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Article

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Ekonomický časopis

Provided in Cooperation with:
Slovak Academy of Sciences, Bratislava

Reference: Pokorný, Tomáš/Chytilová, Helena (2023). Changes in monetary policy responses in the Czech Republic, Hungary and Poland. In: Ekonomický časopis 71 (6/7), S. 369 - 404.
https://www.sav.sk/journals/uploads/121309026-7%2023%20Pokorn%C3%BD%20+%20SR.pdf.
doi:10.31577/ekoncas.2023.06-07.01.

Changes in Monetary Policy Responses in the Czech Republic, Hungary and Poland[1]

*Tomáš POKORNÝ – Helena CHYTILOVÁ**

Abstract

In this paper, we analyze how the monetary policy responses of the Czech National Bank, the Hungarian National Bank, and the National Bank of Poland have changed over the period from 1996 Q1 to 2022 Q3 using the Kalman filter. Our findings are as follows. Firstly, we identify concealed discretion among the monetary policy authorities, which is traced through significant changes in their monetary policy parameters, highlighting the importance of applying a time-varying estimation framework. Secondly, we observe a decreasing response of policy rates to the inflation gap over the observed period. The reduced response in interest rates aligns with the anchored inflation expectations in the Czech Republic and Poland. Thirdly, the interest rate smoothing parameter has declined over the observed period, reflecting the central banks' increased willingness to adjust policy rates more vigorously. Additionally, we find a significant decrease in the long-term equilibrium interest rates, indicated by the policy-neutral rate, across the analyzed countries.

Keywords: *Central bank, monetary policy, inflation, time-varying parameters, Kalman filter*

JEL Classification: C32, E50, E52, E58

DOI: https://doi.org/10.31577/ekoncas.2023.06-07.01

Article History: *Received*: February 2023 *Accepted*: October 2023

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[1] This project was carried out with the financial support of the IGA/A Grant Competition at the Prague University of Economics and Business
(Project No. 8/2022, OP RDE IGA/A, CZ.02.2.69/0.0/0.0/19_073/0016936).

Introduction

Since Taylor (1993) wrote his groundbreaking paper, many economists have empirically analyzed the interest rate inflation-targeting reaction functions of central banks (Taylor rules). Traditionally, Taylor rules have been modeled with time-invariant (constant) parameters to illustrate the average monetary policy response over a specific horizon. To capture changes in the monetary policy response, researchers have increasingly employed time-varying parameters in Taylor rule analyses over the past two decades. Clarida, Galí and Gertler (2000), Kim and Nelson (2006), Baxa, Horváth and Vašíček (2014), and Feldkircher et al. (2016) have demonstrated the significant dynamic behavior of monetary policy parameters. Ignoring this dynamism could overlook crucial central bank decisions.

The central bank's governing council, composed of members with limited tenure, can alter its monetary policy stance. This change can stem from short-term temporary shocks excluded from decision-making or result from various political, economic, and social pressures (Favero and Rovelli, 2003; Valente, 2003). Additionally, central banks update their predictions and monetary policy analysis models, potentially leading to varying optimal monetary policy responses to inflation and business cycles.[2]

Yüksel, Metin-Ozcan and Hatipoglu (2013) proposed that the third source of time-variability in policy parameters is the evolving monetary policy transmission mechanism. As the economy evolves and household and firm behavior changes, the central bank adjusts its response to targeted variables due to differences in the effectiveness of the policy rate transmission mechanism. As discussed in Benati (2008), Zhang, Osborn and Kim (2008), and Baxa et al. (2014), firmly anchored inflation expectations allow the central bank to respond less firmly to the inflation gap.[3]

Drawing on Muth's (1961) rational expectations theory, the expected time-variability of monetary policy parameters is embedded in agents' rational expectations. Conversely, unanticipated changes are not factored in and can lead to erroneous economic decisions, potentially exacerbating business cycle fluctuations.

This paper investigates the evolution of monetary policy in three transition economies: the Czech Republic, Hungary, and Poland.[4] Specifically, we analyze how the Taylor rule parameters of the Czech National Bank, the Hungarian National

[2] For example, the Czech National Bank replaced separate regression models with the Quarterly Prediction Model (QPM) in 2002. In 2008, the model was further replaced by a modern dynamic stochastic general equilibrium (DSGE) model, g3. This model was updated to g3+ in 2019 to provide a more precise representation of the evolution of foreign variables (Brázdik et al., 2020).

[3] Presently, the transmission mechanism of policy rates might be weakened in CEE countries with independent monetary policy, as agents could borrow money in Euros due to lower interest rates. To our knowledge, this potential change in the transmission mechanism has not been thoroughly investigated yet.

Bank, and the National Bank of Poland have evolved. We also discuss the sources of this hidden discretion and its implications for economic performance. Despite the need for this understanding, changes in the response of central banks in Central and Eastern European (CEE) countries have not been adequately explored. The primary goal of this research is to address this gap by focusing on the analyzed central banks.

Our estimation relies on the Kalman (1960) filter combined with the full information maximum likelihood approach. We employ data from the period 1996 Q1 to 2022 Q3, with the investigation confined to the post-adoption of inflation targeting in all three countries. Consequently, the analyzed period is narrowed to 2001 Q2 to 2022 Q3, and the pre-inflation era serves solely for filter initialization.

Our findings reveal significant changes in monetary policy parameters, underscoring the importance of incorporating a framework of time-varying parameters when analyzing shifts in monetary policy. The results for all three central banks indicate a decline in response to inflation gap over much of the observed horizon, recently followed by an increase due mainly to energy price shocks. The decrease in the smoothing preference implies that central banks have become more open to stronger and less moderate policy rate responses. Furthermore, we identify a gradual reduction in policy-neutral rates to levels akin to those in the Eurozone.

The primary contribution of this paper lies in its employment of a more appropriate methodology for estimating time-varying parameters compared to existing CEE countries focused literature. Additionally, it unveils contemporary changes in the monetary policy response.

