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Abstract

Estimated Taylor rules became popular as a description of monetary policy conduct. There are numerous reasons why real monetary policy can be asymmetric and estimated Taylor rule 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) a Generalized Method of Moments (GMM) estimation of models that allow discrimination between the sources of potential policy asymmetry but are conditioned by specific underlying relations (Dolado et al., 2004, 2005; Surico, 2007a,b); and (ii) a flexible framework of sample splitting where nonlinearity enters via a threshold variable and monetary policy is allowed to switch between regimes (Hansen, 2000; Caner and Hansen, 2004). 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.

Keywords: monetary policy, inflation targeting, nonlinear Taylor rules, threshold estimation

JEL Classification: C32, E52, E58

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1. Introduction

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

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

Numerous empirical studies have provided evidence that monetary policy setting of many central banks can really be characterized as asymmetric. The asymmetric loss function was found to affect the decisions of the Bank of England (Taylor and Davradakis, 2006) and the

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

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

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

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

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

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

2. Rationales for asymmetric monetary policy

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

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

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

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

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

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

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

3. Empirical testing of asymmetric monetary policy

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

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

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

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

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

3.1 Nonlinearities in the economic system

Monetary policy asymmetry can be related to nonlinearities in the economic system. In particular, nominal price stickiness can cause a nonlinear trade-off between inflation and output. Dolado et al. (2005) derive nonlinear monetary policy rule when the PC is convex. They propose to augment the standard linear policy rule, such as Eq. (5), by an interaction

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

term of the expected inflation and the output gap given that any inflationary pressures driven by the output gap are larger if the PC is convex, which calls for an additional interest rate increase whenever the output gap is positive.

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

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

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

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

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

3.2 Asymmetric preferences of the central bank

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

This claim can be empirically tested as follows. First, if the conditional inflation variance is time varying, the residuals of the PC (Eq. (6)) shall contain autoregressive conditional heteroscedasticity (ARCH) effects. The null hypothesis of conditional homoskedasticity can be tested by means of an ARCH LM test. If the null is rejected, Eq. (6) can be estimated more

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

efficiently using an ARCH-type of model. We use the common GARCH (1,1) with the variance equation defined as:

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

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

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

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

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

$$\begin{aligned}i_{t}&=\rho i_{i-1}+(1-\rho)\alpha+\beta\left(\pi_{t+12}-\pi_{t+12}^{*}\right)+\gamma y_{t}\\&+\kappa_{1}\left(\pi_{t+12}-\pi_{t+12}^{*}\right)^{2}+\kappa_{2}y_{t}^{2}+\kappa_{3}\left(\pi_{t+12}-\pi_{t+12}^{*}\right)y_{t}+\kappa_{4}\left(i_{t}-\alpha\right)^{2}+\upsilon_{t}\end{aligned} \tag{10}$$

where the asymmetric preferences enter via squared terms of inflation and the output gap while the inflation-output interaction term controls, as in Dolado et al. (2005), for potential rule nonlinearity coming from nonlinearity in the PC. Moreover, the last term track potential asymmetric preferences in terms of the deviation of actual interest rate i_{t} from the estimated equilibrium value α.

3.3 Policy regimes with a threshold effect

The previous frameworks derive the nonlinear monetary policy rule assuming specific functional forms and parameterizations of Eqs. (1), (2) and (3). Such a model-based approach allows linking of the estimated coefficients of the policy rule to the parameters describing the

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

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

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

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

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

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

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

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

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

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

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

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

[5] Assenmacher-Wesche (2006) uses a Markov switching model for the US, the UK and Germany. She finds evidence in favor of low- and high-inflation regimes for all three countries.

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

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

$$
\begin{aligned}
\left(\pi_{t+12}-\pi^{*}_{t+12}\right) &= \prod\nolimits_{1} z_t+\zeta_{1,t} \\
y_t &= \prod\nolimits_{2} z_t+\zeta_{2,t}
\end{aligned}
\tag{13}
$$

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

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

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

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

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):

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

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

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

4. Data description

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

The short-term interest rate is the three-month interbank interest rate for CZE and POL and overnight interbank interest rate for HUN given that the former is not available for the whole period of analysis. The inflation rate is measured by year-on-year changes in the consumer price index (CPI). We assume a forecasting horizon of 12 months and use three measures of the inflation target (inflation gap is always a deviation of expected inflation from the target value): (i) the actual inflation target of each central bank;[9] (ii) the smoothed (HP) trend of inflation target;[10] and (iii) the smoothed (HP) trend of actual CPI inflation.[11] Figure A.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 the seasonally adjusted GDP (in millions of euros in 1995 prices) and the trend value obtained by Hodrick-Prescott (HP) filter (the smoothing parameter set to 14400). Given that the GDP is available only quarterly, we have disaggregated it to monthly frequency using a univariate statistical method of Fernandez (1981) that allows the information to be augmented with the related series. For this purpose

