approaches faster the (lower) long-term inflation target. However, Jonáš and Mishkin (2004) argue that such opportunistic approach to disinflation (Orphanides and Wilcox, 2002) can make monetary policy less predictable, which is in fact problematic for IT credibility. The empirical evidence on asymmetric monetary policy setting in the NMS is very limited. Horváth (2008) employs simple subsample analysis along the sign of the deviation of inflation from the target for the Czech National Bank and finds that monetary policy was asymmetric in the first years after the adoption of IT and symmetric afterwards. The reason was arguably the need to gain credibility and to anchor inflation expectations. On the other hand, IT is flexible enough to allow policy makers not to contract demand when inflation is slightly above the target and the shocks are likely to be shortlived (Blinder, 1997). Similarly, it seems plausible that other concerns, such as economic growth and financial stability, can lead to the temporary dismissal of inflation targets. By extending Surico's (2007a,b) model to small open economies (including the exchange rate vis-à-vis the euro area as a policy objective) Ikeda (2010) finds two conflicting policy asymmetries in Visegrad countries: an aversion to interest rates above the reference value (requiring an expansionary stance) and a preference for exchange rate appreciation relative to the euro area (requiring a policy tightening). However, it seems useful to explore diverse sources of policy asymmetries in a more general setting rather than sticking to a particular source of policy asymmetry (inflation) or a particular model.

In this paper, we test the hypothesis of asymmetric monetary policy in three Central European NMS: the Czech Republic, Hungary, and Poland, who have adopted the IT framework and maintain a flexible exchange rate.[1] We employ two empirical frameworks to test for policy asymmetry: (i) a framework based on an underlying structural model that modifies the LQ framework, which allows discrimination between sources of policy asymmetry but is conditioned by the specific model setting; and (ii) a flexible econometric framework where monetary policy is allowed to switch between two regimes according to a threshold variable. Besides the common choices for the threshold variable, such as the deviation of inflation from the target and the business cycle stance, we use a variable that tackles potential policy asymmetry along positive and negative shocks. In particular, we focus on the degree of financial stress in the economy to see whether central banks behave differently when the economy is distressed.

We find mixed evidence in terms of asymmetric behavior of central banks. First, we find no evidence of asymmetric policy driven by nonlinearities in the economic system. Second, we obtain some indications of asymmetric policy driven by preferences, in particular in terms of inflation and the actual deviation of the interest rate from its long-term equilibrium value. Third, we detect possible policy switches driven by the degree of financial distress in the economy. In particular, central banks seem to alter their monetary policy stance when the economy is faced by severe financial stress. Our empirical findings imply that monetary policy in the three NMS is possibly handled in a slightly different fashion than the legally grounded symmetric IT mandate suggests. It seems especially interesting to further analyze whether financial stability concerns affect policy-making as an (implicit) target or are considered because of their potential effect on inflation.

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 for policy asymmetry, and in section 4 we present our dataset. In section 5, we review the empirical results. The final section concludes.

[1] While the Czech Republic and Poland applied inflation targeting over the whole sample, Hungary adopted it only in 2001. However, given that anti-inflationary policy was well under way before official inflation targeting was adopted, due to homogeneity of analysis as well as the size of the data sample we do not treat 2001 as the a priori moment of regime switch in Hungary. In fact, for similar reasons we do impose any other a priori regime switching, for example, when the central bank leadership or broader institutional environment changed (e.g., EU entry). Nevertheless, although the data sample does not allow multiple structural breaks, the threshold estimation applied in part of our analysis allows splitting into two regimes along different potential threshold variables.

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), 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 output:

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

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

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 shocks, respectively. Eq. (1) represents the aggregate supply (AS) schedule or the PC, and Eq. (2) is the intertemporal 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 the general functional form, which can be quadratic if 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 the 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 where g 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 want 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 combinations of functional forms (1)–(3) give rise to different versions of the 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 domestic inflation π_t and output y_t and the relations themselves can be subject to structural change. Therefore, an alternative is to use an empirical framework that tracks asymmetries in monetary policy setting but does not rely on the specific structure of the model.

3. Empirical Testing of Asymmetric Monetary Policy

There are diverse empirical strategies for testing for monetary policy asymmetry. They typically consist of 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):

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

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 the policy decision, and ε_t is the error term. Given that the 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 for 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}:

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

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

3.1 Nonlinearities in the Economic System

Monetary policy asymmetry can be related to nonlinearities in the economic system. For example, (upward) nominal price stickiness can drive a nonlinear trade-off between inflation and

[2] The choice of instruments we considered as a benchmark is taken from the study of Clarida et al. (1998, 2000), including additionally the foreign (euro) interest rate. However, to avoid the problem of weak instruments pointed out in Stock and Yogo (2005) we limited the number of variables as well as the number of lags. We also tested different combinations of variables and lags, but failed to find any substantial differences. In addition, there is some controversy about additional variables (regressors) that can affect interest rate decisions. In particular, small open economies may adjust the interest rate to the exchange rate or international interest rates, for example. However, the three NMS use IT, where domestic price stability is the only official policy target. Moreover, there is no evidence that the Hungarian and Polish central banks respond to any additional variable (Vašíček, 2010). Although the interest rate of the euro area sometimes turns significant in the estimated policy rule of the Czech National Bank (Horváth, 2008; Vašíček, 2010), it is puzzling whether this means a genuine aim to stabilize domestic interest rate vis-à-vis the euro area or 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 an instrument rather than a regressor.

output. Dolado et al. (2005) derive a nonlinear monetary policy rule when the PC is convex. Under such model specification the linear policy rule, such as Eq. (5), is augmented by an interaction term of expected inflation and the output gap. Given that any inflationary pressures during economic expansions are larger if the PC is convex, there must be an additional interest rate increase whenever the inflation is above target or the output gap is positive.

To implement this framework empirically, we estimate in the first step a very simple backward-looking PC to understand the nature of the inflation-output trade-off:

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

where the present inflation rate π_t depends on its lagged value π_{t-1} and the lagged output gap y_{t-1}. The PC is nonlinear when coefficient ϕ is significantly different from zero. In particular, it is convex when $\phi > 0$ and concave when $\phi < 0$. 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 a positive and statistically significant value of the coefficient accompanying the interaction term of inflation and output κ is evidence of rule asymmetry. In particular, the increase in the interest rate is more than proportional when inflation is above the defined target or the output gap is positive. The logic is the following. When expected inflation is above its target, the current real interest rate is below its equilibrium value. This in turn implies a higher output gap in the short run, which feeds through to additional inflationary pressures in the longer run. If there is a convex relationship between the output gap and inflation, these future inflationary pressures will be stronger and monetary policy must adopt an additional interest rate increase corresponding to the size of the deviation of inflation from the target as well as the output gap.[3]

3.2 Asymmetric Preferences of the Central Bank

Asymmetric preferences with respect to economic outcomes represent another rationale for why central banks can behave asymmetrically. They may disproportionally decrease the interest rate when output is below its potential (to prevent further recession) or increase it when inflation exceeds the specified target (for credibility reasons). In countries such as the NMS where IT was adopted as a disinflation strategy, the narratives (e.g. Jonáš and Mishkin, 2004) suggest that central banks were prone to behave asymmetrically due to temporal existence of multi-year targets. Therefore, they were lenient to the (short-term) target undershooting in order to approach faster the long-term target, which represented the price stability. Orphanides and Wilcox (2002) argue that this opportunistic approach to disinflation implies path dependence in monetary policy making. In particular, policy makers' reaction to a given inflation level depends on the prior inflation history. Dolado et al. (2004) show that under asymmetric preferences (represented by a

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

linex function), the optimal policy rule is nonlinear irrespective of the form of the AS schedule. In their model the central bank can assign different weights to positive and negative deviations of inflation from the target and, therefore, the central bank's loss depends not only on the size of the deviation of inflation from the target, but also on its sign. In this setting, optimal monetary policy is asymmetric and the derived monetary policy rule contains inflation volatility (conditional variance) as an additional regressor. Given that the conditional inflation variance depends nonlinearly on lagged inflation and output, the interest rate in fact becomes a nonlinear function of lagged inflation and output. Therefore, asymmetric preferences imply that central banks respond for prudential reasons not only to inflation, but also to its volatility.