The structure of the paper is as follows: The first section offers an overview of the empirical literature related to time-varying Taylor rules and discusses the implications of these empirical findings for monetary policy. The latter part of this section presents findings from related literature that examine the cases of the Czech Republic, Hungary, and Poland. The second section discusses the data used. The third section outlines the methodology and presents our empirical model. Finally, the last section presents country-specific results and discusses our findings in relation to the related literature.

1. Literature Review

Several methods exist for tracing changes in the monetary policy parameters within central bank reaction functions.

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[4] It is noteworthy that we analyze the Visegrád 4 countries, excluding Slovakia, which has entrusted its monetary policy to the European Central Bank upon joining the Eurozone.

1.1. Sample Split Approach

One approach involves estimating the Taylor rule on sample splits to test whether the estimated parameters varied significantly across different periods. Various studies have employed this method. For instance, Judd et al. (1998) estimated the Fed's Taylor rule using OLS and divided the samples according to the Fed chairmen. They found that during Alan Greenspan's tenure, the Fed placed emphasis on both inflation and the output gap, with the response to the output gap twice that of Taylor (1993). An increase in weights on both inflation and the output gap was observed during Volcker's appointment in response to high 1970s inflation. Conversely, no significant response to inflation was found during Burn's period, while the weight on the output gap increased.

Clarida et al. (2000) used the GMM estimator on sample splits of the pre-Volcker and post-Volcker eras. Similarly, as Judd et al. (1998), they found that in Volcker's period the weight on inflation was considerably higher. Orphanides (2004) also examined these subperiods, noting the major difference was in weight on the output gap, which was higher pre-Volcker.

Martin and Milas (2013) used the GMM estimator and sample splits for UK data, revealing higher inflation gap weights in the pre-2007 crisis years.

Although the sample split approach is straightforward, unbiased, and consistent, it focuses on comparing average responses within subsamples and doesn't capture changes in parameters from period to period. This limitation underscores the promotion of alternative estimation methods, which this paper adopts.

1.2. Kalman Filter Approach

The Kalman (1960) filter is a precise technique that captures data generation processes, including parameter shocks in each period. The method needs to be combined with some other procedure which estimates the covariance matrices that are used for the calculation of the filtering equations.[5] Most of the papers use the maximum likelihood estimator (see the discussion below).

The Kalman filter combined with the maximum likelihood estimation was used in the following papers. Plantier and Scrimgeour (2002) found a decreasing trend in New Zealand neutral real interest rate (NRIR) in Taylor rule with time-varying NRIR. Jalil (2004) noted gradual changes in the monetary policy response to inflation in the second half of the 20th century for the US. Contrary to the results obtained via sample split approach in Judd et al. (1998), Clarida et al. (2000) and Orphanides (2004), his results suggest that the instability of the monetary policy reaction function is determined beyond who the chairman of the Fed is. Boivin

[5] See the Appendix, equation (A.13).

(2005) showed that the Fed's response to inflation was weak in the 1970s but then in the 1980s it was in line with the Taylor principle.[6] Trecroci and Vassalli (2006) emphasized the time-varying approach's explanatory power for UK, German, French, Italian, and US central bank Taylor rules. Trehan and Wu (2007) found significant variability of the parameters in Taylor rule for the US. In advance, they found that when the NRIR moves in the same direction as the trend growth rate, the probability of a change in inflation raised by the unperceived change in trend growth is lower. Mandler (2007) also analyzed the Fed's time-varying monetary policy coefficients and found that the uncertainty of agents about future values of the Fed's policy rate arises from three sources; one of them being the uncertainty about Fed's future policy time-dependent coefficients. This is a very important finding which confirms the theory which postulates that the central bank may de-anchor inflation expectations when it reacts time-variably to the same inflation level. Such a behavior may increase inflation volatility and can also be a secondary source of inflation in a situation when firms and households expected the central bank to react differently and increased prices more than they should have. Mésonnier and Renne (2007) estimated the ECB's Taylor rule with a time-varying natural interest rate and showed that this modified Taylor rule explains data better.

While the studies mentioned provide clear outcomes, a critical aspect often overlooked is the proper initialization of the Kalman filter. This oversight can substantially impact the results obtained, prompting some studies to employ alternative estimation procedures.

Cogley and Sargent (2001) employed a vector autoregression (VAR) with random coefficients in conjunction with the Kalman filter estimator to study post-war Fed Taylor rule changes. They efficiently initialized the Kalman filter by estimating a time-invariant VAR over the initial 40 quarters of their data.

However, this method faces limitations when datasets are insufficiently extensive, as we encounter in this paper, necessitating a distinct initialization approach (see the Appendix).

Kim and Nelson (2006) adopted a Heckman-type two-step maximum likelihood estimation (MLE) combined with the Kalman filter to estimate the Fed's forward-looking Taylor rule with time-varying weights. Their results favored dividing the history of US monetary policy into three periods rather than two as suggested by Orphanides (2004). To address the initialization problem, they employed a two-stage MLE procedure (Heckman procedure). Using this procedure, they use the estimates from the first stage as initial values for the second stage. However, this solution raises two key issues. First, it relies on accurate initial values from the

[6] These results are contrary to the findings of Clarida et al. (2000), who estimated that the response to inflation in the 1970s was in line with the Taylor principle.

first stage for precise values in the second stage. Second, it demands a sufficiently long dataset that isn't compromised by reduced observations in the initial stage. These challenges lead us to depart from this approach in our analysis.

Yüksel et al. (2013) investigated the time-varying Taylor rule of the Central Bank of Turkey and highlighted the structural-extended Kalman filter as superior to the standard Kalman filter for determining the bank's reaction function. Nonetheless, their solution for parameter initialization, based on ordinary least squares (OLS) estimates, might be insufficient due to endogeneity in the Taylor rule variables, potentially introducing bias to the estimated values.

In this study, we align with the mainstream approach, combining the Kalman filter with the full-information maximum likelihood estimator. This method enables the examination of period-to-period changes in the Taylor rules of the Czech National Bank, the Hungarian National Bank, and the National Bank of Poland, providing comprehensive insights into the dynamics of their monetary policy reactions. We address the filter initialization challenge through two strategies. Firstly, we calibrate the initialization values to minimize estimation prediction errors. Secondly, we solely employ pre-inflation targeting data to initialize the filter.