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

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

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

[11] It can be argued that central banks aim rather at eliminating inflation that is significantly above its trend. This seems plausible for the NMS given that inflation targeting was introduced when inflation rates were still relatively high. To anchor the inflation expectation, the central banks had to stick to targets that were lower to what monetary policy could immediately achieve. However, they indicated the intention of monetary authorities to stabilize the price level.

we have used a monthly industrial production index, which is arguably the most related series to GDP available on monthly frequency. The financial stress is measured by EM-FSI elaborated by the International Monetary Fund (IMF) (Balakrishnan et al., 2009). It is a composite index of five subcomponents: (i) 12-months rolling beta (from the capital asset pricing model – CAPM) of bank stock index; (ii) stock market returns (year-on-year change in stock market index multiplied by minus one, so that declines in stock prices implies index increase); (iii) stock market volatility (six-month rolling monthly squared stock returns); (iv) sovereign debt spread (10-year government bond yield minus 10-year US Treasury bill yield); and (v) exchange market pressure index (month-over-month percent changes in the exchange rate and total reserves minus gold). The EM-FSI is constructed as a simple sum of standardized subcomponents and is plotted for each country in Figure A.2.

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 a fundamental uncertainty of what is the best measure of inflation gap, we report for each country the results with inflation gaps derived from the three alternative measures of the inflation target: (i) the actual inflation target of each central bank; (ii) the smoothed (HP) trend of inflation target; and (iii) the smoothed (HP) trend of the CPI inflation. We can see that most of the coefficients have the expected sign. The expected inflation gap (coefficient β) enters significantly in the Czech Republic but not in Hungary and Poland (due to elevated standard errors). This finding is a bit puzzling but it can be an indication that the intensity of the interest rate response to the inflation gap is not linear. The significant response to the output gap (coefficient γ) found in Poland can be interpreted as a policy aimed at price stability as long as the output gap predicts future inflation pressures. At last, we can see that the degree of interest rate smoothing (coefficient ρ) is substantial. Nevertheless, a high smoothing coefficient can be interpreted in terms of true policy inertia only with a lot of caution (see Rudebusch, 2006) since the interest rates at monthly frequency are autocorrelated by construction.

Table 1.: GMM estimates of the linear monetary policy rule (Eq. (5))

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

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

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

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

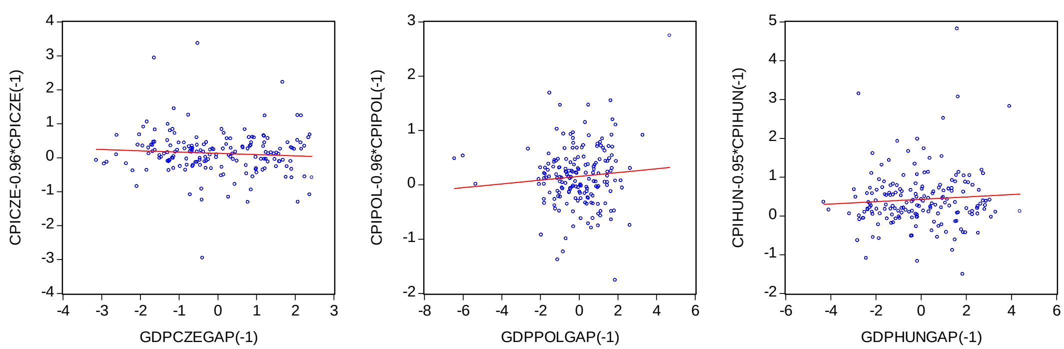
Estimates of the linear and nonlinear version of the simple backward-looking PC (Eq. (6)) appear in Table 2. Besides OLS we also use a GARCH(1,1) model to take into account the potential time-varying volatility of inflation. We are mainly interested in sign and statistical significance of the coefficient of the squared output gap $\gamma\varphi$. The PC is convex when this term is positive. The results show that there is little evidence of any (linear or nonlinear) relationship between inflation and the stance of business cycle in these three NMS. This is also evident by simple visual inspection of Figure 1, showing the scatter plots between the smoothed inflation rate $\left(\pi_t - \hat{\beta}\pi_{t-1}\right)$ and the output gap y_{t-1}. Although the results can be affected by noise in measuring the output gap, there is also some evidence showing that inflation rates in the NMS have significant external determinants (Stavrev, 2009; Vašíček, 2010b).

