Modeling the evolution of monetary policy rules in CESEE

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In this paper we use a novel econometric approach to estimate time-varying monetary policy rules for four inflation-targeting economies in Central, Eastern and Southeastern Europe (CESEE). Our results indicate that monetary policy in the Czech Republic, Hungary, Poland and Romania is strongly anchored to inflation stabilization, which implies that these economies follow a comparatively strict version of inflation targeting. By contrast, there is less evidence for output stabilization playing an important role in the conduct of monetary policy. Other factors that are of relevance in the monetary policy reaction function include the short-term interest rate in the euro area and – depending on the country under consideration – a measure of exchange rate movements. We find that the coefficients on domestic inflation expectations and euro area interest rates have declined since the mid-2000s, but that they still play an important role in central banks' reaction functions. This decline in the size of estimated coefficients may mirror an international environment characterized by loose and unconventional monetary policies that cannot be appropriately captured by euro area interest rates. It may also reflect contained global and domestic price growth.

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Since Taylor's (1993) seminal work on U.S. monetary policy, there has been a vastly growing literature that models central banks' reaction functions. These feedback rules often include a measure of inflation, real activity and other, more specific indicators of the economic environment (e.g. exchange rates, foreign interest rates and financial market conditions). Recent empirical estimates of the monetary policy reaction function take into account the fact that the conduct of monetary policy changes over time. Monetary policy may adapt due to changes in the economic environment or other, more abrupt events such as a switch from exchange rate targeting to inflation targeting. This was the case for most CESEE countries during the early 1990s. Another reason why policy rules change over time may be changes in the composition of monetary policy committees, which has recently been demonstrated by Jung and Kiss (2012). More generally, Orphanides (2004) has shown that most central banks' preferences have changed since the 1970s.

In this paper, we aim to close a gap in the literature by estimating a time-varying monetary policy rule for four CESEE economies which are currently pursuing inflation targeting, namely the Czech Republic, Hungary, Poland and Romania. Research on monetary policy rules in CESEE economies is scant, and the number of studies that estimate feedback rules that adapt to changes in the underlying macroeconomic conditions is even more limited. Also, existing research only covers the time span up to 2009, therefore we aim to shed light on how traditional monetary policy rules fit the most recent past, which has been characterized by major central banks, including the European Central Bank (ECB), pursuing loose monetary policies and low inflation in the euro area and the countries under consideration. To this end, we use a novel econometric approach in the spirit of Nakajima and West (2013) that allows estimating time-varying monetary policy

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rules for a rich dataset covering additional control variables identified in the existing literature. This approach does not only feature time-varying coefficients but a time-varying variable selection in a coherent statistical fashion. This means that we can find out whether interest rates were adjusted to particular domestic macroeconomic developments in one period, while this was not the case in another period.

Our results can be summarized as follows: First, for the period from mid-2000 to early 2015 we find that all CESEE economies under consideration respond strongly to changes in (expected) domestic inflation, while there is almost no evidence of policy rates adjusting to deviations in real activity. This result is in line with existing literature, which suggests that the CESEE countries under consideration seem to follow a comparatively strict version of inflation targeting. Second, we show that short-term interest rates in the euro area play an important role in domestic interest rate setting in the countries covered; however, the significance of this effect has declined recently. This may be explained by the fact that the three-month EURIBOR is not an optimal proxy to reflect the unconventional monetary policies that were adopted after the outbreak of the crisis. Third, we find that in Hungary and Romania, interest rates tend to adjust to movements in exchange rates, while results for the Czech Republic and Poland are less clear. Finally, for all four countries we find that interest rates are rather persistent. That is, changes in interest rates appear to happen gradually rather than abruptly.

The remainder of this paper is structured as follows: Section 1 provides a brief literature review on time-varying monetary policy rules on the one hand and monetary policy rules in CESEE on the other hand. In section 2 we introduce an augmented monetary policy rule in the spirit of Taylor (Taylor, 1993) and the econometric framework that allows for variable selection and time-varying parameters in a coherent way. Section 3 presents the empirical results and section 4 concludes.

1 Literature review

Changes in monetary policy reaction functions over time have been taken into account in several ways. Early work used sample splits at presumed break dates to estimate both samples independently (see, e.g., Clarida et al., 2000) or used dummy intercepts to account for different periods. While splitting samples is a simple way of dealing with time-varying coefficients, it has two disadvantages: First, the optimum point in time for a split has to be assumed, which is not always an obvious choice, and second, this implies that changes in coefficients are modeled as a change-point process, implying abrupt changes of the underlying structural coefficients. Another strand of literature makes use of Markov-switching regimes, typically distinguishing between a high- and a low-inflation regime (see, e.g., Assenmacher-Wesche, 2006). Compared to simple sample splitting, this method offers the advantage that no break date or threshold has to be predefined. The nature of Markov-switching regimes, however, does not allow for gradual adjustments in monetary policy but assumes an abrupt change of how monetary policy is conducted instead. While there may be occasions that would justify abrupt changes (such as a general switch from exchange rate to inflation targeting) a more realistic approach would allow for gradual changes. Such time-varying monetary policy rules can be estimated by drifting coefficients within a state-space model, brought into the monetary policy rule literature by Boivin (2006), who examined Taylor rules estimates for the U.S.A. It enables the estimation of smooth changes without any prior assumption of sample breaks.

Since most empirical work on monetary policy rules has focused on the U.S.A. or other advanced economies, literature on monetary policy rules in CESEE is rather limited. Most studies (e.g. Mohanty and Klau, 2005; Paez-Farell, 2007; or Ghatak and Moore, 2011) use coefficients that are constant over the whole time span, and in some of these papers the time range starts already in the 1990s. The assumption of a linear feedback rule with coefficients that are constant over the sample period seems especially unrealistic for CESEE economies for two reasons: First, these countries underwent a major economic transition in the 1990s, and second, three out of four countries in our sample switched from an exchange rate-targeting framework to inflation targeting in the late 1990s or early 2000s. Petreski (2011) lists inferred and official switch dates from exchange rate to inflation targeting for the Czech Republic, Hungary[2] and Poland. Romania currently follows a managed floating exchange rate regime, in line with using inflation targets since 2005 as a nominal anchor for monetary policy. In addition, the economic environment changed considerably during the last decade, from economic boom phases in the mid-2000s to periods of severe contraction brought about by the global financial crisis and a period of recovery and loose international monetary policy. Given these facts, it seems unlikely that a linear feedback rule can appropriately characterize monetary policy for the countries considered in this study.