1.3. Other Approaches

Various alternative estimation methods can also capture changes in monetary policy parameters beyond sample splits and the Kalman filter. Several noteworthy techniques have been applied:

Wesche (2003) utilized the Markov-switching model to examine Taylor rule parameters for central banks in the UK, Germany, France, Italy, and the US. She classified monetary policy regimes into "dove" (high weight on output, low weight on inflation) and "hawk" (high weight on inflation, low weight on output) categories. Kuzin (2006) employed the Markov-switching model and the Kalman filter to estimate time-varying coefficients in a backward-looking Taylor rule for Germany. The findings suggested that the Bundesbank's aversion to inflation exhibited sudden and significant shifts, indicating an opportunistic approach to disinflation.

Partouche (2007) utilized the generalized method of moments (GMM) with smoothing splines to estimate the time-varying Taylor rule for the Fed. His results aligned with Clarida et al. (2000) and Orphanides (2004), revealing a notably higher response to inflation during Volcker's tenure and confirming that a higher weight on inflation stabilized the US economy.

McCulloch (2007) employed the adaptive least-squares estimation method on US data and identified significant variation in the coefficient on expected inflation. Notably, this coefficient increased during Volcker's chairmanship, decreased in the 1990s, and subsequently rose again in 2003.

Liu et al. (2018) employed the latent threshold time-varying vector autoregression (VAR) model to estimate time-varying Taylor rules for China and the US. Their results highlighted gradual changes in monetary policy parameters, revealing discretionary behavior of policymakers. They found that during economic expansion, the focus was primarily on inflation, while during recessions, the output gap held greater weight.

Baxa et al. (2014) utilized a varying coefficients method on time-varying Taylor rules for countries including Australia, Canada, New Zealand, Sweden, and the UK. They observed a decreasing interest rate smoothing parameter across all analyzed countries, indicating increasing openness by central banks to significant changes in policy rates. The results also suggested that a successfully anchored inflation expectation led to a diminishing response to the inflation gap.

Korhonen and Nuutilainen (2016) employed a rolling data ranges method combined with the GMM to scrutinize specifications of the Russian Central Bank's Taylor rule. They showed that in 2014 and 2015 the Bank's response to the output gap significantly increased and the response to the inflation gap decreased. While their procedure better depicted the data-generating process compared to sample splits, it couldn't fully account for ad hoc shocks to parameters spread across the entire rolling window period.

Although these alternative methods appear robust, their accuracy warrants further testing. In this study, we align with the mainstream literature, preferring the Kalman filter estimator.

In conclusion, findings from literature utilizing diverse estimation techniques provide robust evidence of the time-variability of monetary policy parameters, underscoring the importance to analyze monetary policy changes time-variably.

1.4. Empirical Evidence from Related Papers Analyzing the Targeted Countries

This subsection compiles findings from empirical literature examining time-varying Taylor rule parameters in the Czech Republic, Hungary, and Poland.

Frömmel and Schobert (2006) employed the GMM and sample splits to capture Taylor rule changes in six CEE countries, including the Czech Republic, Poland, and Hungary. They noted these countries transitioned from exchange rate targeting to inflation targeting, aligning with official monetary policy declarations. Horváth (2006; 2009) utilized the GMM and the Kalman filter to estimate diverse Taylor rule forms featuring time-varying equilibrium interest rates in the Czech Republic. His results indicated a gradual decrease in the equilibrium interest rate over time, reaching levels comparable to the euro area. Yilmazkuday (2008) introduced dummies for structural breaks in the Taylor rule to estimate monetary policy shifts in the Czech Republic, Hungary, and Poland. While Czech Republic and Poland showed evidence of inflation targeting introduction, Hungary's data did not support its National Bank's full adoption of inflation targeting.

Frömmel, Garabedian, and Schobert (2011) employed a cointegration approach incorporating shifts in exchange rate regimes to examine Taylor rules in CEE countries, including the Czech Republic, Hungary, and Poland. They established that these central banks shifted from exchange rate targeting to inflation targeting, aligning with their official declarations. Petreski (2011) utilized a switching regression to estimate time-invariable Taylor rules for the Czech Republic, Hungary, and Poland using instrumental variables and GMM. His findings corroborated the switch from exchange rate to inflation targeting in these countries, as found by Frömmel and Schobert (2006), Frömmel et al. (2011).

Michalek et al. (2012) employed the Kalman filter to estimate the natural real rate of interest (NRIR) for Poland. Their results indicated a gradual decrease in the NRIR during the period of 1998 – 2011. Feldkircher et al. (2016) employed a Markov chain Monte Carlo-based estimation procedure to analyze time-varying Taylor rule parameters for the Czech Republic, Hungary, Poland, and Romania. They observed a gradual decrease in central banks' response to the inflation gap, suggesting impacts from unconventional monetary policies and global and domestic price growth. Mackiewicz-Łyziak (2017) estimated time-varying parameters of the Central Bank of Poland's Taylor rule using GMM on sample splits. She found a decrease in smoothing parameter and weight on inflation gap and an increase in weight on output gap after the outbreak of the 2007 crisis.

Abaligeti, Németh and Schepp (2018) applied the Kalman filter to estimate forward-looking, time-varying parameters of the Central Bank of Hungary's Taylor rule. Their results indicated the Hungarian inflation targeting history could be divided into three periods characterized by distinct reactions of the Central Bank of Hungary.

While the methodologies in Frömmel and Schobert (2006), Yilmazkuday (2008), Frömmel et al. (2011), Petreski (2011), Mackiewicz-Łyziak (2017) are limited in capturing full dynamics of time-varying monetary policy parameters, Horváth (2006; 2009) Michalek et al. (2012) focus primarily on equilibrium interest rate dynamics without fully capturing other monetary policy parameter changes. Feldkircher et al. (2016) potentially introduced bias by using censored policy rates, noted by Podpiera (2008). Moreover, specifications in Feldkircher et al. (2016); Abaligeti et al. (2018) assume central banks are not forward-looking regarding the output gap.