Table 2.: OLS/GARCH estimates of simple linear/non-linear Phillips curves (Eq. (6))

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

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

Figure 1: Scatter plots between smoothed inflation rate and output gap (the Phillips curve) and fitted linear trend



Although the previous results put into question the convexity of the AS schedule, we continue to estimate Eq. (7), where the inflation-output interaction term appears as an additional regressor. These results are reported in Table 3. As expected, this term is mostly insignificant and there is no indication of asymmetric central bank reaction driven by a nonlinear PC. In any case, it is important to keep in mind that the results are conditioned by the underlying model.[12]

Table 3.: GMM estimates of the nonlinear monetary policy rule (Eq. (7))

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

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

5.3. Nonlinear monetary policy rules due to asymmetric preferences

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

First, we explore whether the central banks of the three NMS applied nonlinear policy rule due to higher weight assigned to positive deviation of expected inflation from the target. Dolado et al. (2004) suggested tracking such nonlinearity by the inclusion of conditional inflation variance (Eq. (9)) to an otherwise linear policy rule. Therefore, first, we need to

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

check whether the inflation volatility is truly time-varying to be used as an regressor in Eq. (9) . The inflation is again modeled by simple backward-looking PC (Eq. (6)) and the ARCH LM test is used to check the neglected ARCH in residuals. The test gives affirmative evidence for the Czech Republic and Poland but cannot reject the null of no conditional heteroskedasticity for Hungary. Conditioned on these results, we re-estimate the PC using GARCH (1,1). The results of the corresponding mean and variance equation both for linear and quadratic speciation of the PC appear in Table 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.

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

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

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

An alternative way to test whether monetary policy rule is nonlinear due to asymmetric preferences is suggested by Surico (2007a,b). His approach does not require estimation of conditional inflation variance to test asymmetric response to inflation. In addition, it allows the testing of whether the central bank has asymmetric preferences with respect to the output gap and the interest rate gap, the latter is defined as a deviation of the current interest rate from its long-term equilibrium value. The asymmetric preferences enter the policy rule by square components for inflation, output and interest rate gaps (Eq. (10)). We adjust the nonlinear rule derived in Surico (2007a,b) to be more plausible for the inflation-targeting NMS. In particular, we replace the response to contemporaneous inflation by response to expected inflation gap given that inflation-targeting central banks are forward-looking and the inflation target is not constant. The estimates of such nonlinear policy rule appear in Table 5. The columns with estimates of κ_1, κ_2 and κ_4 refer to nonlinearities related to asymmetric preferences for inflation, output and interest rate gaps, respectively, and κ_3 captures a response to nonlinearities in economic structure. First, the only country where we find some evidence of asymmetric response to the inflation gap is Hungary, though the sign of coefficient κ_1 is negative, implying a stronger response when inflation is below its target. This contra-intuitive finding is in fact consistent with the evidence from Eq. (9) where we find a negative response to conditional inflation variance. On the other hand, we do not confirm the previous finding that the Czech National Bank treated the positive inflation deviation asymmetrically from its target. Second, the coefficient κ_2 of the squared output gap is insignificant for the three countries and if their central banks considered the stance of business cycle (see Tables 2–4), they did it in a symmetric manner. Third, for all countries we reveal a preference to limit the volatility of the current interest rate from its equilibrium value (proxied by the intercept α). The positive value of κ_4, found in the Czech Republic and Hungary, reflects a distaste for actual interest rates exceeding the equilibrium value. The negative value found for Poland can be a sign that the Polish National Bank was resistant to keeping the interest rates too low. In fact, the preference for higher interest rates (κ_4 negative) can also be an indication of a preference for price stability, while the opposite (κ_4 positive) can also indicate a preference to avoid contraction. As compared to the benchmark linear case (Eq. (5)), the interest rate smoothing has substantially decreased to more plausible levels (Rudebusch, 2002, 2006) as compared with the linear case (Eq. (5)). Finally, the inflation response coefficient β is not altered for the Czech Republic and Poland but it turns significant and higher than unity for Hungary, indicating a stabilizing nature of monetary policy conduct when the nonlinear nature of monetary policy is taken into account. These findings are promising as compared to Surico (2007a), who obtains for the ECB less plausible results such as a negative and insignificant response to inflation rate.[13] Due to reasons of space, we do not report the autocorrelation and over-identification tests but they provide a very similar picture as in previous tables.