In the early literature addressing the non-linear nature of the monetary policy rule, usually sample splits or simple dummy intercepts are used to capture different regimes of monetary policy (see, e.g., Frömmel and Schobert, 2006; Yilmazkuday, 2009). More recently, Frömmel et al. (2011) and Petreski (2011) both used a Markov-switching approach to estimate monetary policy rules for CESEE economies. All these authors find that the countries under consideration in this study reacted strongly to inflation in the recent past, reflecting the move from exchange rate targeting to inflation targeting. With regard to the Czech Republic and Hungary, there is some evidence that contemporary monetary policy decisions are additionally driven by the exchange rate, the output gap and foreign interest rates (only for the Czech Republic). In contrast, for Poland and Romania most authors find that none of these additional variables significantly influence the interest-setting process.[3]

2 Data and empirical methodology

In this section we describe the methodology and data we use to evaluate Taylor rules for the Czech Republic, Hungary, Poland and Romania. Specifically, we estimate an augmented forward-looking Taylor rule where the coefficients are allowed to change over time.

[2] *Hungary kept an official exchange rate band of +/–15% against the euro until the beginning of 2008 (Frömmel et al., 2011).*

[3] *The impact of the leu's exchange rate on interest rate setting in Romania has not been investigated yet.*

2.1 Monetary policy rules

In his seminal paper, John Taylor (Taylor, 1993) proposed a rule that describes the reaction function of the U.S. Federal Reserve (Fed). In its simplest version, the Taylor rule postulates nominal interest rates as a function of inflation and real activity:

$$i_t^T = \bar{r} + \pi^* + \kappa\left(\pi_{t+12} - \pi^*\right) + \gamma(y_t - \bar{y}) \tag{1}$$

The rule states that interest rates (i_t^T) should rise if expected inflation (π_{t+12}) exceeds the inflation target set by the central bank (π^*) or if output (y_t) increases above its trend value ($\bar{y}$). As a by-product, the Taylor rule pins down the long-run neutral interest rate $\bar{r}$.[4] Note that this is a forward-looking version of the Taylor rule[5] since we include expected inflation as opposed to historical or contemporaneous inflation. These rules, as advocated in Clarida et al. (1998, 2000), provide a more realistic characterization of monetary policy as the interest-setting behavior of central banks is generally forward looking. Related to this debate is a discussion raised by Orphanides (2001) about the importance of using real-time data instead of data that are revised after a monetary policy decision has been made. While we do not have real-time data on output available for use, we construct the output gap based on an expanding-window estimation to better mimic the central bank's historical information set at each point in time compared to filtered data using all ex-post available data.

By estimating κ and γ it is possible to investigate whether the central bank stabilizes deviations of inflation from target and the output gap (Assenmacher-Wesche, 2006). More specifically, and to fulfill the Taylor principle, the coefficient on the deviation of inflation from target should be greater than 1 and the one attached to the output gap positive (see e.g., Woodford, 2001). Since nominal interest rates naturally respond one-for-one to increases in inflation (if the Fisher equation holds), a coefficient that is exactly unity would imply that the central bank would not sufficiently counteract inflation movements.

Since the Taylor rule was originally put forward to characterize monetary policy in the U.S.A., we extend the rule to account for structural features of small open economies. Albeit all countries considered in this study switched from exchange rate to inflation targeting prior to our sample period, several authors have argued more generally to control for exchange rates when estimating a Taylor rule (e.g. Clarida et al., 1998; Taylor, 2001), and especially so when looking at catching-up economies. In these economies, the exchange rate plays a more vital role than in advanced economies, where most domestic and foreign transactions are in local currency, markets are deeper, and the private sector is better equipped for absorbing exchange rate changes (Ostry et al., 2012). A number of studies have found that emerging market inflation targeters often (implicitly) include the exchange

[4] *Estimating the neutral rate via a Taylor rule has recently been shown to improve inflation forecasts (Horváth, 2009). See also Mehrotra and Slacík (2009), who show that actual money growth in relation to a McCallum rule provides information about inflation developments over a horizon of ten quarters for four CESEE economies.*

[5] *In the literature, the Taylor rule is often stated as $i_t^T = \pi_t + 2 + 0.5(\pi_t - 2) + 0.5(y_t - \bar{y})$. This is simply a rearranged version of equation (1), in its backward-looking form using actual inflation along with Taylor's findings for $\kappa = 1.5$, $\gamma = 0.5$. Moreover Taylor assumed that the Fed effectively followed an inflation target of 2% between 1987 and 1993, and that the long-run real interest rate was also 2%.*

rate in their interest rate reaction function (see, e.g., Mohanty and Klau, 2005, and Aizenman et al., 2011).[6] In line with the majority of the literature, we include a contemporaneous measure and a lagged measure for exchange rate movements (er_t, er_{t-1}), where generally an increase in the measure signals a depreciation of the local currency. This comes with the benefit that, depending on the size and sign of the estimated coefficients, the reaction of the central bank can be classified by rules of thumb discussed in Taylor (2001) and Frömmel and Schobert (2006). For example, a country pursuing an explicit exchange rate target is expected to cut interest rates when faced with appreciation pressures. This would be reflected in significant positive coefficients on both the current and lagged values of the exchange rate. In other cases discussed in Frömmel and Schobert (2006), the central bank reacts only temporarily to movements in the exchange rate. With respect to CESEE countries, Frömmel et al. (2011) experiment with different indicators of exchange rate movements and conclude that augmenting the Taylor rule improves estimation results. In what follows we will experiment with levels and period-on-period growth rates of the national nominal exchange rate against the euro, the real effective exchange rate, and the exchange market pressure index (EMP), which captures depreciation pressures on the currency and deviations of the currency from its long-run trend. In addition we include the three-month EURIBOR as a foreign interest rate and the EONIA as a robustness check.[7] Including a foreign interest rate in the Taylor rule can be motivated from an uncovered interest rate parity perspective. The uncovered interest rate parity basically relates domestic interest rates to foreign interest rates and exchange rate expectations. While it is hard to empirically find evidence of the validity of the uncovered interest parity in the short run, Chinn and Meredith (2004) show for a set of G-7 countries that it holds in the long run. Consequently, and with respect to the countries considered in this study, foreign interest rates have been frequently included in monetary policy rules (see e.g., Arlt and Mandel, 2014; Horváth, 2009; Vašíček, 2010).

To complete the model, note that policy rates typically show a very persistent pattern over time since central banks tend to adjust interest rates gradually. The dynamics of adjustment of the actual level of the interest rate to its target is then given by:

$$i_t = \left(1 - \sum_{p=1}^{2} \rho_p\right) i_t^T + \sum_{p=1}^{2} \rho_p i_{t-p} \tag{2}$$

That is, the central bank conducts interest rate smoothing by setting the actual rate as a linear combination of what is implied by the Taylor rule (first part in

6 *Albeit in a textbook inflation-targeting setting, the exchange rate should only affect an inflation-targeting central bank's interest rate through its impact on expected inflation, a more pragmatic approach should admit the importance of the exchange rate in the case of catching-up economies and provide some leeway within the inflation target framework (Ostry et al., 2012).*