Our research presents novel aspects. Firstly, we employ the Kalman filter as in Horváth (2006; 2009) while specifying the Taylor rule based on Clarida et al. (2000). Except for Horváth (2006; 2009), the Taylor rule specifications in the related literature mentioned in this subsection do not follow Clarida et al. (2000) in his specification. Unlike Horváth (2006; 2009) and Michalek et al. (2012), we allow all monetary policy parameters to be time-varying and we have more current data. Similar to Feldkircher et al. (2016), we analyze the time-varying Taylor rule of the Czech Republic, Hungary, and Poland. However, our study differs by using a distinct Taylor rule specification and mainstream Kalman filter estimation, rather than Markov chain Monte Carlo. Our paper aims to bridge the knowledge gap in tracking monetary policy changes in CEE countries using the widely-accepted Kalman filter estimator. Our results intend to validate findings from other approaches while encompassing the period of unconventional monetary policy tool usage and addressing the ongoing Covid-19 crisis, high inflation, energy crises, and the Ukrainian-Russian war.

2. Data

In the realm of related literature that examines the targeted countries, researchers typically rely on monthly data, see Frömmel and Schobert (2006), Horváth (2006), Yilmazkuday (2008), Horváth (2009), Frömmel et al. (2011), Petreski (2011), Feldkircher et al. (2016). However, with the exception of Feldkircher et al. (2016), this choice is typically necessitated by the scarcity of data observations. The rationale behind the use of higher frequency monthly data lies in their ability to capture changes in monetary policy parameters on a monthly basis. Nevertheless, it's worth noting that, as argued by Vašíček (2012), monthly data can be noisy, particularly for variables like inflation and interest rates that exhibit high persistence at this frequency.

A further challenge with employing monthly data arises when incorporating the output gap into the model specification. Unlike inflation and short-term interest rates, the GDP is not regularly measured on a monthly basis by statistical agencies. To derive monthly estimates for the output gap, one must resort to adjusting quarterly GDP time series data to the monthly frequency using interpolation methods, which could potentially introduce bias. This issue contributes to the motivation behind central banks' preference for employing quarterly data in their primary prediction models, see Budnik et al. (2009), Szilágyi et al. (2013), Brázdik et al. (2020).

Our quarterly time series are notably extensive, rendering the use of monthly data unnecessary. Thus, we align with the approach undertaken by Mackiewicz-Łyziak (2017), Michalek et al. (2012) and Abaligeti et al. (2018), opting to utilize the more robust fully observed quarterly data.

In this study, we utilize quarterly time series data encompassing the short-term interest rate, the inflation gap, and the output gap, spanning from 1996 Q1 to 2022 Q3. Importantly, it's worth noting that the period preceding the implementation of inflation targeting is exclusively used for filter initialization (discussed in the Appendix). Given that the Czech National Bank adopted inflation targeting in December 1997, the National Bank of Poland in September 1998, and the Hungarian National Bank in June 2001, our analysis restricts itself to a uniform data range spanning from 2001 Q2 to the 2022 Q3.

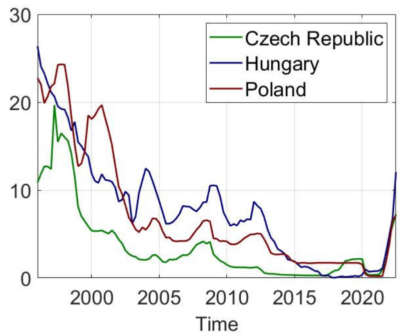
Descriptive statistics tables can be found in the Appendix, specifically Tables 2, 3, and 4.

2.1. Short-Term Interest Rate

For this variable, we utilize the time series of the short-term interest rate provided by the OECD. This interest rate specifically corresponds to short-term (3-month) government bonds issuance. The short-term interest rate is a crucial indicator, determined by the policy rate, and is preferred over policy rates themselves to avoid potential censoring issues, as elaborated in Podpiera (2008). It's worth noting that we encountered a challenge with a few missing values in the Hungarian short-term interest rate series, which we addressed using the interpolation method proposed by Stineman (1980).

Figure 1

Short-Term Interest Rate in The Czech Republic, Hungary and Poland



Note: The short-term interest rate in the Czech Republic, Hungary, and Poland declined significantly to approximately 0.5% before experiencing a notable increase in 2022.

Data source: OECD.

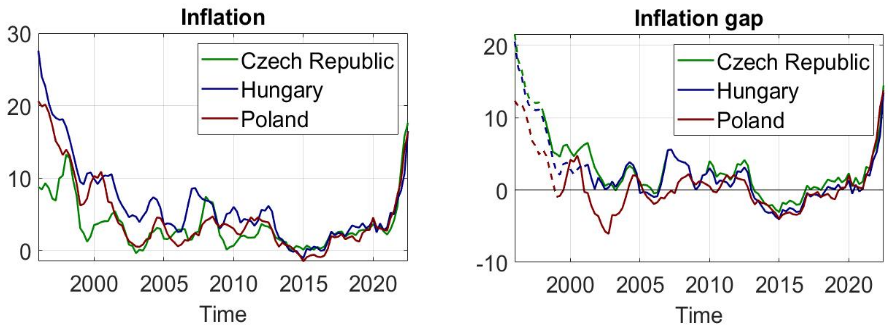
Figure 1 illustrates a notable trend: the short-term interest rate experienced a rapid decline in all three countries during the period 1995 – 2010. Starting from values of approximately 20%, it progressively dropped to levels well below 10% by 2010. This pronounced decrease was primarily propelled by the implementation of inflation targeting, which in turn led to the stabilization of inflation around the 2% mark (see Figure 2). Over the subsequent decade, interest rates exhibited significantly reduced volatility compared to the previous decade, fluctuating within the range of approximately 0.5 – 2% since 2015. This phase was largely characterized by persistent disinflationary forces, with periods of deflationary pressure between 2010 and 2020. However, the landscape changed dramatically following the Covid-19 pandemic and the Ukrainian-Russian conflict, prompting central banks to raise policy rates in order to counter the heightened inflationary pressures that emerged.