This model can be brought to the data as follows. First, the conditional inflation variance can be obtained from the estimated PC. If the conditional variance is time varying, the residuals of the PC (Eq. (6)) will contain autoregressive conditional heteroskedasticity (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. Consequently, we reestimate the (linear or nonlinear) PC allowing for ARCH effects in the residuals. In particular, we use the common GARCH (1,1) model with the variance equation defined as:

$$\sigma^2_{\pi,t} = \omega_\pi + \nu_1 \xi^2_{\pi,t-1} + \nu_2 \sigma^2_{\pi,t-1} \tag{8}$$

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

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

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 a nonlinear PC, which leads to an exponential monetary policy rule. The way to bring such a nonlinear equation to the data is by linearization using Taylor series approximation around points where the asymmetry-driving parameters are zero. This results in the policy rule:

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

[4] It should be noted that like Dolado et al. (2004) we use the contemporary inflation variance even though the rule is forward-looking in inflation. The reason is that it seems highly unrealistic to assume that any central bank is able to predict, besides the inflation rate in one year's time, also its variance and adjusts the contemporaneous interest rate to it. To obtain consistent results of this estimation, it is necessary to ensure that the previous ARCH model has not been misspecified and the estimated conditional variance is not noisy. Misspecification is tested for by means of an LM test applied to the standardized residuals from the GARCH model, which must not be serially correlated.

where asymmetric preferences enter via the 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 tracks potential asymmetric preferences in terms of the deviation of the actual interest rate i_t from the estimated equilibrium value α.[5]

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 parameters describing policy preferences and the structure of the economy. However, the results are greatly conditioned given that the underlying relations are not observable and may 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, 2011) 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 aggregate demand (AD) schedule (Eq. (2)). The high economic openness of these countries again suggests that domestic output may have external determinants. In addition, the loss function of the monetary authorities (Eq. (3)) is not observable. Although all three countries officially apply IT aimed at price stability, other objectives are not discarded as long as they do not jeopardize price stability. Finally, as argued before, asymmetric monetary policy can arise due to a different response to positive versus negative shocks that cannot be tracked within the previous models.

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 for nonlinearities in the Fed policy rule using the flexible framework of Hamilton (2001), which takes into account uncertainty about the function forms. Cukierman and Muscatelli (2008) employ smooth transition regression to test for nonlinearities of the Taylor rules in the US and the UK. Florio (2006) augments their model with the possibility of nonlinearities in interest rate smoothing using the change in the 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 modify the policy stance in the face of information about (realized or expected) inflation or financial shocks rather than considering 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 the policy rule estimated in each regime (e.g., Bec et al., 2002):

[5] Ikeda (2010) proposes an interesting but also very specific extension of Surico's (2007a,b) model for small open economies. In particular, the exchange rate vis-à-vis the euro area is introduced into the policy loss function.

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

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

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 or below the target, which is the approach adopted for the Czech Republic by Horváth (2008), or whether output is above or below its potential (the threshold value is assumed to be zero).

In reality, the threshold value may not be known. For example, central bankers may 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 the threshold variable. Gredig (2007) estimates the threshold value of different variables (the inflation gap, the output gap, and gross domestic product (GDP) growth) for the Central Bank of Chile (CBC) and finds two different regimes according to the business cycle stance.[6] 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 statistical criteria to estimate consistently the threshold value (of a continuous variable) that splits the sample into two regimes. Although his method requires both the regressors and the threshold variable to be exogenous, Caner and Hansen (2004) suggested an extension to endogenous regressors.[7] We follow this framework, given that we are estimating a forward-looking policy rule from ex-post data. The model can be written as:

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

where the function *f* indicates whether the threshold variable q_i takes a 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.[8] The procedure is sequential. The first step consists of OLS estimation of the endogenous variables (in our case inflation and the output gaps) on a set of exogenous instruments:

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

[6] 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.

[7] Taylor and Davradakis (2006) employ GMM estimation (of a three-regime policy rule for the Bank of England) with a grid search for two threshold values (of the inflation rate) that minimize the GMM criterion function.

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

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

where z_t are the instruments – in our case the lagged values of the 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 $\hat{y}_t$, which 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 the criterion:

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

where S_n is the squared residual of Eq. (14) and ¤ 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):

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

Finally, we estimate by the GMM the monetary policy rule for sub-samples allowing for all the parameters switching between the two regimes. Unlike Caner and Hansen (2004), we again use the Newey and West (1994) heteroskedasticity 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 the inverted likelihood ratio statistics (more details below).

We use three threshold variables: (i) the 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 the 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) 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.

4. Data Description

Our dataset consists of monthly data ranging from 1998/M1 until 2010/M3.[9] The principal data source is the Main Economic Indicator database of the OECD and Eurostat.

The short-term interest rate is the three-month interbank interest rate for CZE and POL and the 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 perform the whole analysis using three measures of the inflation target (the inflation gap is always the deviation of expected inflation from the target value): (i) the actual inflation target of each central bank;[10] (ii) the smoothed (HP) trend of the inflation target;[11] and (iii) the smoothed (HP) trend of actual CPI inflation.[12] This seems to be a rather crucial issue in terms of robustness, given that the inflation targets in these NMS were time-varying with a downward-sloping trend, and there is no obvious argument for which of these three methods is superior. Figure A.1 compares the inflation gaps constructed by the three methods. The output gap is measured as the difference between the logarithm of the current value of seasonally adjusted GDP (in millions of euros at 1995 prices) and the trend value obtained by the Hodrick-Prescott (HP) filter (smoothing parameter set to 14,400). Given that GDP is available only quarterly, we have disaggregated it to monthly frequency using the univariate statistical method of Fernandez (1981), which allows the information to be augmented with the related series. For this purpose we have used the monthly industrial production index, which is arguably the most closely related series to GDP available at monthly frequency. Financial stress is measured by the EM-FSI elaborated by the International Monetary Fund (IMF) (Balakrishnan et al., 2009). It is a composite index of five subcomponents: (i) the 12-month rolling beta (from the capital asset pricing model – CAPM) of the bank stock index; (ii) stock market returns (the year-on-year change in the stock market index multiplied by

[9] We use monthly data to have a sufficient number of observations to apply the sample splitting techniques. Unfortunately, this comes at a cost. First, monthly data tend to be noisier than quarterly or yearly data. Second, some variables, such as inflation rates and interest rates, are highly persistent at monthly frequency. The persistence of the dependent variable in a model with partial adjustment drives the result that the coefficient of the 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 the interest rate to the inflation rate is very limited in the short term, while its long-term multiplier is very high. The results of the analysis can be confronted with the analysis with quarterly data contained in Vašíček (2010).

[10] 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 a band, whose width changes over time as well. Therefore, we always use 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.

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

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

minus one, so that a decline in stock prices implies an increase in the index); (iii) stock market volatility (six-month rolling monthly squared stock returns); (iv) the sovereign debt spread (the 10-year government bond yield minus the 10-year US Treasury bill yield); and (v) the exchange market pressure index (month-over-month percentage changes in the exchange rate and total reserves minus gold). The EM-FSI is constructed as the simple sum of the standardized subcomponents and is plotted for each country in Figure A.2. As reported by Balakrishnan et al. (2009), the EM-FSI captures most episodes of financial stress detected in previous studies.