7 *The results based on the EONIA are available from the authors upon request. Replacing the three-month EURIBOR by the EONIA yielded very similar results to those presented in section 3. In fact, correlations of estimated coefficients based on the estimations presented in section 3 and the robustness exercise including the EONIA are mostly close to 0.99. Only in Hungary, the coefficient on exchange rate developments shows a slightly smaller correlation of about 0.7.*

equation 2) and the historical interest rate (second part of equation 2). Our augmented Taylor rule that includes interest rate smoothing then becomes:

$$i_t = \left(1 - \sum_{p=1}^{2} \rho_p\right)\left[\bar{r} + \pi^* + \kappa\left(\pi_{t+12} - \pi^*\right) + \gamma(y_t - \bar{y}) + \psi er_t + \psi_1 er_{t-1} + \lambda eur_t\right] + \sum_{p=1}^{2} \rho_p i_{t-p} \quad (3)$$

where we have opted to include $p=2$ lags based on the marginal likelihood (approximated through the deviance information criterion).[8] Following Assenmacher-Wesche (2006), we assume that the central bank is able to control interest rates only up to a stochastic error u_t and subsume the long-run real interest rate and the inflation target into the constant $\alpha=\bar{r}-(\kappa-1)\pi^*$. We can then re-write equation (3) into

$$i_t = \left(1 - \sum_{p=1}^{2} \rho_p\right)\left[\alpha + \kappa\pi_{t+12} + \gamma(y_t - \bar{y}) + \psi er_t + \psi_1 er_{t-1} + \lambda eur_t\right] + \sum_{p=1}^{2} \rho_p i_{t-p} + u_t \quad (4)$$

Equation (4) is the model that is typically estimated in the literature. In this paper, however, we pursue a more flexible approach that allows estimated coefficients to vary over time. Accordingly the model becomes:

$$i_t = \left(1 - \sum_{p=1}^{2} \rho_{t,p}\right)\left[\alpha_t + \kappa_t\pi_{t+12} + \gamma_t(y_t - \bar{y}) + \psi_t er_t + \psi_{t,1} er_{t-1} + \lambda_t eur_t\right] + \sum_{p=1}^{2} \rho_{t,p} i_{t-p} + u_t \quad (5)$$

2.2 Econometric framework

This section introduces the econometric setting for the empirical analysis that follows. Let us assume that a time series $\{z_t, t=1,..., T\}$ is described by the observation equation

$$z_t = x'_t b_t + u_t, \quad (6)$$

where $x_t=(x_{1t},..., x_{Kt})'$ denotes a K-dimensional vector of possible explanatory variables measured in time t and $b_t=(b_{1t},..., b_{Kt})'$ is a $K \times 1$ vector of dynamic regression coefficients. Furthermore, let u_t be a normally distributed white noise error with zero mean and variance σ^2.

Following Nakajima and West (2013), we assume that the elements of b_t, b_{jt} $(j=1, ..., K)$ are related to a latent stochastic process β_{jt} as follows

$$b_{jt} = \beta_{jt} s_{jt}, \quad s_{jt} = I\left(\left|\beta_{jt}\right| > d_j\right). \quad (7)$$

Here $I(|\beta_{jt}|>d_j)$ denotes the indicator function which equals unity if the latent parameter β_{jt} exceeds a threshold $d_j \in \mathbb{R}$ to be estimated from the data. This implies that if β_{jt} is small, $s_{jt}=0$ and thus $b_{jt}=0$. Since β_{jt} evolves over time, this implies that b_{jt} could be non-zero for some points in time whereas for other periods it could equal zero. Thus $s_{jt}=0$ implies that there is no regression relationship between z_t and x_{jt} in time. The assumption that b_{jt} arises as a thresholded variant of β_{jt} provides a flexible and parsimonious means of modeling dynamic relationships and account-

[8] *Results based on p=3 lags generally yielded a worse fit and are available from the authors upon request.*

ing for model uncertainty, effectively capturing the notion that some variables may be relevant in some periods and less relevant in other periods.

To complete the model we also have to impose a law of motion for $\beta_t=(\beta_{1t},\ldots,\beta_{Kt})'$ where we assume for simplicity that it follows a random walk process

$$\beta_t = \beta_{t-1} + e_t. \tag{8}$$

Here, e_t is a vector white noise process with zero mean and a $K\times K$ dimensional variance-covariance matrix $V=diag(\vartheta_1, \ldots, \vartheta_K)$. Equation (8) is typically called a state equation for β_t. In general, it would be straightforward to assume that equation (8) is a stationary process. However, since the data are typically not very informative on the actual persistence of the latent states, and the length of our data is rather limited we take the simpler route and estimate equation (8).

The model outlined above belongs to the general class of state-space models and provides a large degree of flexibility in terms of modeling. For instance, our model is capable of unveiling changes in the underlying structural behavior of the time series analyzed. This proves to be of prime importance when used to investigate the behavior of a central bank. It is worth noting that our model nests a plethora of simpler models. For instance, if e_t equals the zero vector for all t then $\beta_t=\beta_{t-1}$ and equation (6) collapses to a standard linear regression model.

The model is estimated in a Bayesian fashion. This implies that suitable prior distributions have to be imposed on all parameters, which are described in more detail in the next section. Estimation is done using the Markov chain Monte Carlo (MCMC) algorithm described in Nakajima and West (2013). This implies that conditional on all other parameters the latent states are sampled by means of a Metropolis Hastings (MH) step. Likewise, we adopt a direct MH step to simulate the thresholds controlling the inclusion/exclusion of a given variable. For the remaining steps simple conditional posterior distributions exist, which makes Gibbs sampling feasible.

2.3 Prior distributions and implementation

We take a Bayesian stance to estimation and inference. More specifically, this implies that we have to specify suitable prior distributions on the parameters of the model. In the present application we use the following prior setup. We specify a standard inverted gamma prior on σ^2,

$$\sigma^2 \sim IG(\underline{a}_0, \underline{a}_1), \tag{9}$$

with $\underline{a}_0=0.01$, $\underline{a}_1=0.01$ being hyperparameters. The specific values chosen render the prior effectively uninformative. Furthermore, following Griffin and Brown (2010) and Bitto and Frühwirth-Schnatter (2014), we impose the following hierarchical priors on the elements of the initial state β_0,

$$\beta_{0j} \sim N\left(0,\tau_j^2\right),\ \tau_j^2 \sim G\left(a_\tau, \frac{a_\tau \xi_i}{2}\right),\ \xi_i \sim G(\underline{b}_0, \underline{b}_1). \tag{10}$$

Let $a_\tau=0.2, \underline{b}_0=0.01$ and $\underline{b}_1=0.01$ denote scalar hyperparameters chosen by the researcher. Finally, we impose a prior setup similar to the one described in the previous equation on the square root of the variances in V,

$$\sqrt{\vartheta_j} \sim N\left(0,\phi_j^2\right),\ \phi_j^2 \sim G\left(a_\phi, \frac{a_\phi \zeta_i}{2}\right),\ \zeta_i \sim G\left(\underline{c}_0, \underline{c}_1\right). \tag{11}$$

Similar to the prior on β_{0j} we set $a_\phi=0.2$, $\underline{c}_0=0.01$ and $\underline{c}_1=0.01$. This prior controls the degree of time-variation in the parameters and is thus crucial for the present application. Recently, researchers searched for priors that allow for significant variation in the autoregressive parameters while keeping the model stable, which is also the approach we follow in this paper.