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

5.4. Nonlinear monetary policy rules via threshold effects

As argued before, the previous methods of inference on policy asymmetry rests on the specific assumption about the structure of the economy and the central bank's loss function. In what follows, we use the empirical forward-looking policy rule proposed by Clarida et al. (1998, 2000) and allow the response coefficients to switch between two regimes according to the evolution of threshold variable. Given that the method of threshold estimation (Hansen, 2000; Caner and Hansen, 2004) requires the threshold to be exogenous, we use as a threshold observed (rather than expected) values.[14] Using inflation and the output gap as thresholds, we

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

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

want to see whether the interest rate setting differs in high and low inflation regimes and in recession and expansions. Moreover, we include a new variable that can arguably give some insight on asymmetries in a monetary policy setting, the financial stress index (EM-FSI). In this case, we try to uncover whether central banks alter their consideration of common policy targets in the face of financial instability and whether they directly adjust policy rates according to the degree of financial stress in the economy.

The inference on monetary policy asymmetry has so far been carried out by means of conventional t-tests of statistical significance of additional nonlinear terms (inflation – output interaction term, conditional inflation variance or squared terms of inflation, output and interest rate gaps). With the current method, the policy asymmetry is tested by means of threshold effects. Unfortunately, a standard Wald test comparing the point estimates in each regime cannot be used because the method provides a sample split even in the absence of true threshold effects, which makes estimates inconsistent.[15] Given that the threshold estimation is based on the minimization of the squared residual of Eq. (14), we can draw the inverted LR statistics (Eq.(16)) for the entire set of possible threshold values ¤ 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, irregular shape where $\mathrm{LR}_n(Q)$ crosses the confidence interval more than once and the minimum is less evident, is an indication that the sample may be split more than once or that there is not threshold effect at all.

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

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

Figure A.4 plots the respective LR sequences when the output gap is used as the threshold variable. We discard again the threshold model for Hungary as the LR reaches its minimum only at very high values of the inflation gap, making the sample split unfeasible. For the Czech Republic, we find a well-defined threshold only with the inflation deviation from the trend (the right-most panel). In this model, when the output gap exceeds the threshold value (estimated at 0.73), its coefficient γ is 2.43 versus 1.64 when it is below the target (in both cases this is highly significant). This suggests that the Czech National Bank handles monetary policy in an asymmetric way along the business cycle. In particular, it is ready to increase the interest rate by a larger amount during periods of economic expansion.[16] In the first two panels, we can see that the LR crosses the horizontal line more than once. However, the sample size does not allow another split. For Poland, we find a precise threshold in the first two models (with an inflation gap derived from the actual inflation target and from the HP trend of the target). The threshold value is estimated at -0.05 in both cases. While the corresponding response coefficient γ is insignificant in the regime below the threshold (i.e.

[16] These findings must be interpreted with caution given that the linear monetary policy rules (Table 1) features a substantially smaller and statistically insignificant response to the output gap. Similarly, the estimates of nonlinear policy rules in line with Surico (2007 a,b) reported in Table 5, do not indicate statistical significance of the squared term for the output gap.

when the output is below its potential), it turns significant and reaches a value of 14 when the threshold is breached. This finding is interesting in view of the linear model estimates (Table 1.) showing that the National Bank of Poland (NBP) responds rather to the output gap than inflation. The results of the threshold model suggest that Polish monetary policy is asymmetric along the business cycle. However, this evidence cannot be directly interpreted that the NBP, as a long-term inflation targeter, aims at business cycle stabilization instead of the inflation target. It might mean that the output gap affects the NBP's inflation forecast that is the driver of interest rate setting.[17]

At last, we use the financial stress indicator (EM-FSI). The evolution of this variable (normalized to have zero mean) is depicted in Figure A.2. It is notable that all three countries experienced a degree of financial stress during the recent global turmoil unseen in the previous decade but that the stress was also high as a consequence of the Russian crises in late 1998. On the other hand, unlike many developed countries, the NMS did not suffer an increase in financial stress on the eve of the new millennium following the NASDAQ crash (2000), the terrorist attack on the US (2001) or the US corporate scandals (2002). EM-FSI allows, unlike binary crisis variables (Leaven and Valencia, 2008), the intensity of financial stress to be measured and can be used for the threshold estimation. Nevertheless, it is not evident whether EM-FSI should enter directly in the estimated policy rule as regressor or "stay behind" as threshold variable driving the regime switches. In other words, it is puzzling whether the central bank may directly respond to some stress measure or only to modify its consideration of other objectives. Consequently, we estimate the threshold model with and without the financial stress as an additional regressor. Figure A.5 depicts the LR evolution when EM-FSI is included as a regressor, which is almost identical with EM-FSI dropped. We can see that the threshold value is clearly delimitated in all three figures for the Czech Republic and first two figures for Poland. For Hungary, the LR sequence reaches its minimum at very high values of the stress but there are still 28 observations in the upper regime. We split all the samples and pursue GMM estimation for each regime. These results are reported in Table 6.[18]