5. Empirical Results

5.1 Linear Monetary Policy Rules

The GMM estimates of the linear monetary policy rules (Eq. (5)) are presented in Table 1. As noted above, given the fundamental uncertainty about what the best measure of the inflation gap is, 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 the inflation target; and (iii) the smoothed (HP) trend of 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 rather puzzling but may indicate that the intensity of the interest rate response to the inflation gap is not linear. In particular, a changing size of the response coefficient in different regimes may lie behind the elevated standard errors. Another possible interpretation is that the assumed forecasting horizon of 12 months might not fit all countries or periods of time. The significant response 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. Although we can see that the estimated degree of interest rate smoothing (coefficient ρ) is substantial, we must be aware that it can be interpreted in terms of true policy inertia only with a great deal of caution (see Rudebusch, 2006) since the interest rates, especially at monthly frequency, are autocorrelated by construction. The partial adjustment specification of the policy rule also determines that the elevated policy inertia affects the other coefficients, which may lie in principle behind some differences compared to the analysis conducted with quarterly data (Vašíček, 2010). In the standard Taylor rule the constant term (coefficient α) is usually interpreted as being the policy neutral rate. However, we must be careful in our case as we depart substantially from the standard framework (e.g., Clarida et al., 1998, 2000) using a time-varying inflation target. Finally, the benchmark results allow us to perform the first test of monetary policy rule nonlinearity. Following Siklos and Wohar (2005), who found evidence of ARCH-type effects in the monetary policy rule of the Fed, we apply the LM test for omitted ARCH to the residuals of each specification. The presence of ARCH is rejected at conventional significance levels for all specifications.

Table 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: The dependent variable is the short-term interest rate. The number of observations is 132. The three estimates provided for each country correspond to the three methods of proxying the inflation target: (i) the actual inflation target, (ii) the inflation target HP trend, and (iii) the inflation HP trend. HAC standard errors in parenthesis. *, **, and *** denote significance at 10, 5, and 1%. LB is the p-value of the Ljung-Box test for 1st order serial correlation. J-stat is the p-value of the Sargan overidentification test.

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 constraints) when the output gap is positive. Hence, as a first step we must test whether there is any evidence of a 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 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 the 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 the business cycle in these three NMS. This is also evident from a simple visual inspection of Figure 1, showing that the relation between inflation and the output gap is relatively weak. We have also departed from Dolado et al. (2005) by including other quadratic terms in Eq. (6), but we do not find them to be significant. Although the results may be affected by noise in measuring the output gap, there is also some evidence that inflation rates in the NMS have significant external determinants (Stavrev, 2009; Vašíček, 2011).

Table 2: OLS/GARCH Estimates of Simple Linear/Nonlinear Phillips Curves (Eq. (6))

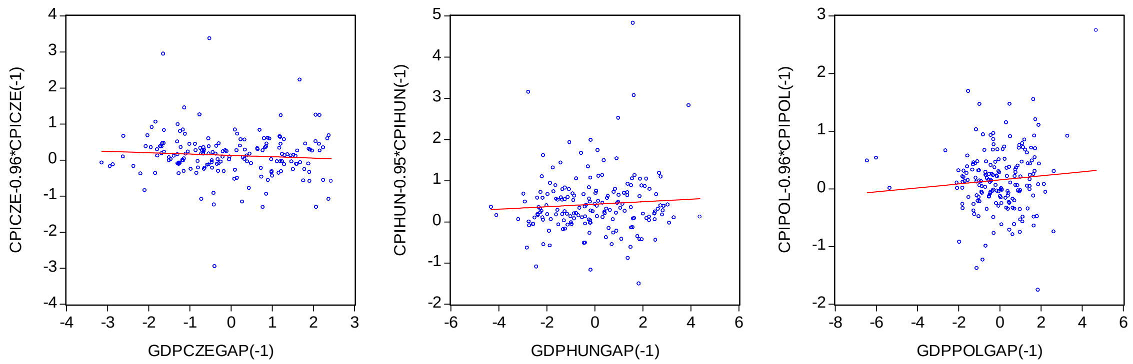
Country	α (const.)	β (π_{t-1})	γ (y_{t-1})	$\gamma\varphi$ (y_{t-1}^2)	ω (const.)	ν_1 (ξ_{t-1}^2)	ν_2 (σ_{t-1}^2)	R^2	LB
CZE (OLS.)	0.10	0.96***	0.06					0.95	0.00
	(0.07)	(0.01)	(0.04)						
CZE (GARCH)	0.16	0.94***	0.02		0.05***	-0.04**	0.87***	0.95	0.00
	(0.11)	(0.04)	(0.06)		(0.01)	(0.02)	(0.03)		
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: The dependent variable is the short-term interest rate. The number of observations is 132. The three estimates provided for each country correspond to the three methods of proxying the inflation target: (i) the actual inflation target, (ii) the inflation target HP trend, and (iii) the inflation HP trend. Standard errors in parenthesis. *, **, and *** denote significance at 10, 5, and 1%. LB is the p-value of the Ljung-Box test for 1st order serial correlation.

Although the previous results put into question the convexity of the AS schedule or any other nonlinear relationship between output and inflation, we continue to estimate Eq. (7), where the inflation-output interaction term appears as an additional regressor. These results are reported in Table 3. As expected, this term is mostly insignificant and there is no indication of an 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.[13]

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

Figure 1: Scatter Plots between the Inflation Rate and the Output Gap (the Phillips Curve)



Notes: The y axis depicts the smoothed inflation rate from the estimated linear PC $\left(\pi_t - \hat{\beta}\pi_{t-1}\right)$ and the x axis represents the lagged output gap (y_{t-1}) for the Czech Republic (left), Hungary (middle), and Poland (right). The linear trend is fitted so as to proxy the output gap slope coefficient (γ) in the estimated PC.

Table 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: The dependent variable is the short-term interest rate. The number of observations is 132. The three estimates provided for each country correspond to the three methods of proxying the inflation target: (i) the actual inflation target, (ii) the inflation target HP trend, and (iii) the inflation HP trend. HAC standard errors in parenthesis. *, **, and *** denote significance at 10, 5, and 1%. LB is the p-value of the Ljung-Box test for 1st order serial correlation. J-stat is the p-value of the Sargan overidentification test.

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 rules due to a higher weight assigned to positive deviations of expected inflation from the target. Dolado et al. (2004) suggested tracking such nonlinearity by the inclusion of the conditional inflation variance (Eq. (9)) in an otherwise linear policy rule. Therefore, first, we need to check whether the inflation volatility is truly time-varying so that it can be used as a regressor in Eq. (9). Inflation is again modeled by the simple backward-looking PC (Eq. (6)) and the ARCH LM test is used to check the neglected ARCH in the 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 for both the linear and quadratic speciation of the PC appear in Table 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 4.

The short-term interest rate responds significantly to the conditional inflation variance in the Czech Republic, which suggests that the Czech National Bank handled inflation in an asymmetric manner, and in particular that it weighted positive deviations from the target more heavily than negative ones. This result seems to be mainly driven by the period of disinflation of 1998-1999 where the conditional inflation variance achieved maximum values and was the most volatile.[14] These findings are consistent both with narrative (Jonáš and Mishkin, 2004) and previous empirical evidence (Horváth, 2008). On the contrary, the conditional inflation variance enters with a counter-intuitive negative sign for Hungary, which is probably related to the noisiness (the residuals of the PC for Hungary do not contain ARCH effects) and very low variance of this series (standard deviation of 0.02 as compared to 0.43 for the Czech Republic and 0.16 for Poland). In any case, the results are again conditioned by the PC specification that was used to derive the conditional inflation variance.

[14] When we exclude observations for years 1998 and 1999 from the regression, the corresponding coefficient is not significant anymore.