The traditional prior setup used in this model relies on a pre-sample that is used to inform the corresponding prior distributions. In addition, hyperparameters are typically set such that the model is stable. Typically, this can be regarded as one of the main limitations of this modeling approach. However, in this paper we pursue a hierarchical approach that limits the problems regarding prior elicitation.

2.4 Data

For each country, our sample contains 131 monthly observations spanning the period from 7/2004 to 5/2015. The reason we do no extend the sample span to cover the period prior to 2004 is mainly due to the lack of reliable data on inflation expectations. Moreover, and as discussed above, by the start of the time period considered, all countries had already switched to inflation targeting. Hence, we use the longest sample available for estimating a forward-looking Taylor rule and cover various economic regimes and different phases of movements in exchange rates, prices and, ultimately, monetary policy. Data on policy rates are from Bloomberg and provided in percent per annum. As a measure for real activity, we select industrial production (including the construction sector, seasonally and working day adjusted, volume index). We calculate the output gap using a standard Hodrick-Prescott (HP) filter with $\lambda=129{,}600$. As mentioned earlier, the HP filter is estimated by using an expanding window to ensure that we take only information into account that was available to the central bank at that particular point in time. Inflation expectations for all countries considered are taken from the Hungarian central bank's quarterly inflation report and reflect one year-ahead inflation expectations of households. We use various measures to capture interest rate sensitivity to exchange rate movements. To this end, we collected data on the real (CPI-based) effective exchange rate from the Bank for International Settlements. Moreover we include national currencies' nominal bilateral exchange rates vis-à-vis the euro and the exchange market pressure index as advocated in Aizenman and Pasricha (2012) and Feldkircher et al. (2014). The EMP is defined as $emp_t=\left(\frac{e_t-e_{t-1}}{e_{t-1}}-\frac{ir_t-ir_{t-1}}{ir_{t-1}}\right)$, with e_t denoting the local nominal exchange rate per EUR 1 and ir_t standing for international reserves (minus gold) in U.S. dollars. Finally, we also include deviations of national currencies' nominal exchange rates vis-à-vis the euro from their long-run trend, estimated by the HP filter (fx_gap_t). For all measures related to exchange rates, an increase reflects a depreciation (pressure) of the currency and the related coefficient in the Taylor rule is expected to be positive. Lastly, data on the three-month EURIBOR, the average interest rate at which European banks are willing to lend money to each other over the three-month horizon, and the EONIA, the interbank rate at which banks lend overnight, are retrieved directly from the ECB's statistical data warehouse.

3 Empirical results

In this section we provide evidence whether monetary policy in the Czech Republic, Hungary, Poland and Romania can be adequately described by time-varying augmented Taylor rules. The main results are summarized in charts 1 to 4. In each chart, the first plot in the top panel on the left-hand side shows the domestic policy rate. The subsequent charts show, on the left-hand scale, the policy rate, coefficients attached to inflation expectations, the output gap, a measure for the exchange rate and the three-month EURIBOR. Moreover, we plot the coefficients associated with the lagged interest and exchange rates. In general, we present results for the specification that yielded the best in-sample fit and coefficients with the expected sign. The long-run transform of the coefficients is achieved by multiplying the respective coefficients with $\left(\frac{1}{1-\sum_{p}^{2}\widehat{\rho_{t-p}}}\right)$. In the charts, we further show the associated posterior inclusion probabilities (PIPs) plotted on the right-hand scale. They indicate how often a particular variable has been included in the Taylor rule model compared to how often the coefficient has been pushed toward zero. Variables with PIPs greater than 0.5 – as indicated by exceeding the dashed blue line in the charts – are considered as important regressors (Barbieri and Berger, 2004). Finally, the vertical line marks the collapse of the investment bank Lehman Brothers in September 2008, which is generally seen to mark the outbreak of the global financial crisis.

Chart 1 shows the results for the *Czech Republic*. Looking at the main components of the Taylor rule first, we see that our findings reveal both positive coefficients on inflation expectations and deviations of output from its long-run trend. The coefficient on deviations from trend output, however, receives only little support in the data as indicated by small posterior inclusion probabilities throughout the sample period. This finding is in line with Horváth (2009), who examines a broad range of Taylor rules, including backward- and forward-looking rules and rarely finds evidence for output stabilization for the Czech Republic. Vašíček (2010) argues that output gaps are typically estimated with a large noise component, which can be even more pronounced for emerging economies. The coefficient on inflation expectations is small but positive and hovers around 0.2 at the beginning of the sample. Hence our estimation results differ from those of Ghatak and Moore (2011) and Mehrotra and Sánchez-Fung (2011), who employ a linear regression framework and report insignificant (and negative) responses of interest rates to inflation. With the outbreak of the crisis, the coefficient on inflation starts to decrease strongly. Note that the size of the estimated coefficients does not directly reflect the central bank's preferences regarding output relative to inflation stabilization. This is so because the coefficients and the weight the central bank puts on inflation versus output stabilization are related in a non-linear fashion (Svensson, 1998).[9] To recover central banks' preferences, we would have to pin down an optimal monetary policy rule in connection with a structural model for

[9] *Assenmacher-Wesche (2006) estimates non-linear Taylor rules for three industrialized countries and, in a second step, maps the reduced form estimates into a loss function of the central bank.*