2.2. Inflation Gap

We derive inflation data from the OECD time series, which represents the year-on-year relative change in the consumer price index. To calculate the inflation gap, we subtract the annual inflation target of each respective central bank from the quarterly inflation rate.[7]

F i g u r e 2

Inflation and Inflation Gap in the Czech Republic, Hungary and Poland



Note: The first graph illustrates the gradual decline in inflation across the Czech Republic, Hungary, and Poland following the implementation of inflation targeting. In the second graph, we observe the inflation gap, representing the deviation from the central bank's target at assumed $t+2$ horizon. The dashed line demonstrates the theoretical evolution of the inflation gap if the central bank had announced the inflation target before adopting inflation targeting.

Data source: Inflation rate – OECD; inflation targets – Czech National Bank, Hungarian National Bank, National Bank of Poland; inflation gap – own calculations.

Figure 2 illustrates the inflation trend (first graph) and the inflation gap (second graph) for the studied countries. The dashed line in the second graph represents the pre-inflation targeting era, characterized by relatively high and volatile inflation

[7] The inflation targets for each year can be found in Table 1 located in the Appendix. When an interval was specified as the target, we consider the mean of that interval as the target value for the corresponding period.

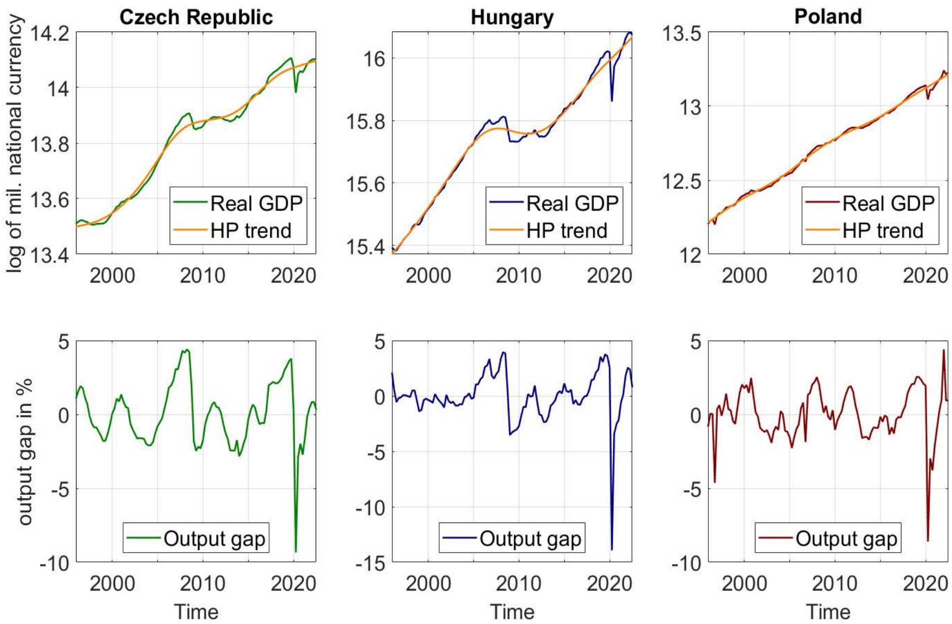
in all three countries. Over the last decade, a notable stabilization of inflation around 4% is evident, accompanied by reduced volatility in the inflation gap. The Czech National Bank stands out for effectively maintaining price stability compared to the Hungarian National Bank and the National Bank of Poland. Recent quarters have witnessed an escalation in inflation dynamics due to disruptions in global value chains, pandemic-induced production shutdowns, energy crises, and the Russian-Ukrainian conflict.

2.3. Output Gap

For the output gap calculation, we rely on real GDP data. The seasonally adjusted real GDP for the Czech Republic is derived from the ARAD database of the Czech National Bank, while the corresponding data for Hungary and Poland is obtained from the Eurostat database. To measure the output gap as a percentage deviation, we utilize the logarithm of these series.

Figure 3

Output and Output Gap in the Czech Republic, Hungary and Poland



Note: The upper graphs illustrate consistent GDP growth in the observed countries, with fluctuations during events like the 2020 COVID-19 pandemic and the 2007 financial crisis (Czech Republic and Hungary). The lower graphs display output gap deviations from potential, with the 2020 pandemic leading to significant business cycle fluctuations.

Data source: Czech GDP – ARAD, Hungarian and Polish GDP – Eurostat; HP trend and the output gap – own calculations.

To estimate the potential output of the Czech, Hungarian, and Polish economies, we employ the two-sided Hodrick and Prescott (1997) filter as in Horváth (2006; 2009) and Mackiewicz-Łyziak (2016; 2017).[8] The resulting cyclical component from the filter serves as a direct measure of the output gap. The graphs in Figure 3 showcase the time series of output and the estimated HP trend, displayed in a logarithmic scale of millions of national currency for each country (upper three graphs). Correspondingly, their respective output gaps are illustrated in the lower three graphs.

The real GDP in the Czech Republic, Hungary and Poland has a growing trend. Note that the Polish economy was more resistant to the financial crisis of 2007, in contrast to the Czech and Observing the real GDP trends, all three countries – the Czech Republic, Hungary, and Poland – exhibit a growing trajectory. Worth noting is Poland's resilience during the 2007 financial crisis, in contrast to the Czech and Hungarian economies. The subsequent crisis caused by COVID-19 and the associated economic restrictions led to a similar adverse impact on Polish GDP as it did on the Czech and Hungarian economies. The effect of this crisis is primarily reflected in higher output gaps, as depicted in the lower three graphs.

3. Methodology

Inflation targeting policy can be performed using a simple monetary policy rule, exemplified by Taylor's (1993) proposal, where the nominal policy rate r_t^* responds to inflation π_t and the output gap x_t:

$$r_t^* = rr^{nat} + \pi_t + (\pi_t - \pi^*) + 0.5x_t \tag{1}$$

where rr^{nat} is the equilibrium real interest rate and π^* is the inflation target.