Table 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^2_t)	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: The dependent variable is the short-term interest rate. The number of observations is 132. The three estimates provided for each country correspond to the three methods of proxying the inflation target: (i) the actual inflation target, (ii) the inflation target HP trend, and (iii) the inflation HP trend. HAC standard errors in parenthesis. *, **, and *** denote significance at 10, 5, and 1%. LB is the p-value of the Ljung-Box test for 1st order serial correlation. J-stat is the p-value of the Sargan overidentification test.

An alternative way to test whether a monetary policy rule is nonlinear due to asymmetric preferences is suggested by Surico (2007a,b). His approach does not require estimation of the conditional inflation variance to test for an asymmetric response to inflation. In addition, it allows for testing of whether the central bank has asymmetric preferences with respect to the output gap and the interest rate gap, the latter being defined as the deviation of the current interest rate from its long-term equilibrium value. The asymmetric preferences enter the policy rule by square components for the inflation, output, and interest rate gaps (Eq. (10)). We adjust the nonlinear rule derived in Surico (2007a,b) to make it more plausible for the inflation-targeting NMS. In particular, we replace the response to contemporaneous inflation by a response to the 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 5. The columns with estimates of κ_1, κ_2, and κ_4 refer to nonlinearities related to asymmetric preferences for the inflation, output, and interest rate gaps, respectively, and κ_3 captures the response to nonlinearities in the economic structure. First, the only country where we find some evidence of an 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 counter-intuitive finding is in fact consistent with the evidence from Eq. (9), where we find a negative response to the conditional inflation variance. On the other hand, we do not confirm the previous finding that the Czech National Bank treated positive deviations of inflation from its target asymmetrically. Second, coefficient κ_2 of the squared output gap is insignificant for the three countries, so if their central banks considered the stance of the business cycle (see Tables 2–4), they did so 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 for the Czech Republic and Hungary reflects distaste for actual interest rates exceeding the equilibrium value. The negative value found for Poland may be a sign that the Polish National Bank was resistant to keeping interest rates too low. In fact, a preference for higher interest rates (negative κ_4) can also be an indication of a preference for price stability, while the opposite (positive κ_4) can also indicate a preference for avoiding contraction. While our evidence for the Czech Republic is consistent with Ikeda (2010), the results for Poland are just the opposite. The reasons are probably related to the fact that Ikeda does not take into account the time-varying inflation target and uses a different inflation targeting horizon as well as a different output gap proxy. As compared to the benchmark linear case (Eq. (5)), the interest rate smoothing has substantially decreased to more plausible levels (Rudebusch, 2002, 2006). 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 less plausible results for the ECB such as a negative and insignificant response to the inflation rate.[15] 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.

[15] Rather surprisingly, he interprets these results as evidence that the ECB follows a nonlinear policy rule.

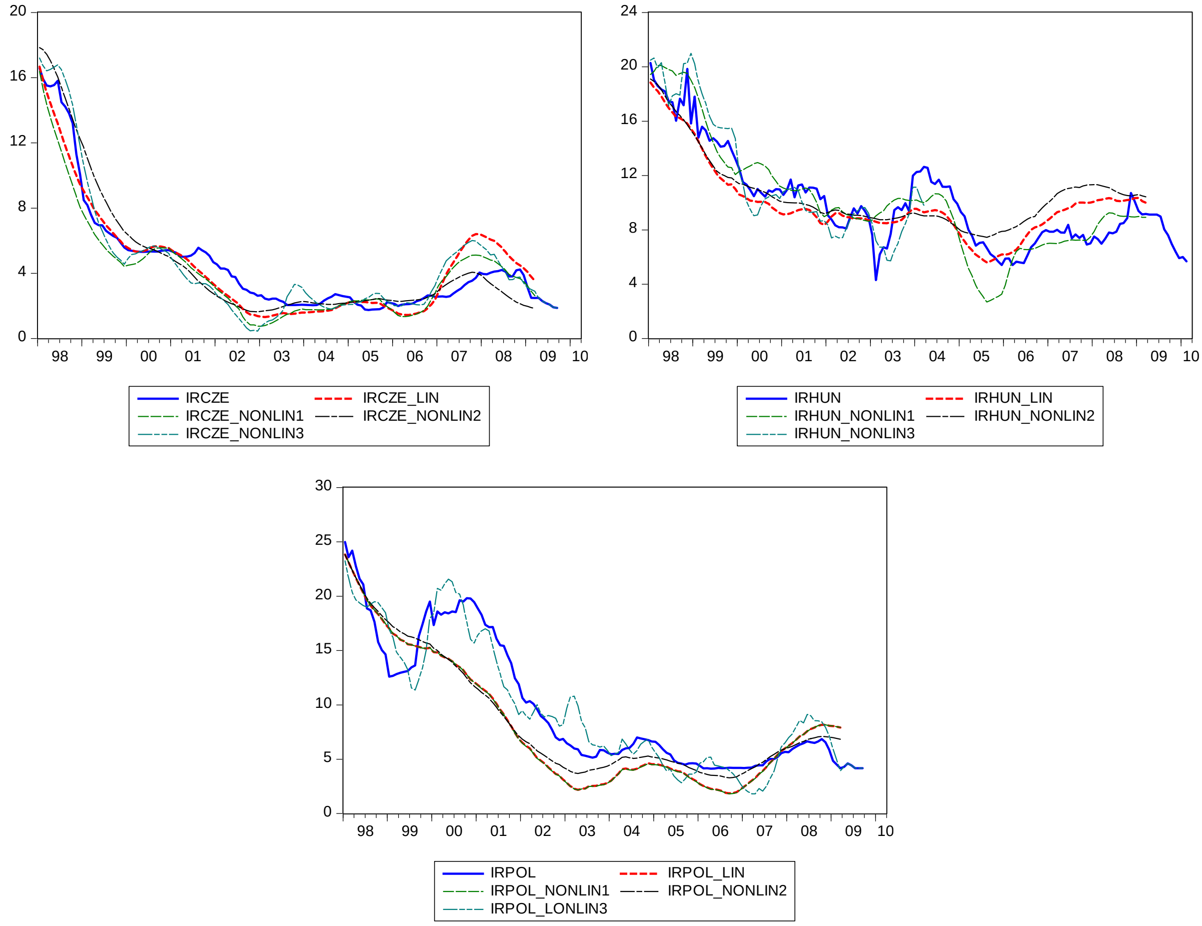
Table 5: GMM Estimates of the Nonlinear Monetary Policy Rule (Eq. (10))

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: The dependent variable is the short-term interest rate. The number of observations is 132. The three estimates provided for each country correspond to the three methods of proxying the inflation target: (i) the actual inflation target, (ii) the inflation target HP trend, and (iii) the inflation HP trend. HAC standard errors in parenthesis. *, **, and *** denote significance at 10, 5, and 1%.

So far we have been looking mainly at the statistical significance of the results. To shed more light on the economic significance of the results, it seems instructive to compare the relative performance of symmetric versus asymmetric descriptions of monetary policy. Figure 2 compares the short-term interest rate (the dependent variable of all specifications) with the in-sample forecast from the benchmark linear model (Eq. (5)) versus the three nonlinear alternatives tested above (Eqs. (7), (9), and (10)). The results clearly suggest that the in-sample forecast from at least one nonlinear model tracks the interest rate dynamics better than the one from the linear model, but the differences are not substantial most of the time. These findings are not surprising in light of the previous discussion and empirical estimates. In fact, a salient feature of policy asymmetry is that it pops up only in some delimited periods.

Figure 2: In-sample Forecast of Linear vs. Nonlinear Monetary Policy Rule



Notes: The inflation target is proxied in all specifications by the actual inflation target (see Figure A.1 for comparison of inflation target proxies). IR… is the short-term interest rate, IR…_LIN is the in-sample forecast from the linear policy rule (Eq. (5)), IR…_NONLIN1 is the in-sample forecast from the nonlinear policy rule of Dolado et al. (2005) (Eq. (7)), IR…_NONLIN2 is the in-sample forecast from the nonlinear policy rule of Dolado et al. (2004) (Eq. (9)), and IR…_NONLIN3 is the in-sample forecast from the nonlinear policy rule of Surico (2007a, b) (Eq. (10)).