the economy under consideration first. This is in contrast to the modeling approach taken in this paper, which relies on a standard but ad hoc formulation of the monetary policy reaction function. See, among others, Assenmacher-Wesche (2006), Favero and Rovelli (2003) and Castelnuovo and Surico (2003) for structural estimations of central banks' preferences. The small coefficient attached to inflation (expectations) might be seen as a violation of the Taylor principle which ensures inflation stabilization only if the respective coefficient exceeds unity. However, as noted in Linnemann and Schabert (2006), the interpretation for small open economies is less straightforward since here additional variables typically appear as important ingredients in empirical Taylor rules. Estimates of sensitivity to inflation are close to those of Petreski (2011), who uses a Markov switching approach to estimate the Taylor rule, and Vašíček (2010) using a linear model. Next, we examine the reaction of the central bank to deviations of the exchange rate from its long-run trend. A positive coefficient would indicate that the central bank increases the policy rate in case the exchange rate depreciates. Looking at the coefficients, we see that the contemporaneous coefficient is negative. However, the coefficient on lagged deviations of the exchange rate from its long-run trend is positive and the sum of both coefficients seems balanced. In general, this finding indicates that the exchange rate is not included in the central bank's reaction function – or, put differently, that the interest rate is not used to stabilize the exchange rate. This does not, however, imply that the Czech central bank does not pay close attention to exchange rate movements. In fact, it decided in autumn 2012 to use the exchange rate as a monetary policy instrument and commenced foreign exchange interventions a year later with the aim of letting the koruna not appreciate well above CZK/EUR 27. That is, while the policy rate is the main monetary instrument to achieve price stability, foreign exchange market operations are used as an additional central bank instrument operating alongside the monetary policy/Taylor rule. See Ostry et al. (2012) and Benes et al. (2013) for analytical approaches to model exchange rate interventions for inflation-targeting economies. Our empirical results on the exchange rate also comply with the findings of Petreski (2011), who reports a positive coefficient for the period when the Czech central bank pursued an exchange rate target – not covered in our sample – whereas during the period of inflation targeting the coefficient on the exchange rate is not significant. Next, and in line with Arlt and Mandel (2014), Horváth (2009) and Vašíček (2010), we find that the policy rate of the Czech central bank reacts significantly and strongly to the three-month EURIBOR. More specifically, a 100 basis point increase in the three-month EURIBOR would suggest a 65 basis point increase in domestic interest rates at the beginning of the sample. In the most recent sample period this effect has somewhat abated, which might be related to the fact that the EURIBOR does not adequately reflect the monetary policy stance in times when monetary policy is characterized by unconventional measures. Lastly, our results point to a significant degree of interest rate smoothing. The sum of the two lags of the policy rate is close to unity. This implies that even in the case of strong and significant coefficients on some of the variables discussed above, the Czech central bank does not necessarily change its policy rate immediately. Whereas the other coefficients have shown marked variation over time, coefficients on lags of the policy rates have been pretty stable over the period considered. This implies that while the relative importance of output, inflation and exchange rate stabilization has varied over time, there has been no general regime shift of the Czech central bank's interest rate-setting behavior such as altering rates more frequently and aggressively during our sample period.

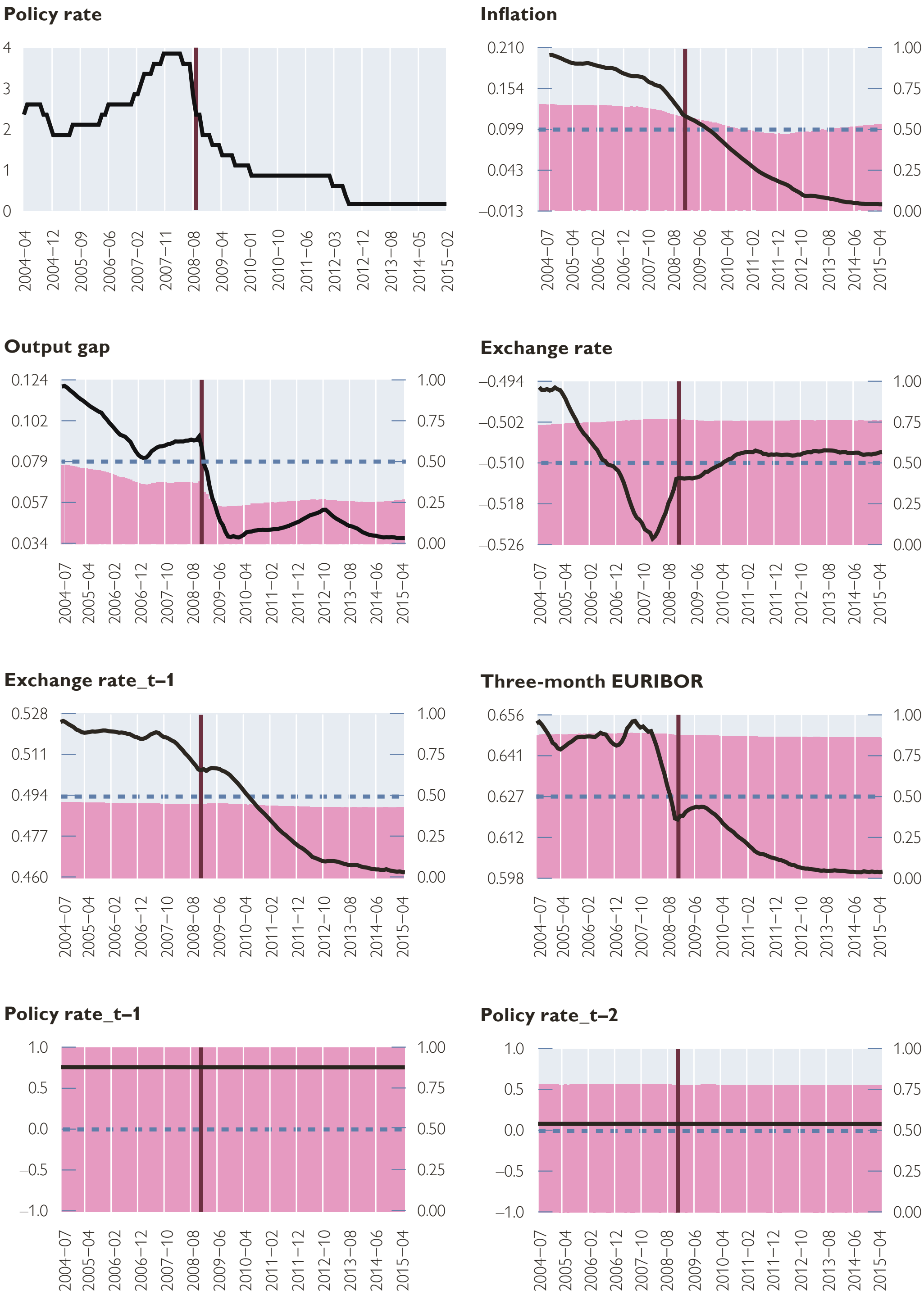
In chart 2 we depict the results for *Hungary*. Both estimated coefficients on inflation and output are positive but the latter is not significant throughout the sample period. In contrast to the results for the Czech Republic, the coefficient on inflation is close to unity at the beginning of the sample, after which sensitivity to inflation declines to about 0.7 in the most recent period of the sample, which is close to estimates provided in Frömmel et al. (2011) and slightly higher compared to the findings of Petreski (2011). The lack of significance in sensitivity of domestic interest rates to the output gap is in line with the existing literature (see e.g., Ghatak and Moore, 2011; Petreski, 2011; Jung and Kiss, 2012; Vašíček, 2010). Both coefficients on the exchange rate are positive, complying with the results of Yilmazkuday (2009) and Frömmel et. al (2011). Although the exchange rate band of +/–15% against the euro was officially not abandoned until February 2008, our results reveal that the importance of the exchange rate already decreased in the run-up to this date. The coefficient then rose slightly again in the subsequent years but lost ground from 2012 – a period in which the Hungarian authorities started to implement several measures to reduce household foreign currency debt and, consequently, decoupled movements in the exchange rate from national financial stability considerations. Note, however, that posterior inclusion probabilities throughout the sample period are above but close to the 0.5 threshold, indicating positive but weak support in the data. Next, we find that the Hungarian short-term interest rate adjusted also to the short-term interest rate in the euro area. This effect declined strongly from the beginning of the sample period to late 2007. From that period on, the effect started to rise again (to about 0.85). This result might indicate that – due to the abandonment of the exchange rate band – the euro area interest rate replaced the exchange rate in influencing monetary policy decision making. Compared to the Czech Republic, adjustments to the interest rate in the euro area are in general more pronounced in Hungary. Lastly, we also find a high degree of persistence in interest-setting behavior.