We can generalize (1), following Orphanides (2010) or Wang and Wu (2012):

$$r_t^* = r^{nat} + \xi_\pi(\pi_t - \pi^*) + \xi_x x_t \tag{2}$$

where ξ_π and ξ_x are the central bank's weight parameters for the inflation and output gaps respectively, and r^{nat} represents the policy-neutral rate.[9]

Considering the delayed impact of monetary policy, a forward-looking Taylor rule can be more appropriate:

[8] We do not use one-sided HP filter as in Feldkircher et al. (2016), because of the Wolf et al.'s (2020) critique. Wolf et al. find that the one-sided filter fails to remove low-frequency fluctuations and has the undesirable feature if dampening fluctuations that one wishes to extract.

[9] Following I. Fisher and Barber (1907), the *policy-neutral* rate can be approximated as the sum of inflation and the equilibrium real rate.

$$r_t^* = r^{nat} + \xi_\pi (E\{\pi_{t+n} | \Omega_t\} - \pi^*) + \xi_x E\{x_{t+n} | \Omega_t\} \tag{3}$$

where Ω_t is the set of information that the central bank has at its disposal when it decides on the policy rate level. *E* denotes the expectations of the central bank over the targeted economic variable at the monetary policy horizon *n*.[10]

Additionally, central banks might prefer gradual policy changes. Following Clarida et al. (2000), the smoothing of the interest rate can be tracked by a simple partial adjustment mechanism:

$$r_t = \rho r_{t-!} + (1-\rho) r_t^* \tag{4}$$

Combining equations (4) and (3) gives:

$$r_t = (1-\rho)[r^{nat} + \xi_\pi (E\{\pi_{t+n} | \Omega_t\} - \pi^*) + \xi_x E\{x_{t+n} | \Omega_t\}] + \rho r_{t-1} \tag{5}$$

Following Kim and Nelson (2006), Baxa et al. (2014), Feldkircher et al. (2016) and other empirical literature discussed in subsection 1.4, it is appropriate to assume that the Taylor rule parameters are time-varying rather than constant. Therefore, we rewrite the equation (5) in a time-varying

$$r_t = (1-\rho_t)[r_t^{nat} + \xi_{(\pi,t)} \left(E\{\pi_{t+n} | \Omega_t\} - \pi_{t+n}^* \right) + \xi_{x,t} E\{x_{t+n} | \Omega_t\}] + \rho_t r_{t-1} \tag{6}$$

Note that we allow the inflation target π_{t+n}^* to be time-varying, which captures the changes in the inflation targets of the Czech National Bank, Hungarian National Bank and National Bank of Poland.[11]

3.1. Model

We adopt the mainstream approach used in studies like Cogley and Sargent (2001), Jalil (2004), Boivin (2005), Kim and Nelson (2006), Trecroci and Vassalli (2006), and Trehan and Wu (2007) to estimate the time-varying Taylor rule parameters using the Kalman (1960) filter. The Kalman filter operates within a state-space framework,[12] enabling the estimation of unobservable variables and parameters.

To apply the Kalman filter, we need to rewrite the model into a state-space form. We can perform that by using simple transformation of parameters in (6) such that $\psi_{r^{nat},t} \equiv (1-\rho_t) r_t^{nat}$, $\psi_{\pi,t} \equiv (1-\rho_t)\xi_{\pi,t}$ and $\psi_{x,t} \equiv (1-\rho_t)\xi_{x,t}$. Using the simplifying assumption of the central bank's perfect foresight, we get:

[10] The *horizon* of the monetary policy is about 4 – 6 quarters.

[11] See Table 1.

[12] A state-space model is a system of difference equations describing the state of some variables given the dynamics of these variables in other periods, for details see Hamilton (1994).

$$r_t = \psi_{r^{nat},t} + \psi_{\pi,t}(\pi_{t+n} - \pi^*_{t+n}) + \psi_{x,t} x_{t+n} + \rho_t r_{t-1} + \varepsilon_t \tag{7}$$

where ε_t is the error term which is composed of forecasts error and is orthogonal to the information set available at time t (Ω_t). We set the horizon of the monetary policy (indexed by n) to 2 in line with Baxa et al. (2014). This is a simplifying assumption because central banks usually consider horizon of 4 – 6 quarters. We face similar reasons as Baxa et al. (2014) and prefer a lower horizon to mitigate rising prediction errors with longer horizons.[13]

By introducing an inflation gap ($\pi^{gap}_{t+n} \equiv \pi_{t+n} - \pi^*_{t+n}$), equation (7) can be rewritten as an observation equation[14] of the state space model in the following matrix expression:

$$r_t = [1 \quad \pi^{gap}_{t+2} \quad x_{t+2} \quad r_{t-1}] \begin{bmatrix} \psi_{r^{nat},t} \\ \psi_{\pi,t} \\ \psi_{x,t} \\ \rho_t \end{bmatrix} + \varepsilon_t \tag{8}$$

Following the empirical literature,[15] we assume that the time-varying Taylor rule parameters adhere to a random walk process. Hence the transition equation of the state-space model[16] can be written in the following matrix representation:

$$\begin{bmatrix} \psi_{r^{nat},t} \\ \psi_{\pi,t} \\ \psi_{x,t} \\ \rho_t \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \psi_{r^{nat},t-1} \\ \psi_{\pi,t-1} \\ \psi_{x,t-1} \\ \rho_{t-1} \end{bmatrix} + \begin{bmatrix} u_{1,t} \\ u_{2,t} \\ u_{3,t} \\ u_{4,t} \end{bmatrix} \tag{9}$$

To run the Kalman filter, we must select the initial values of the parameters. In the absence of prior relevant information about the parameters, this selection is quite problematic (Zhao and Huang, 2020). Since none of the analyzed countries had targeted inflation before our data range, we can't employ the approach of isolating a portion of the dataset solely for pre-estimation, as done in studies like Cogley and Sargent (2001) and Kim and Nelson (2006). Instead, our strategy involves initializing the Kalman filter with parameter values that yield the lowest prediction error, as detailed in Table 5 in the Appendix.

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[13] Note that the assumption of a shorter horizon is a common practice in empirical literature, as demonstrated by studies like Clarida et al. (2000), Boivin (2005), Kim and Nelson (2006), and Baxa et al. (2014).