5.4 Nonlinear Monetary Policy Rules via Threshold Effects

As argued earlier, the previous methods of inference on policy asymmetry rest on a 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 a threshold variable. Given that the threshold estimation method (Hansen, 2000; Caner and Hansen, 2004) requires the threshold to be exogenous, we use observed (rather than expected) values as the threshold.[16] Using inflation and the output gap as thresholds, we want to see whether the

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

interest rate setting differs in high and low inflation regimes and in recessions and expansions. Moreover, we include a new variable that can arguably give some insight into asymmetries in 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.

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 (the inflation-output interaction term, the conditional inflation variance or the squared terms of inflation, output, and interest rate gaps). With the current method, 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 the estimates inconsistent.[17] Given that the threshold estimation is based on 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 ¤ to evaluate the precision of the estimated threshold (see Figures A.3–A.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, an 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 no threshold effect at all.

In Figures A.3–A.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, the threshold effect of the inflation gap is not evident and depends on the measure of the inflation target. Although the LR sequences feature a usually well-defined minimum, it leads to a very asymmetric sample split, leaving one regime with only the minimum number of observations permitted (when the inflation rate very substantially exceeds the target value for the Czech Republic and Poland and when it is significantly below it for Hungary). This disqualifies the reasonability of 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 the deviation of inflation from its HP trend (the right-most figure; the

[17] The method splits the sample at the value of the threshold variable that minimizes the residuals of Eq. (14). When the splits imply that one regime contains only the minimum possible number of observations (10% of the total sample), while the other contains the remaining majority, it is an indication that that there is 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 for the presence of the threshold. However, given the uncertainty about the threshold variable, the threshold value as well as the number of policy regimes, we assess the presence of the threshold effect intuitively by graphical inspection of the LR statistics described below.

estimated threshold is 1.25), and to Poland, when using the deviation of inflation from the target HP trend (the middle figure; the 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.4 plots the respective LR sequences when the output gap is used as the threshold variable. We again discard 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 when we proxy inflation trend by means of the inflation trend value (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). 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. Given the ambiguity of these finding as well as in face of the previous results,[18] one cannot conclude that the Czech National Bank handled monetary policy in an asymmetric way over the business cycle. For Poland, we find a precise threshold in the first two models (with the 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 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 1) showing that the National Bank of Poland (NBP) responds to the output gap rather than to inflation. The results of the threshold model suggest that Polish monetary policy could be asymmetric over the business cycle. However, this evidence cannot be directly interpreted as meaning that the NBP, as a long-term inflation targeter, aims at business cycle stabilization instead of the inflation target. It might mean that the output gap affects the NBP's inflation forecast, which is the driver of interest rate setting.[19]

Finally, we use the financial stress indicator (EM-FSI). The evolution of this variable (normalized to have a zero mean) is depicted in Figure A.2. It is notable that during the recent global turmoil all three countries experienced a degree of financial stress 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). Unlike binary crisis variables (Leaven and Valencia, 2008), the EM-FSI allows the intensity of financial stress to be measured and can be used for threshold estimation. Nevertheless, it is not evident whether the EM-FSI should enter directly into the estimated policy rule as a regressor or "stay behind" as a threshold variable driving the regime switches. In other words, it is puzzling whether the central bank responds directly to some stress measure or only to modify its consideration of other objectives. Consequently, we estimate the threshold model with and without financial stress as an additional regressor. Figure A.5 depicts the evolution of the LR when the EM-FSI is included as a regressor, which is almost identical

[18] The linear monetary policy rules (Table 1) feature a substantially smaller and statistically insignificant response to the output gap. Similarly, the estimates of nonlinear policy rules in line with Surico (2007a,b), reported in Table 5, do not indicate statistical significance of the squared term for the output gap.

[19] The GARCH estimates of the Polish PC reported in Table 2 indicate that the output gap has a significant effect on the inflation rate.

with the EM-FSI dropped. We can see that the threshold value is clearly delimited in all three figures for the Czech Republic and the first two figures for Poland. For Hungary, the LR sequence reaches its minimum at very high values of 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 6.[20]

In all but one case, the upper regime has substantially fewer observations than the lower one. The threshold values file into the upper regime of high stress for an extended period during the period of the Russian crisis (starting in August 1998) and the global financial crisis of 2008 (from August 2008 onwards, peaking in late 2008) and for the Czech Republic also in the first half of 2001 due to stock market volatility and for Hungary in the first half of 2006 due to banking stress. The coefficient of the EM-FSI is mostly significant, suggesting that central bankers adjust policy rates when they are faced with financial stress. Since central bankers might respond to increasing financial stress by monetary easing, the expected sign of the coefficient is negative. Yet the EM-FSI also includes a sub-component representing the exchange rate pressures, in particular domestic currency depreciation,[21] whose prevalence in the overall index can drive an interest rate increase in an attempt to support the domestic currency. For the Czech Republic and Poland, we find that the coefficient κ accompanying the 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 overall Hungarian EM-FSI, which Hungarian monetary policy faced by means of interest rate increases.[22] Our results are slightly different from Ikeda (2010), who uses sub-sample analysis assuming that the crisis arrived in the CEECs as early as in January 2007, whereas most evidence as well as the EM-FSI suggest that the region did not become subject to financial stress until mid-2008. In general, he finds that the recent crisis did change the policy course in Poland, but not in the Czech Republic.

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 threshold effects.[23] 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

[20] We report results with the EM-FSI included as a regressor given that this specification has a better fit and the accompanying coefficient of the EM-FSI is mostly statistically significant.

[21] This subcomponent is not present in the financial stress index proposed by the IMF for advanced economies (Cardarelli et al., 2011).

[22] Baxa et al. (2011) study the response of the 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 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 the time series available, we cannot apply such a framework to the NMS.

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

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, which 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 IT because the response to the inflation gap is mostly insignificant. This is arguably related to the fact that during the period of analysis Hungary implemented an exchange rate band along with the inflation target. Therefore, interest rates might not be increased even though the inflation target is being jeopardized as long as the exchange rate is close to the lower fluctuation band. 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).

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

Table 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	0.99	0.16
	(0.63)	(0.55)	(0.28)	(0.01)	(0.01)				
	5.93***	0.31**	-0.89***	0.90***	-0.05***	> 1.12	32	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	0.99	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: The dependent variable is the short-term interest rate. The three estimates provided for each country correspond to the three methods of proxying the inflation target: (i) the actual inflation target, (ii) the inflation target HP trend, and (iii) the inflation HP trend. HAC standard errors in parenthesis. *, **, and *** denote significance at 10, 5, and 1%. Observ. stands for the number of observations in each regime. Q is the estimated value of the threshold. LB is the p-value of the Ljung-Box test for 1st order serial correlation.

6. Conclusions

Numerous empirical studies try to describe monetary policy decisions by means of estimated Taylor rules. There are different reasons why monetary policy can in fact be asymmetric, in the sense that the intensity of the central bank response varies according to economic developments. Our empirical analysis tries to reveal whether monetary policy could be described as asymmetric in three NMS that apply IT (the Czech Republic, Hungary and Poland). However, this study aims at providing some intuition about the sources of policy asymmetry rather than a specific test of whether policy is asymmetric or not. We find that the overall evidence is mixed. When we use GMM estimation of nonlinear policy rules derived from specific underlying models (Dolado et al., 2004, 2005; Surico, 2007a,b) we do not find any rationale 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 seemed to have weighted situations where inflation exceeded the target more heavily than those where it was below it during the initial period of the IT, while the opposite pattern was identified for Hungary. While the former finding is consistent with asymmetric policy handling during the disinflation period, the latter does not have any clear interpretation but can be possibly linked to inconsistencies between inflation and exchange rate targets pursued in Hungary. Interestingly, for all three countries we reveal a preference to limit the volatility of the current interest rate from its equilibrium value. 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, a preference for lower rather than higher interest rates can be an indication of a preference to avoid contractions, while the opposite points to a preference for price stability.