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

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

In all but one case, the upper regime has substantially less observations than the lower one. The coefficient of EM-FSI is mostly significant suggesting that central bankers adjust policy rates when they are faced by financial stress. Since the central bankers might respond to increasing financial stress by monetary easing, the expected sign of the coefficient is negative. Yet, EM-FSI also includes a sub-component representing the exchange rate pressures, in particular domestic currency depreciation,[19] whose prevalence in the overall index can drive an interest rate increase in an attempt to support domestic currency. For the Czech Republic and Poland, we find that the coefficient κ, accompanying EM-FSI is mostly negative and significant when the financial stress exceeds the estimated threshold. This suggests that both central banks decrease policy rates when the economy suffers high financial stress. On the contrary, the response is mostly insignificant when the stress falls below the threshold value. Hungary seems to be the opposite case; the interest rate response to financial stress is significantly positive and does not differ substantially between the two regimes. This could be related to the forint depreciation pressures that were a significant driver of the Hungarian overall EM-FSI and Hungarian monetary policy faced them by means of interest rate increase.[20]

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

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

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

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

because the response to the inflation gap is mostly insignificant. The coefficient of the output gap γ suggests that the real economy raises concerns only when the inflation stress is low (the Czech Republic and Poland) if at all (Hungary).

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

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

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

6. Conclusions

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

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

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

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Appendix

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

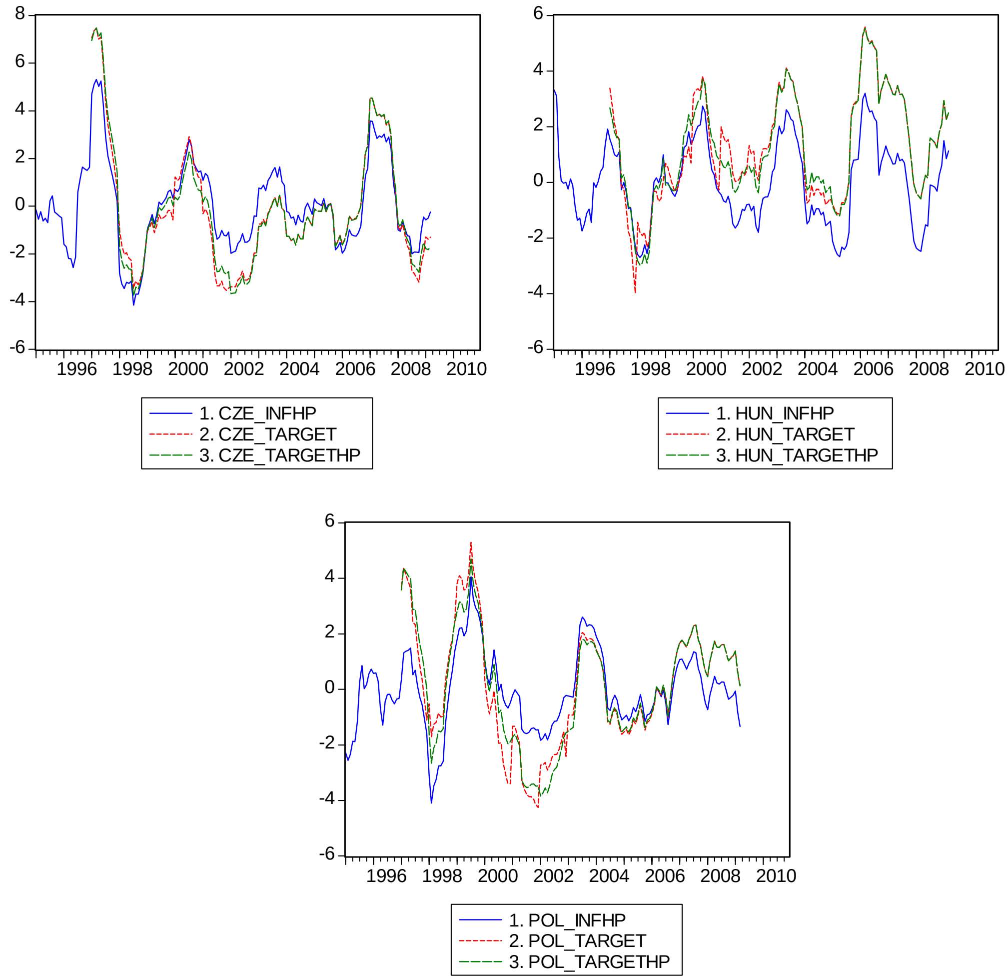


Figure A.2: IMF's Emerging-Markets Financial Stress Index (EM-FSI)