Chart 3 summarizes the results for *Poland*. Two observations stand out: First, among all the countries under investigation, the-long term inflation coefficient is the highest exceeding unity throughout the sample period. Moreover, the coefficient receives strong posterior support in the data, with PIPs of close to 0.8 for all time periods considered. In other words, Polish monetary policy is characterized by a strong commitment to inflation stabilization (Frömmel and Schobert, 2006). Second, the estimated Taylor rule for Poland is the only one that yields a positive and significant coefficient on the deviation of real activity from its long-term trend. This evidence is rather weak, however, since posterior inclusion probabilities exceed the 0.5 threshold only marginally. Next, we investigate whether interest rates adjust to movements in the exchange rate. As in the Taylor rule for the Czech Republic, the best fit is achieved by including a measure that indicates deviations from the złoty's long-run trend. Both coefficients on current and lagged exchange rate deviations are estimated with opposite signs, while they sum up to approximately zero, which indicates a zero net effect. Non-significant effects of interest rate adjustment to exchange rate movements are also reported in Petreski (2011). Interest rate adjustments in Poland are also influenced by short-term interest rates in the euro area. However, the estimated coefficient is much smaller relative to that of the other countries covered in this study. Also, the coefficient decreased somewhat, from about 0.4 at the beginning of the sample period to 0.36 in the most recent period of the sample. In line with the results for the Czech Republic and Hungary, we find evidence for interest rate smoothing with lagged coefficients on the policy rate summing up to close to unity.

Lastly, we investigate estimates for the augmented Taylor rule in *Romania* (chart 4). The variables affecting interest setting are inflation expectations, period-on-period change in the nominal exchange rate and short-term interest rates in the euro area. Interest rates adjust significantly to movements in inflation expectations. In line with the results for the other countries covered in this study, the coefficient attached to inflation was more pronounced at the beginning of the sample period (about 0.6) and less so during the most recent part of the sample period (about 0.4). There is no evidence in the data that monetary policy reacted to deviations of output from the long-run trend, as indicted by PIPs far below the 0.5 threshold; this complies with the results of Frömmel et al. (2011), Ghatak and Moore (2011) and Vašíček (2010). Both coefficients on exchange rate changes are positive, indicating that monetary policy tightened when the Romanian leu weakened against the euro, but of low significance. The estimated policy rule for Romania also features short-term interest rates in the euro area. Compared to the other countries in this study, the estimates for Romania are by far the most pronounced, ranging from about 1.5 at the beginning of the sample period to about 0.8 in the most recent period. The decline in the estimated coefficient is similar to the results for the Czech Republic, Poland and, partially, Hungary. Finally, we find evidence for interest rate smoothing. Compared to the other countries, the degree of smoothing is somewhat smaller and interest rate setting seems less persistent.

Chart 1

Czech Republic

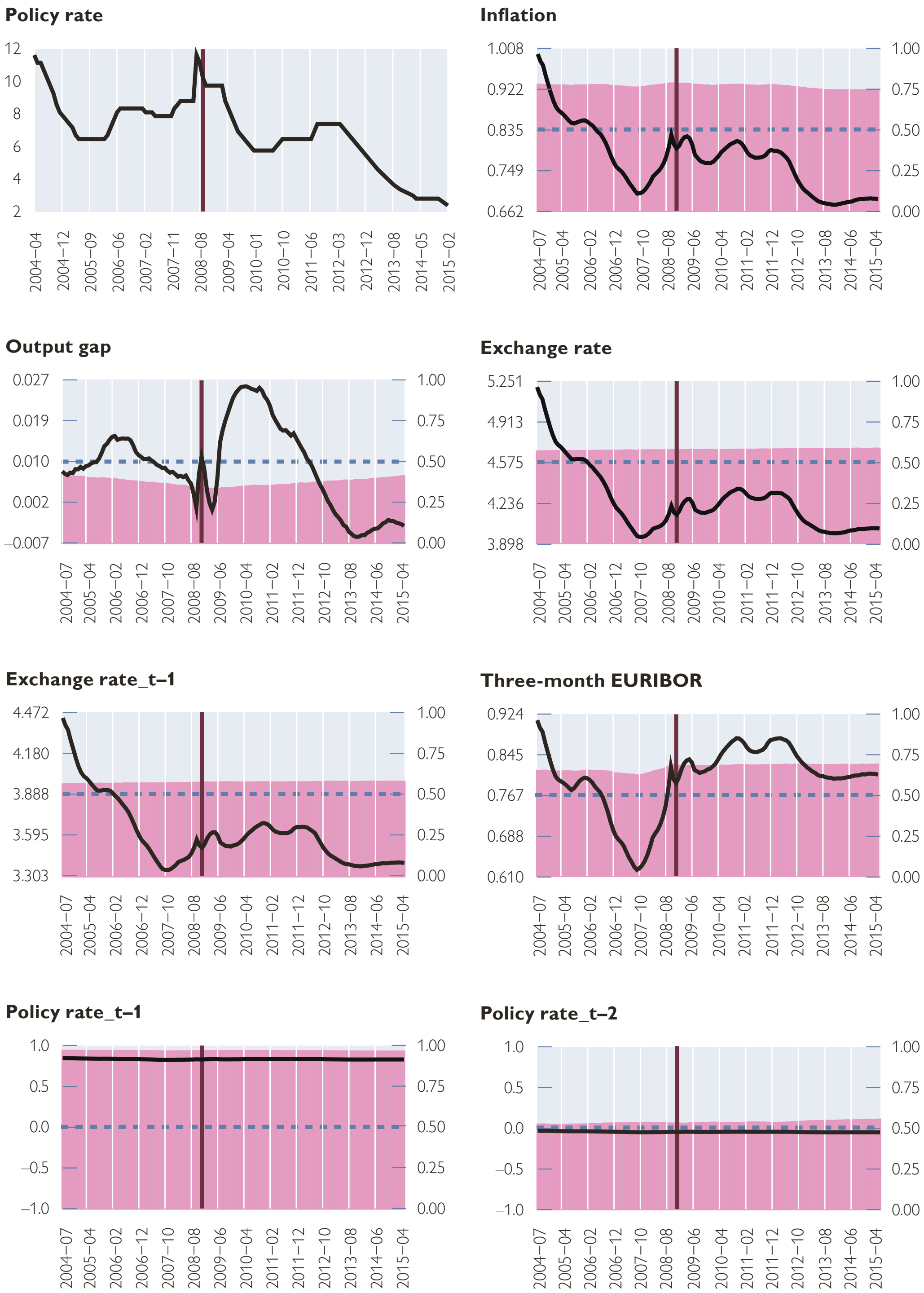


Source: Authors' calculations.