The previous results rely on the 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, as an alternative we use 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 a competing threshold variable. The threshold effects are most evident with the financial stress index. While the Czech and Polish central banks seem to face financial stress by decreasing their policy rates, the opposite pattern is found for Hungary.

The policy implications of our empirical findings can be summarized as follows. First, the mixed evidence on asymmetric responses to expected inflation should be confronted with the fact that the IT regime, as implemented by most central banks, legally de facto implies symmetric policy handling. The target is usually expressed by a point value or band with no recognition that positive deviations are less desirable than negative ones. However, the narratives suggest that policy handling can be often asymmetric when IT is used as disinflation strategy, which can be either related to the aim to approach faster the long-term inflation target or a significant uncertainty, which is related to the achievement of the mid-term targets. In this regard, it seems interesting to ask to what extent the asymmetric preferences in inflation revealed for the Czech National Bank contributed (along with shocks such as unexpected koruna appreciation) to the inflation target undershooting that often occurred during the first years of Czech inflation targeting.[24] Second, the indication that the degree of financial distress can alter the monetary policy stance has two possible interpretations; it is an indication either that financial stability is a goal which is pursued (even implicitly) by central banks (irrespective of its effect on future inflation and output), or that central bankers adjust policy rates being aware that financial instability can affect the future path of the macroeconomic variables it targets. The latter view can be linked to the financial accelerator literature (Bernanke et al., 1996) suggesting that financial vulnerabilities can amplify adverse shocks to the economy. In the aftermath of the recent crisis, the former approach seems to be favored over the latter. Still, there is a discussion whether monetary policy that pays attention to financial stability should be reactive or pre-emptive (see, for example, Borio and White, 2004, or Cúrdia and Woodford, 2010). Our study de facto tests reactive policy, as the EM-FSI is a coincident indicator of the stability of the financial system. Finally, the forecasting models employed in central banks commonly describe monetary policy in a symmetric form (for example, in the case of IT central banks only in terms of inflation). Therefore, the potential asymmetry of actual policy decisions could open a gap between the modeling apparatus of central banks and practical policy decisions.

There are different avenues of future research. First, it could be interesting to compare the behavior of central banks in the NMS and in other emerging countries that use IT 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 to include different aspects of small open economies to derive model-based nonlinear policy rules that are more suitable for the NMS. In this vein, Ikeda (2010) proposed a model where central bankers have preferences in terms of the exchange rate. Third, with respect to the threshold model, the assumption of an exogenous threshold variable can be too restrictive given the forward-looking nature of IT. Recently, Kourtellos et al. (2009) extended the model of Caner and Hansen (2004) to include 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 the possibility that monetary policy is asymmetric but also that it evolves over time.

[24] Šmídková (2008) and special Issue 58 (2008) of the *Czech Journal of Economics and Finance* deals in great detail with the first ten years of inflation targeting in the Czech Republic.

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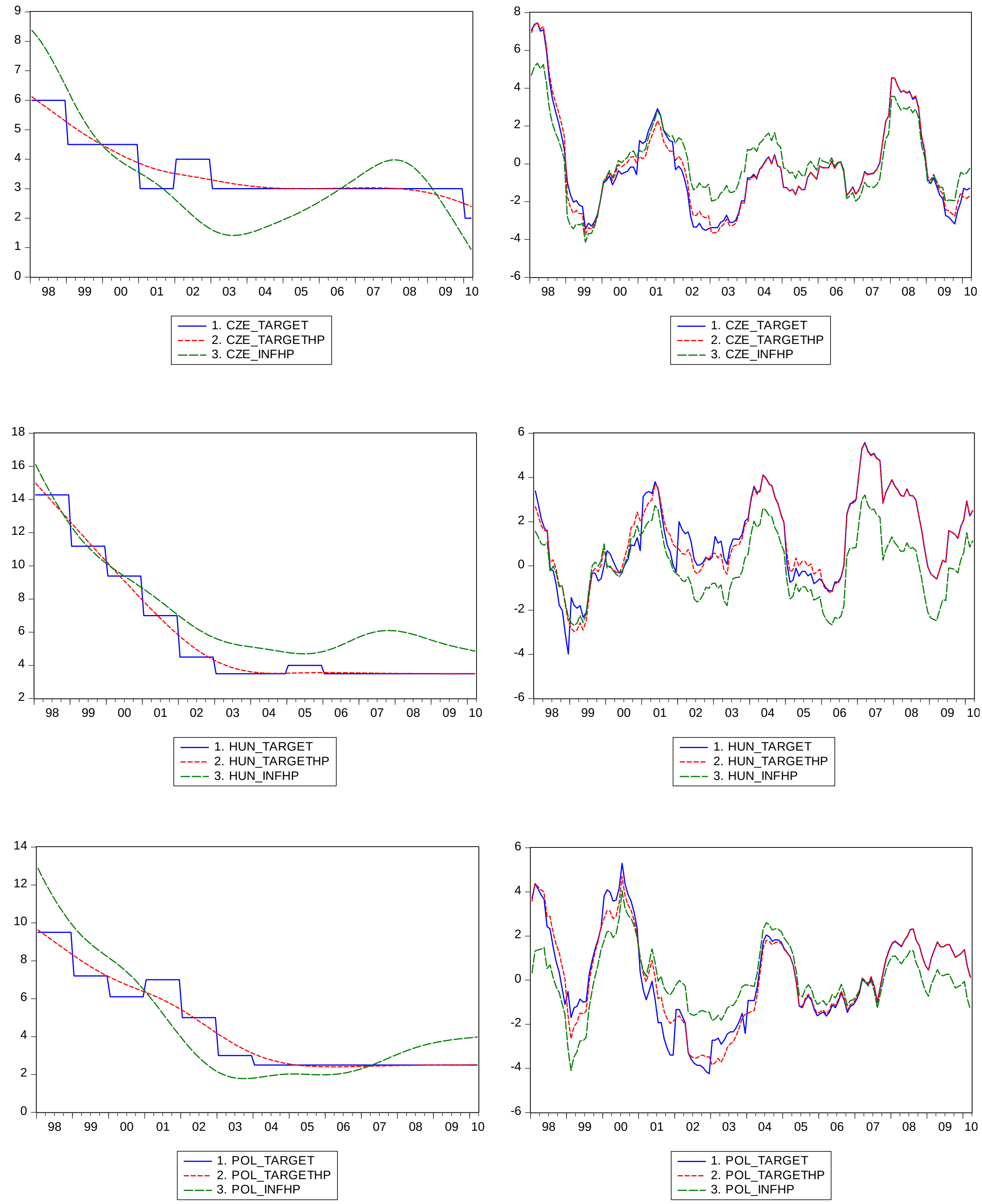
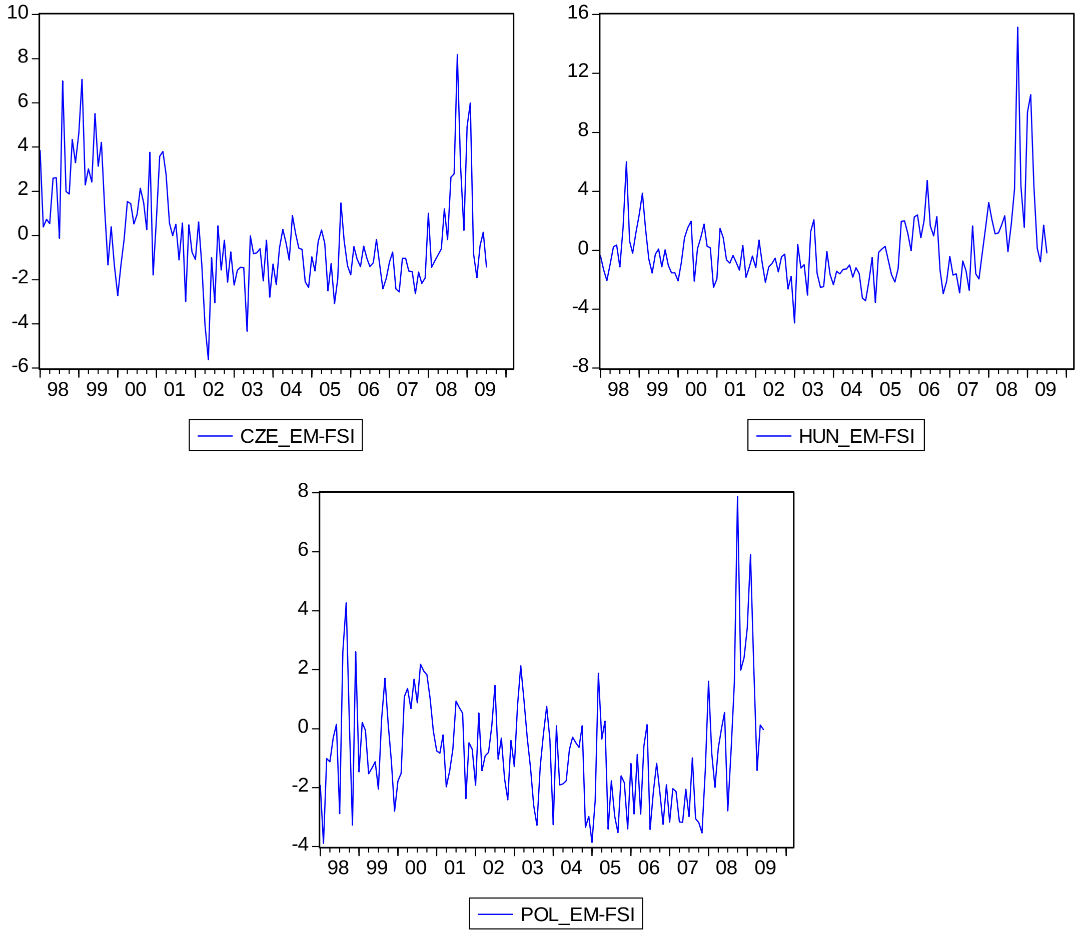
Appendix