[14] See the Appendix, equation (A.13).

[15] See, for example, Horváth (2006), Yüksel et al. (2013) or Baxa et al. (2014).

[16] See the Appendix, equation (A.14).

Additionally, we exclusively utilize disposable data from the pre-inflation targeting era for initialization purposes. This approach minimizes the potential adverse impact of incorrect initialization values on the estimation process.[17]

Note that the estimated parameters then can be recomputed in order to gain the original parameters of the weight on the inflation gap and output gap and the policy-neutral rate described by equation (6).[18]

4. Results and Discussion

In this section, we present the results of the estimation process outlined in section 3 and engage in a discussion of the findings.

4.1. The Czech Republic

Our findings are illustrated in Figure 4 below. The results indicate that the Czech National Bank's response varies over time, with all monetary policy parameters significant at the 5% level, except for the weight on the output gap ($\xi_{x,t}$).

The declining trend of the policy-neutral rate (r_t^{nat}) is primarily driven by decreasing inflation trends (see Figure 2). This demonstrates that in the money market, the equilibrium interest rate at which money demand meets supply has decreased to lower levels. This trend aligns with Horváth (2006; 2009), who estimated a decrease in the policy-neutral rate from around 5% in 2002 to approximately 2% in 2005.[19] However, our results exhibit higher volatility in the policy-neutral rate compared to Horváth's estimates.

Regarding the smoothing parameter (ρ_t), our estimates appear reasonable in light of critiques of excessive policy inertia by Rudebusch (2005). The results suggest that the Czech National Bank exhibited a significant preference for interest rate smoothing, which decreased after the Great Recession, only to rise again when foreign exchange interventions ceased in 2017. The volatility at the end of the sample cannot be reliably interpreted due to an increase of model's prediction

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[17] Keep in mind that an inaccurate initialization primarily affects the initial periods of estimation, yet the estimator remains consistent and the filter's performance improves with more observations.

[18] The recomputation is performed as follows: $r_t^{nat} = \dfrac{\psi_{r^{nat},t}}{(1-\rho_t)}$, $\xi_{\pi,t} = \dfrac{\psi_{\pi,t}}{(1-\rho_t)}$, $\xi_{x,t} = \dfrac{\psi_{x,t}}{(1-\rho_t)}$. Notably, substantial values of the smoothing parameter ρ_t possess the potential to lead to the model's explosion. Therefore, to ensure the stability of the filter, an upper constraint must be imposed on the smoothing parameter.

[19] In Horváth (2009), the analyzed data range from Horváth (2006) is extended to 2001 – 2006 and other Taylor rule specifications are analyzed.

error shown in Figure 5, which is caused by Covid-19 pandemic and Russian-Ukrainian war shocks. This shows that our model is unable to explain these shocks and performs poor after 2020. Similarly, the model has a problem to explain the Great Recession. Our estimates of monetary policy inertia are lower than Horváth's (2006) for the period of 2002 – 2005, and generally consistent with Horváth's (2009) for the period of 2001 – 2006.[20] For the period of 2004 – 2015, our estimated values are lower than in Feldkircher et al. (2016).[21] In Mackiewicz-Łyziak's (2016) time-invariable Taylor rule specification with the forward looking inflation gap assuming central bank's perfect foresight, which is close to ours, she estimated the parameter to be 0.78 for the active and 0.98 for the passive monetary policy.[22] Our results of the smoothing parameter vary around 0.5, hence, we estimated lower inertia.

The weight parameter on the inflation gap ($\xi_{\pi,t}$) encapsulates the magnitude of the Central Bank's response to the inflation gap at the monetary policy horizon. An intriguing pattern emerges from our observations regarding this parameter. Notably, the Czech National Bank swiftly curtailed its response in 2003. This was facilitated by the rapid anchoring of inflation expectations around the targeted level and the bolstering of the Bank's credibility. In response to the overall price escalation in 2007, the Czech National Bank had to address the surge in inflation expectations spurred by swift inflation (see Figure 2).[23] Following this, a period of stability and reduced response to the inflation gap ensued from 2010 onwards, except for the post-pandemic period when inflation expectations experienced a de-anchoring effect. The restrained response during this phase was predominantly due to the constraint posed by the zero lower bound, impeding the ability to implement lower interest rates.[24] Our estimates of this parameter are close to the values estimated in the related empirical literature. For the period of 2002 – 2005 in Horváth

[20] Horváth (2006) estimates the smoothing parameter to be above 0.9. In a subsequent paper, Horváth (2009) introduces expectations into the model, resulting in a decrease of the smoothing parameter to 0.4. It is important to note that in both of these works, the author does not permit the smoothing parameter to be time-varying in his specifications.

[21] On the other hand, Feldkircher et al. (2016) treat this parameter as time-invariant within their Taylor rule specification. The divergence in results could potentially stem from the utilization of the time series of the central bank policy rate, which is subject to censorship and may not adequately capture genuine changes in the parameters (Podpiera, 2008).

[22] She employs the Markov switching method to discern between two distinct monetary policy regimes – active and passive. Her results for the active regime shall be subject to Rudebusch's critique of high monetary policy inertia. Similarly to Feldkircher et al. (2016), the utilization of censored policy rates (Podpiera, 2008) may potentially introduce bias into her results.