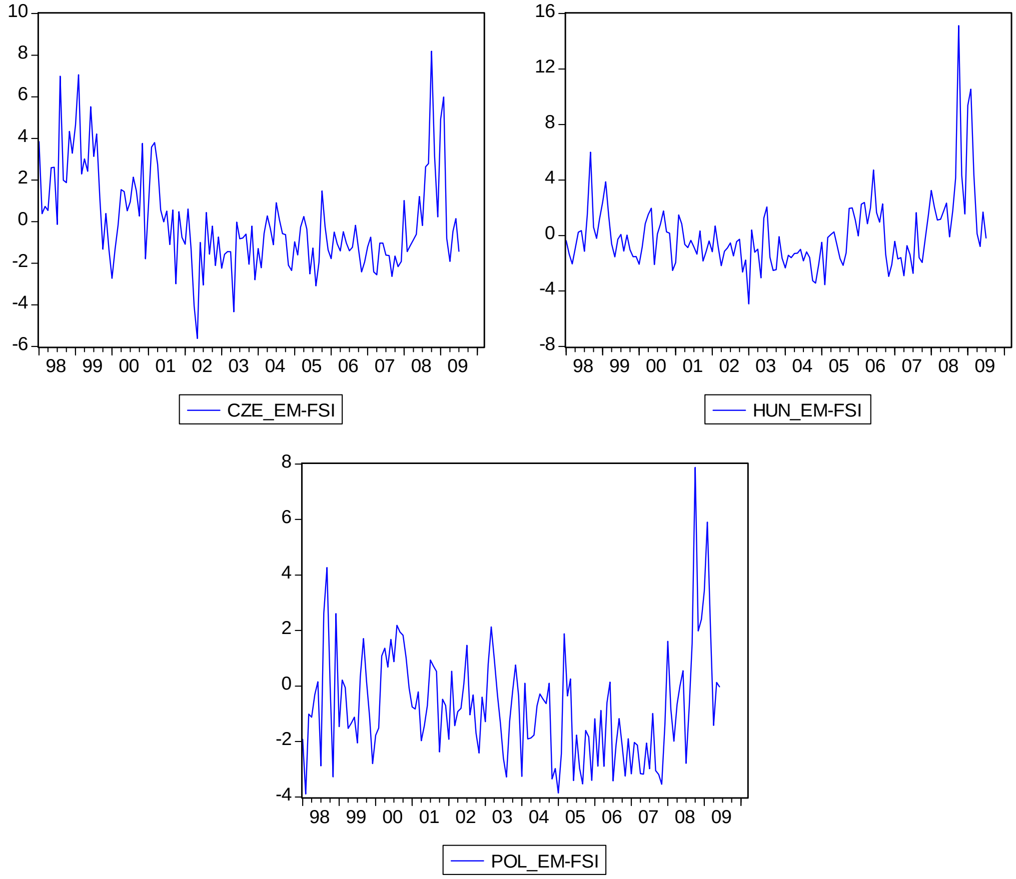
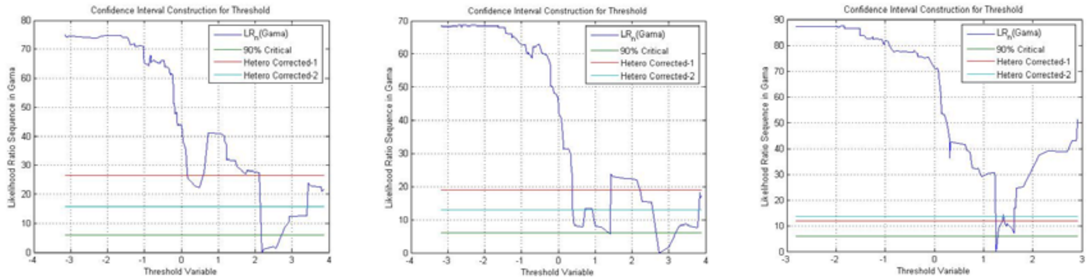
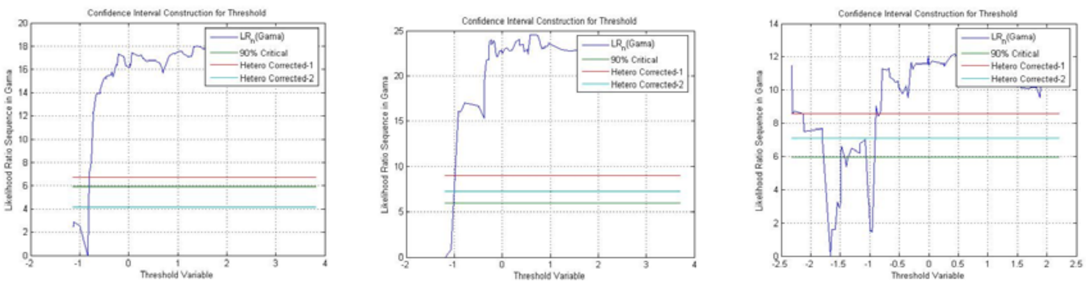