Figure A.1: Proxies of the Inflation Target and the Inflation Gap



Notes: The inflation gap (right panel) in is defined as the difference between expected CPI inflation in *t* and the inflation target in *t* (left panel). Expected inflation is proxied by actual CPI inflation in *t*, and three methods are used to proxy the inflation target: 1. the official inflation target (the value is set for December of each year and is used for all the months of the same year), 2. the HP trend of the official inflation target, 3. the HP trend of CPI inflation.

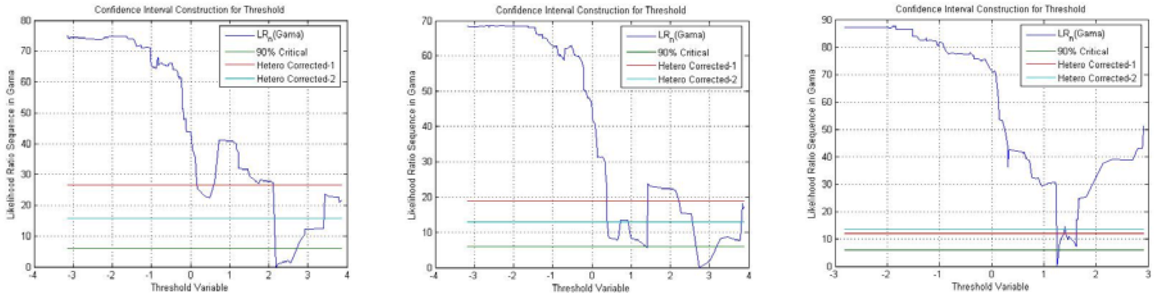
Figure A.2: The IMF's Emerging-Markets Financial Stress Index (EM-FSI)



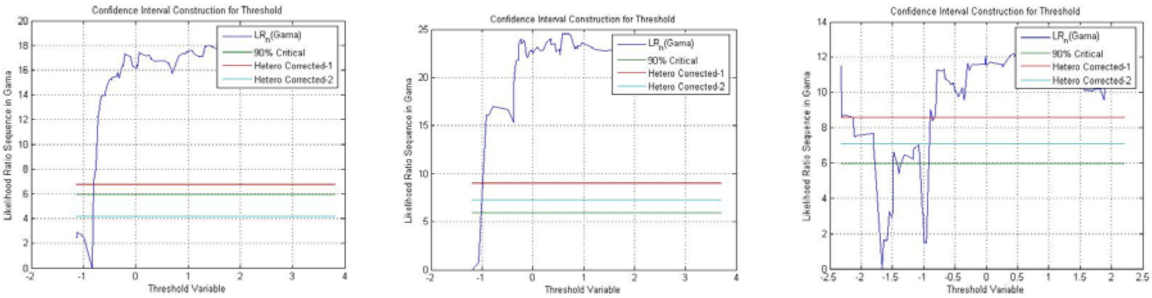
Notes: The EM-FSI is a simple sum of five subcomponents as defined in Balakrishnan et al. (2009).

Figure A.3: Likelihood Ratio Sequences for Different Values of the Threshold Variable (the Inflation Gap)