Note: The first plot depicts the domestic policy rate, whereas the remaining plots show the long-run coefficients (black solid line, left-hand scale) of inflation expectations (12 months ahead), the output gap, deviations of the exchange rate from its long-run trend, the three-month EURIBOR and short-run coefficients for two lags of the policy rate. The pink area refers to the time-varying posterior inclusion probability of the respective variable (right-hand scale). A variable should be included in the model if it receives a higher posterior inclusion probability than 0.5 (dashed blue line). The vertical line marks the outbreak of the global financial crisis.

Chart 2

Hungary

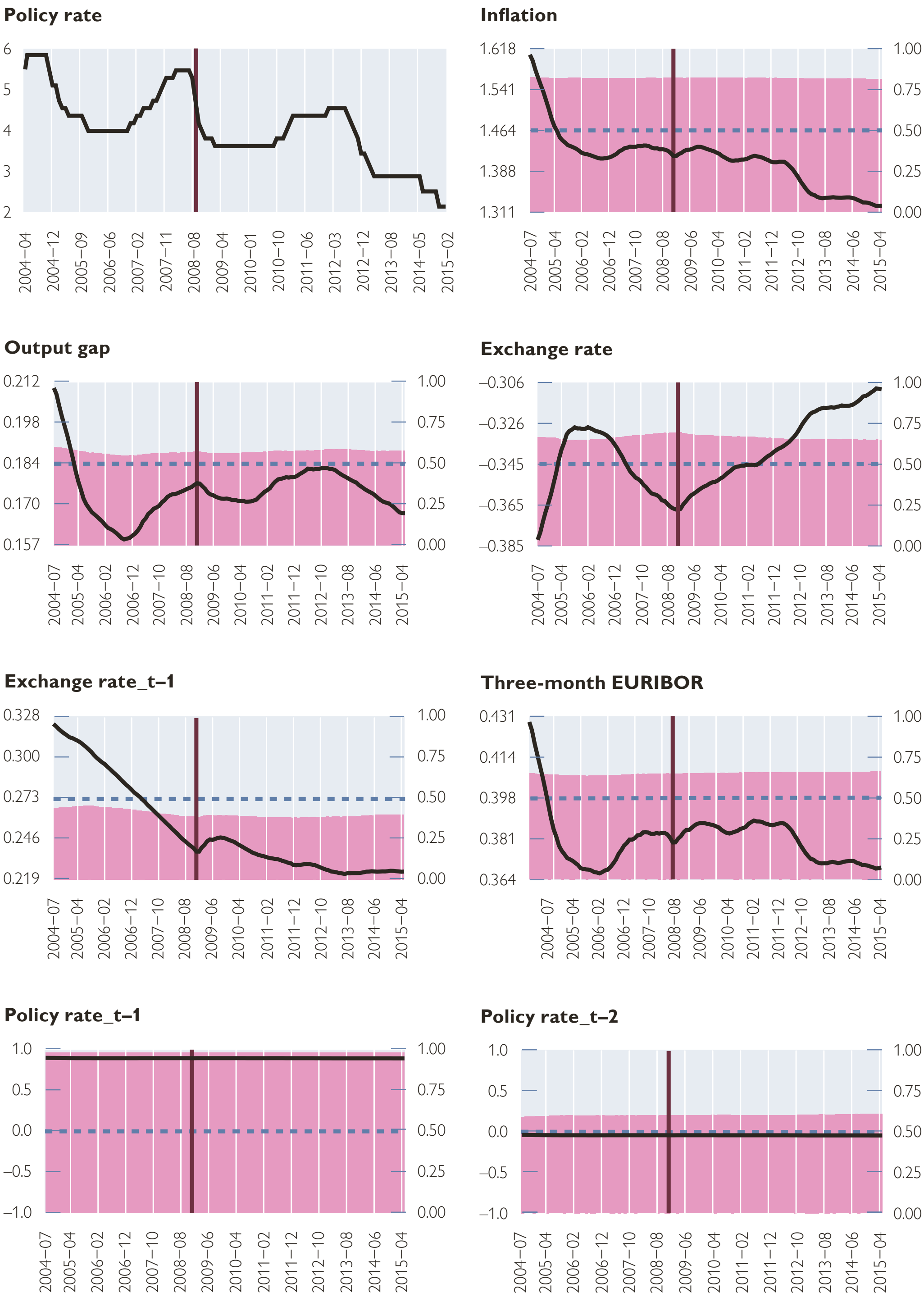


Source: Authors' calculations.

Note: The first plot depicts the domestic policy rate, whereas the remaining plots show the long-run coefficients (black solid line, left-hand scale) of inflation expectations (12 months ahead), the output gap, deviations of the exchange rate from its long-run trend, the three-month EURIBOR and short-run coefficients for two lags of the policy rate. The pink area refers to the time-varying posterior inclusion probability of the respective variable (right-hand scale). A variable should be included in the model if it receives a higher posterior inclusion probability than 0.5 (dashed blue line). The vertical line marks the outbreak of the global financial crisis.