Figure A.3: Likelihood ratio sequence for different values of the threshold variable (the inflation gap)

The Czech Republic



Hungary



Poland

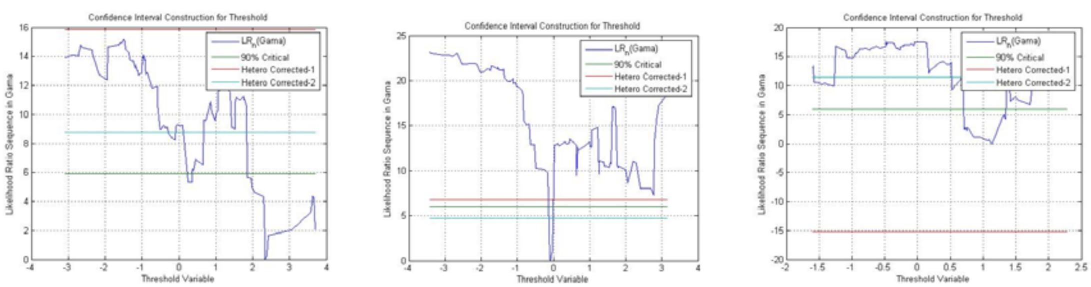
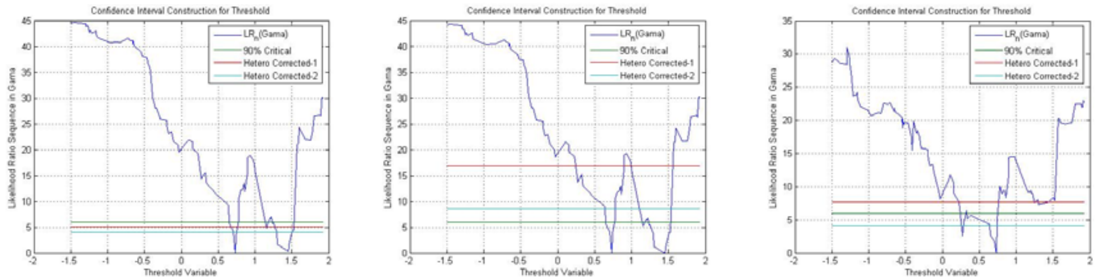
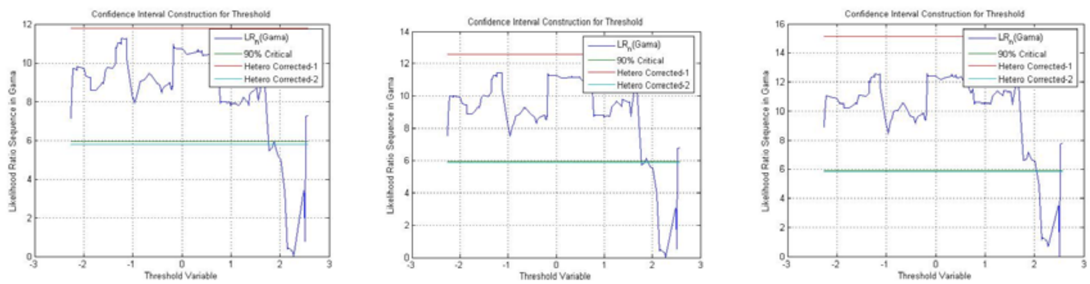