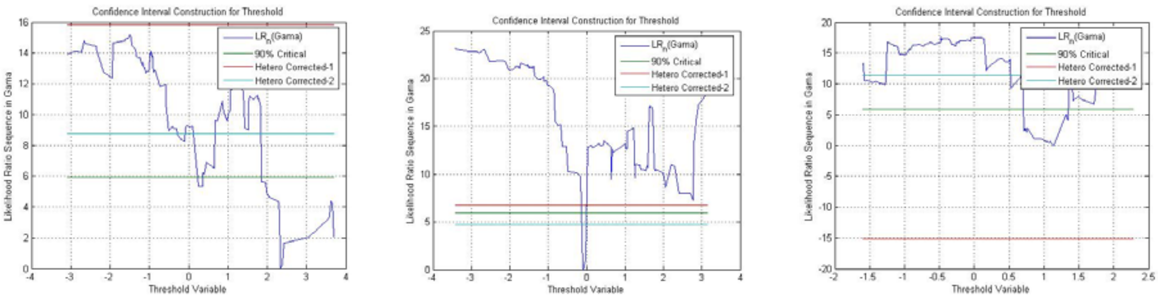
Czech Republic



Hungary



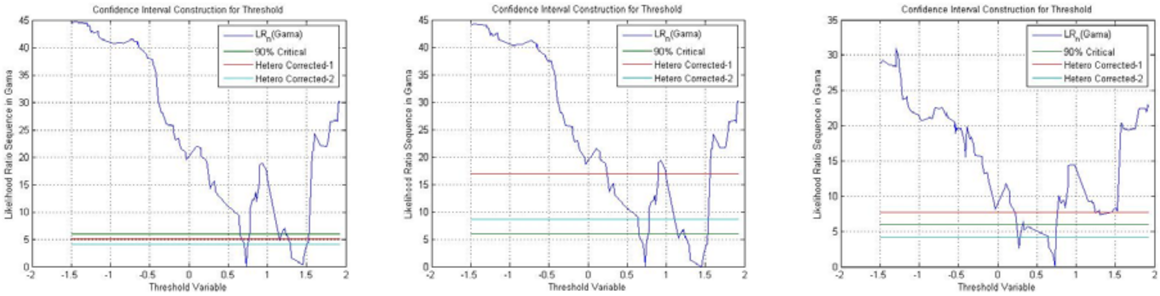
Poland



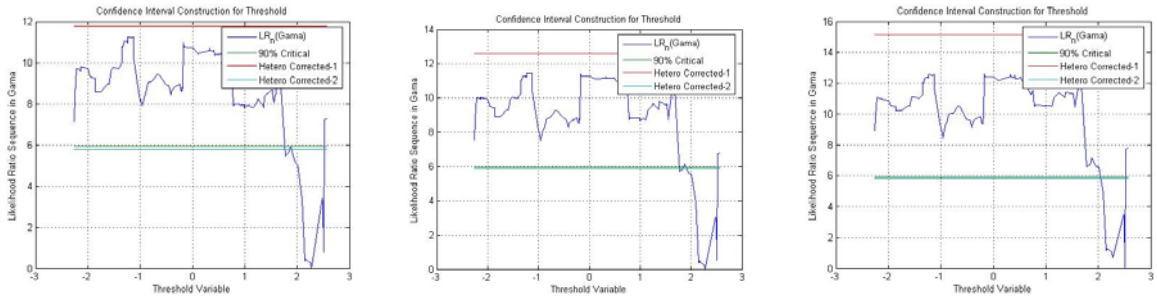
Notes: The likelihood ratio sequence represents the value of the inverted likelihood ratio statistics (Hansen, 2000) calculated for all values of the threshold variable (the inflation gap) according to Eq. (16). The green horizontal line depicts the 90% critical values and the red/blue line the 90% heteroskedasticity-corrected critical values. A particular value of the threshold variable belongs to the threshold interval when its inverted likelihood ratio is lower than the critical value. The three figures for each country correspond to the three different methods of proxying the inflation target.

Figure A.4: Likelihood Ratio Sequences for Different Values of the Threshold Variable (the Output Gap)

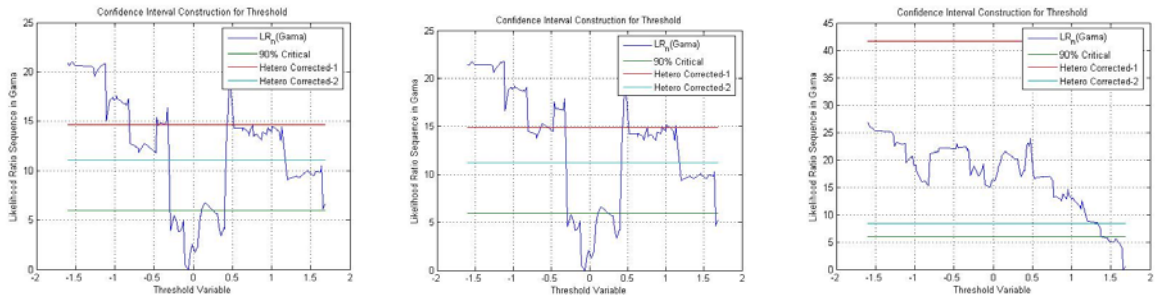
Czech Republic



Hungary



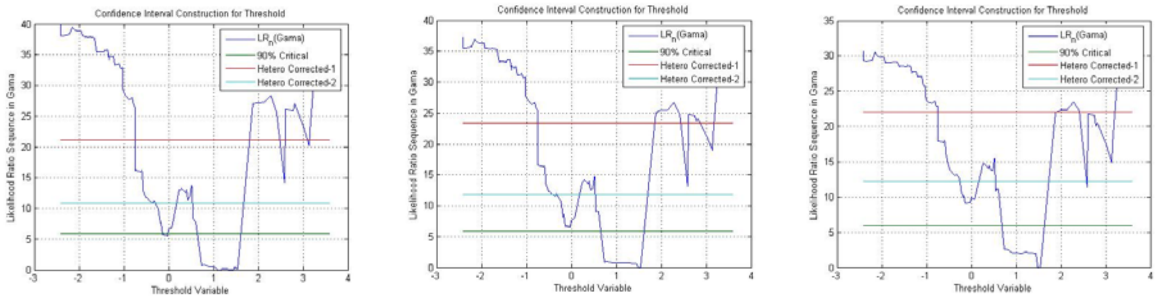
Poland



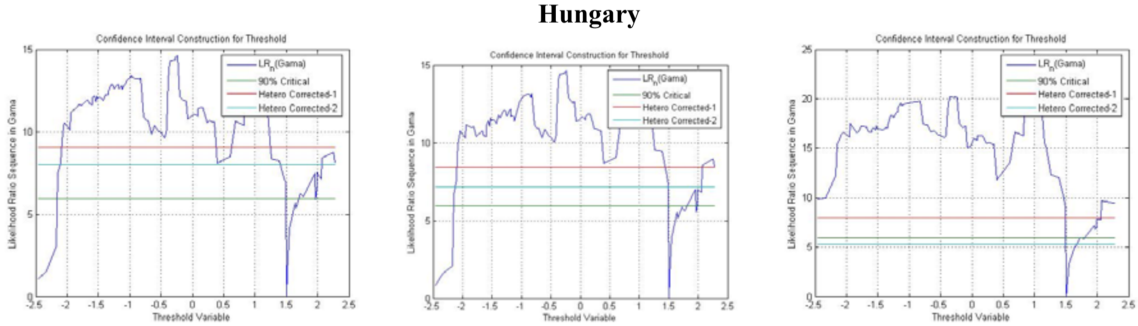
Notes: The likelihood ratio sequence represents the value of the inverted likelihood ratio statistics (Hansen, 2000) calculated for all values of the threshold variable (the output gap) according to Eq. (16). The green horizontal line depicts the 90% critical values and the red/blue line the 90% heteroskedasticity-corrected critical values. A particular value of the threshold variable belongs to the threshold interval when its inverted likelihood ratio is lower than the critical value. The three figures for each country correspond to the three different methods of proxying the inflation target.

Figure A.5: Likelihood Ratio Sequences for Different Values of the Threshold Variable (the EM-FSI)

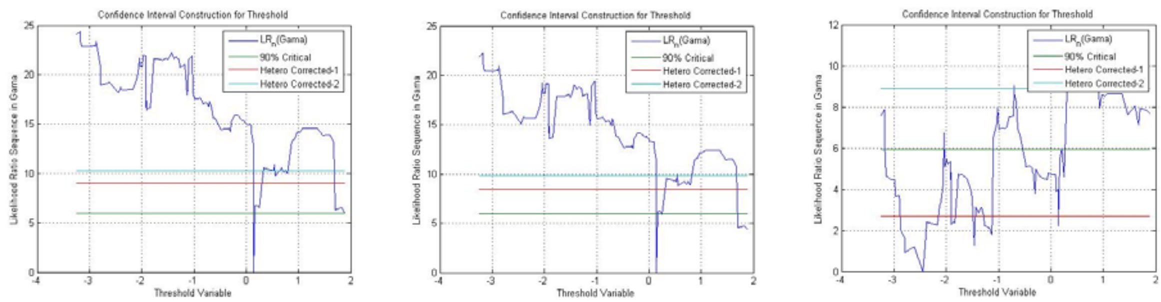
Czech Republic



Hungary



Poland



Notes: The likelihood ratio sequence represents the value of the inverted likelihood ratio statistics (Hansen, 2000) calculated for all values of the threshold variable (EM-FSI) according to Eq. (16). The green horizontal line depicts the 90% critical values and the red/blue line the 90% heteroskedasticity-corrected critical values. A particular value of the threshold variable belongs to the threshold interval when its inverted likelihood ratio is lower than the critical value. The three figures for each country correspond to the three different methods of proxying the inflation target.

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ISSN 1803-7070