Chart 3

Poland

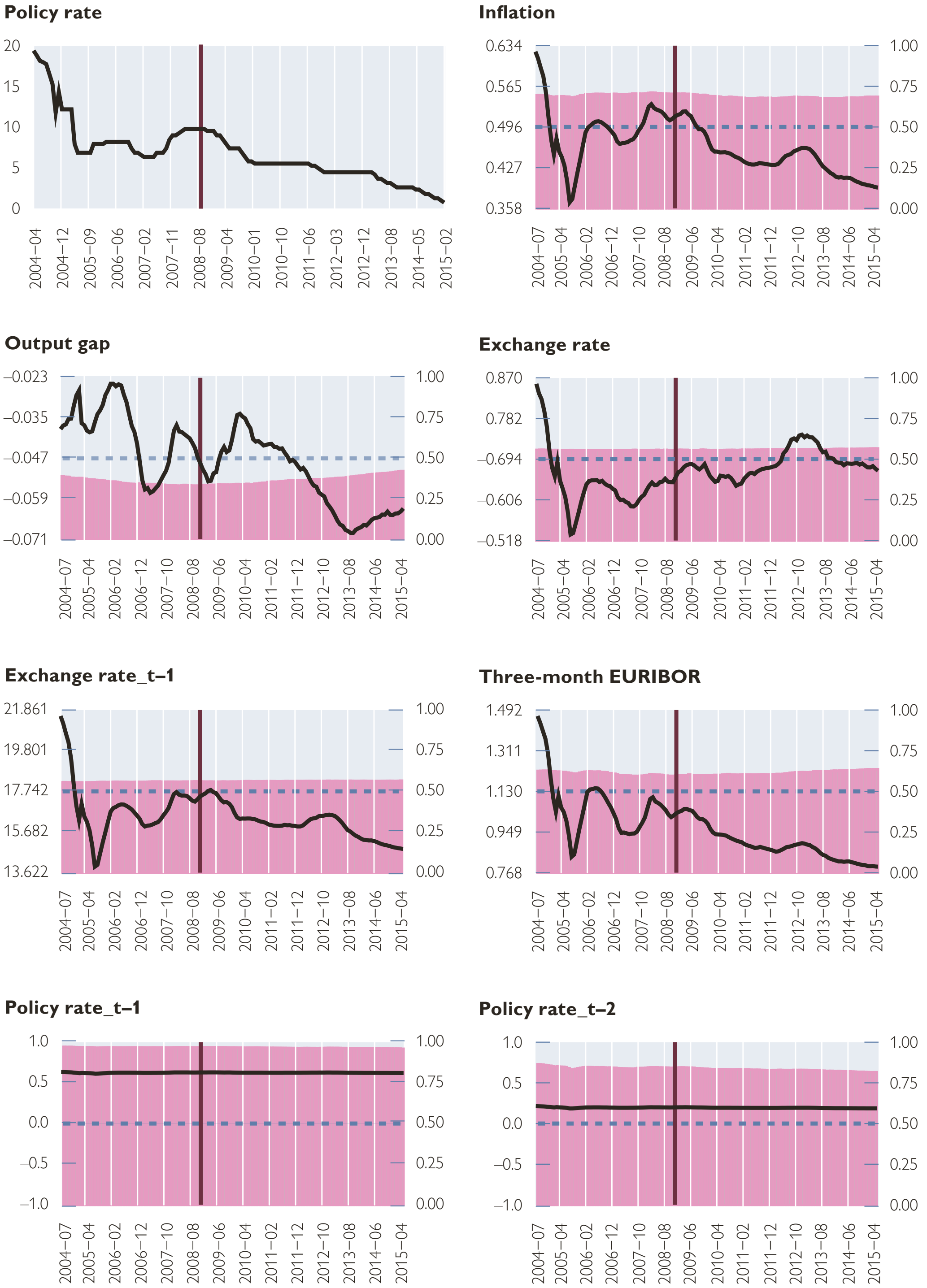


Source: Authors' calculations.

Note: The first plot depicts the domestic policy rate, whereas the remaining plots show the long-run coefficients (black solid line, left-hand scale) of inflation expectations (12 months ahead), the output gap, deviations of the exchange rate from its long-run trend, the three-month EURIBOR and short-run coefficients for two lags of the policy rate. The pink area refers to the time-varying posterior inclusion probability of the respective variable (right-hand scale). A variable should be included in the model if it receives a higher posterior inclusion probability than 0.5 (dashed blue line). The vertical line marks the outbreak of the global financial crisis.

Chart 4

Romania



Source: Authors' calculations.

Note: The first plot depicts the domestic policy rate, whereas the remaining plots show the long-run coefficients (black solid line, left-hand scale) of inflation expectations (12 months ahead), the output gap, nominal exchange rate vis-à-vis the euro (period-on-period growth), the three-month EURIBOR and short-run coefficients for two lags of the policy rate. The pink area refers to the time-varying posterior inclusion probability of the respective variable (right-hand scale). A variable should be included in the model if it receives a higher posterior inclusion probability than 0.5 (dashed blue line). The vertical line marks the outbreak of the global financial crisis.

4 Conclusions

In this study we estimate monetary policy rules for four of the largest CESEE economies, namely the Czech Republic, Hungary, Poland and Romania. During the last decade, these economies underwent different business cycle regimes, ranging from prolonged boom phases in the mid-2000s[10] to periods of contraction during the global financial crisis and a recovery phase characterized by a loose monetary policy stance of major central banks, including the ECB. Hence, we estimate monetary policy reaction functions that make it possible to take into account those changes by drawing on a novel econometric framework that features time-varying coefficients and variable selection in a coherent fashion.

Our results can be summarized as follows: First, we find evidence for a significant response of the policy rate to domestic expected inflation. Among all countries considered, the responses were most pronounced in Poland and less so in the Czech Republic. This does not necessarily imply that monetary policy in Poland has been more strongly anchored to price stability than in the Czech Republic. To clarify this issue, it would be necessary to assess how the estimated coefficients of the Taylor rule depend on the preference parameters of the central bank and on the structural parameters of the economy (Hayo and Hofmann, 2006).[11] Second, we find that output stabilization seems to play a minor role for interest setting in the countries considered, with the exception of Poland (Vašíček, 2010). Our results hence comply with the official communication of these countries' central banks, which seem to follow a comparatively strict version of inflation targeting. Third, we find that in Hungary and Romania, interest rates tend to adjust to movements in exchange rates. That is, if the national currencies face depreciation pressures, central banks will respond by increasing the policy rate. For the Czech Republic and Poland, the results are less clear. Fourth, we find that euro area short-term interest rates feature prominently in the estimated Taylor rules for all countries considered in this study. This does not necessarily imply that these countries' central banks aim to stabilize the domestic exchange rate via anchoring monetary policy to euro area short-term rates – which would contradict the weak evidence for exchange rate stabilization discussed above – since additional factors besides foreign interest rates determine exchange rates (e.g. country risk premia). Rather, this result shows how strongly these countries are integrated with the euro area and that changes in the euro area's monetary policy stance are likely to feed directly into the domestic economies (Babecká-Kucharčuková et al., 2014). Finally, looking at the time variation of the estimated monetary policy rules, we find that interest rates' adjustment to inflation has decreased in all CESEE countries under consideration, mirroring the low interest rate and low inflation environment in the most recent part of our sample period. That is, during the boom years in the mid-2000s interest rates adjusted more strongly to inflation than during the global financial crisis and its aftermath. Also, the importance of the three-month EURIBOR in the estimated Taylor rules decreased over time for all countries (except Hungary), but to different extents. Note that even against the

[10] *Hungary, which experienced a higher degree of volatility of economic growth, represents the only exception from this pattern.*

[11] *In other words, estimating central banks' preferences requires modeling the economy under consideration jointly with the Taylor rule – and even in this case some restrictive assumptions have to be made to recover the preferences of the monetary policymaker. (see, e.g., Castelnuovo and Surico, 2003, or Favero and Rovelli, 2003).*

backdrop of a declining trend of coefficients attached to inflation expectations and the EURIBOR, the data still show pronounced evidence of both variables being important components of monetary policy rules. The decline in the estimates for the three-month EURIBOR might be explained by the fact that this rate may be a good proxy for overall monetary policy conditions in the euro area during normal periods, but less so during times when monetary conditions are driven by unconventional measures (Babecká-Kucharčuková et al., 2014). Future work might look more closely into the direct effects of euro area quantitative easing on interest rate setting and, more generally, monetary policy in CESEE economies.[12]

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[12] *One approach that has resurfaced recently is capturing unconventional monetary policy via a "shadow interest rate" as proposed e.g. in Lombardi and Zhu (2014) and Whu and Xia (2015) that resembles a conventional policy rate during "normal" times and can turn negative during times the central bank provides additional stimulus by implementing nonstandard measures. However, as yet, there is no theoretical underpinning for embedding shadow rates into monetary policy rules.*

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