Figure A.4: Likelihood ratio sequence for different values of the threshold variable (the output gap)

The Czech Republic



Hungary



Poland

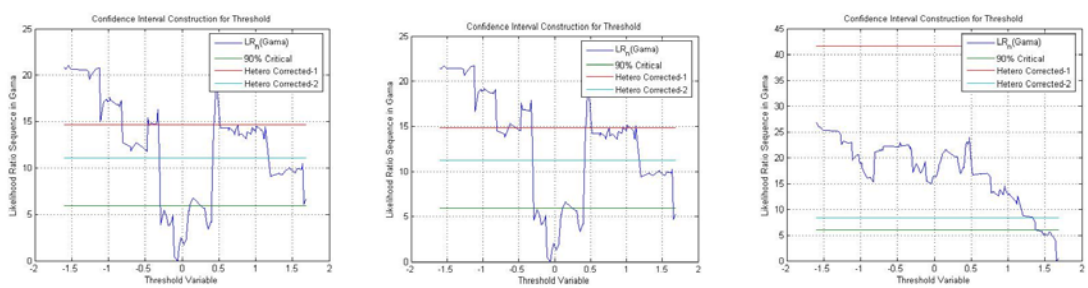
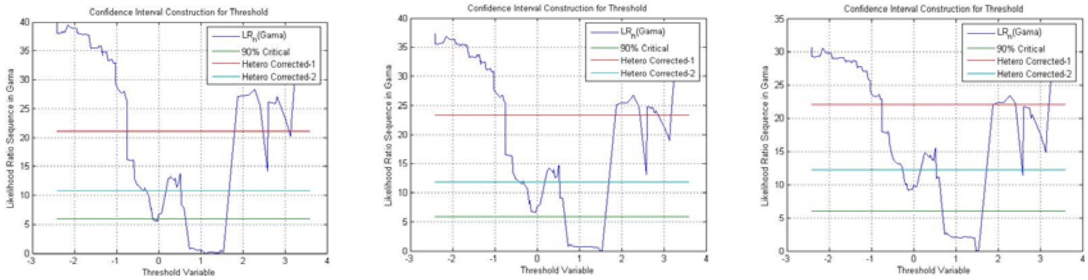
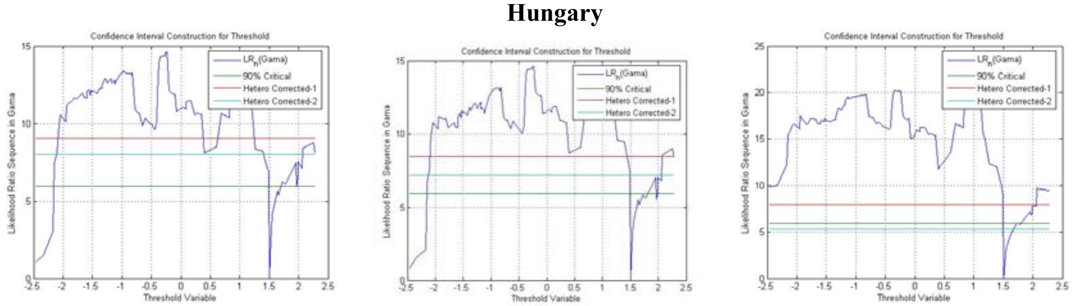


Figure A.5: Likelihood ratio sequence for different values of the threshold variable (the EM-FSI)

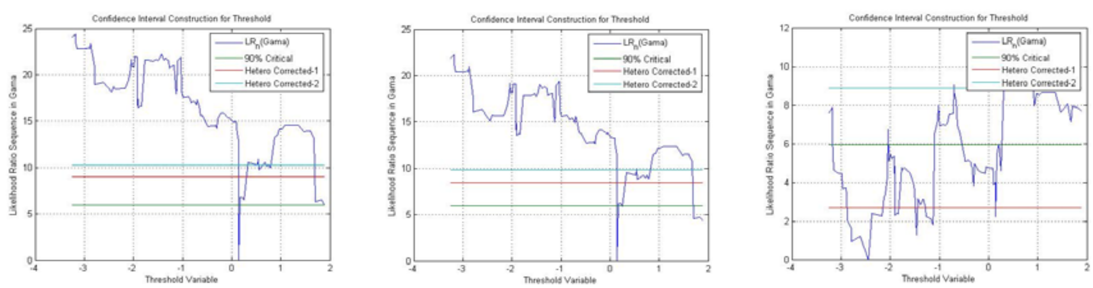
The Czech Republic



Hungary



Poland



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