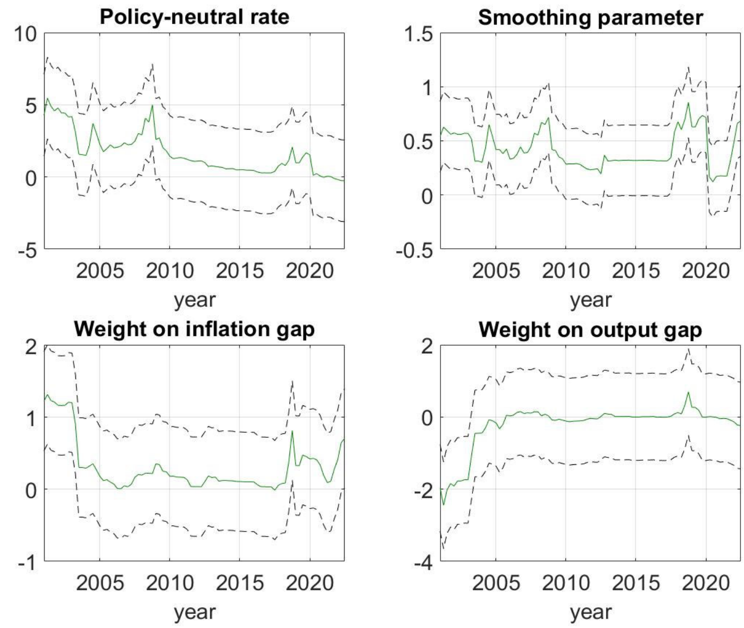
[23] These observations resonate with the existing literature that examines inflation expectations in the Czech Republic, such as Horvath et al. (2008) and Sousa and Yetman (2016).

[24] Worth noting is the period from 2013 to 2017 when the Czech National Bank implemented unconventional monetary policy through foreign exchange interventions.

(2006) and 2001 – 2006 in Horváth (2009), the estimates of the inflation gap coefficient are significant only for a few specifications that differ from ours; nevertheless, the estimates vary from 0.59 – 0.69, which is very close to the average value of our estimate for the same period. Feldkircher et al. (2016) use a similar specification and find a gradual decrease in the parameter from the value of 0.2 in the second half of 2004 to –0.013 in the first half of 2015. Although the values of our estimates are close, unlike Feldkircher et al. (2016), we do not observe a smooth and gradual decrease of the parameter in the 2004 – 2015 period.[25]

F i g u r e 4

Evolution of the Taylor Rule Parameters of the Czech National Bank



Note: The results exhibit significant time-variability in the Czech National Bank's monetary policy parameters. All parameters are significant at the 5% level except for the weight on the output gap, shown with 95% confidence bands (dashed lines) in each graph.

Source: Own calculations.

Our estimates of the weight parameter on the output gap ($\xi_{x,t}$) – which reflects the central bank's responsiveness to business cycles – are insignificant. This outcome aligns with the constitutional mandate set forth by Article 98 of the Constitution, emphasizing the Czech National Bank's primary focus on ensuring price

[25] This discrepancy in results may stem from the utilization of policy rates in the regression equation. As noted, the policy rates are censored and thus do not capture the real dynamics of the parameters (Podpiera, 2008).

stability rather than output stability. This finding is also in line with the official statement of the Czech National Bank regarding its reaction function within the g3 and g3+ model, as outlined in Czech National Bank (2009). Furthermore, this result harmonizes with the related empirical literature. Both Horváth (2006; 2009) and Feldkircher et al. (2016) arrive at similar conclusions, either detecting insignificance across various specifications (Horváth) or estimating the parameter as significant but extremely close to 0 (Feldkircher et al.).

The next Figure 5 illustrates the prediction error of the estimated time-varying parameters for the Czech National Bank. In contrast to the estimations for the Hungarian and Polish central banks' Taylor rules, the Kalman filter is effectively initiated from the beginning of the analyzed period. However, like the Hungarian and Polish cases, the filter's predictions deteriorate during times of crisis. Particularly, the results following the Covid-19 pandemic and the Russian-Ukrainian conflict are deemed unreliable due to their impact on the model's accuracy.

F i g u r e 5

Prediction Error of Estimation of the Czech National Bank's Taylor Rule Parameters

Note: The graph shows the prediction error of the estimated time-varying parameters for the Czech National Bank.
Source: Own calculations.

4.2. Hungary

Similar to the Czech National Bank's case, we observe time-variability in the monetary policy parameters of the Hungarian National Bank, as shown in Figure 6. Unlike in the Czech case, all observed parameters are significant at a 5% significance level, including the weight on the output gap ($\xi_{x,t}$).

In contrast to the Czech case, higher volatility is evident, particularly in the weight on the output gap ($\xi_{x,t}$) and the weight on the inflation gap ($\xi_{\pi,t}$). Note that this volatility might be partly attributed to the substantial prediction error observed at the beginning of the analyzed period (see Figure 7). Such time-variability makes it challenging for households and firms to anticipate the Central Bank's reaction to inflation levels, leading to unanchored inflation expectations. Unstable inflation expectations can exacerbate inflation instability itself, as discussed in Mandler (2007). Our findings align with empirical literature discussing unanchored inflation expectations in the Hungarian economy, as seen in Gábriel (2010), Gábriel, Rariga, and Várhegyi (2014). Yilmazkuday (2008) even concludes that the Hungarian National Bank did not genuinely implement inflation targeting. Despite this, we observe that during this period, the Hungarian National Bank decreased its response to the inflation gap. The Hungarian case underscores the lesson that reducing the response to the inflation gap while facing unanchored inflation expectations can lead to heightened inflation and its volatility in the economy (see Figure 2).

The policy-neutral rate's (r_t^{nat}) evolution is primarily influenced by the inflation level (see Figure 2) and the equilibrium interest rate in the money market, where money demand meets money supply. In comparison with the Czech results, the policy-neutral rate is notably higher and decreases to lower values much later, around 2016. This pattern aligns with the unanchored inflation expectations of economic agents who reacted in the money market to unexpected shifts in inflation and other key variables through abrupt changes. After the volatile period ending in 2012, the policy-neutral rate declined to relatively lower levels, suggesting that in the money market, loanable funds became available at a significantly reduced interest rate. Our policy-neutral rate estimates align with those provided by Baksa et al. (2013), who analyzed the period of 2003 Q1 – 2012 Q1.[26]

The smoothing parameter (ρ_t) exhibits a decreasing trend over the observed period. The estimated values are sensible in light of Rudebusch's (2005) critique of high monetary policy inertia. The parameter's evolution indicates that the Hungarian National Bank prefers gradual shifts in policy rates; however, this preference is inconsistent and often varies from one period to another. This interpretation may, however, be influenced by the relatively large prediction error in our estimates. A sudden increase in the smoothing parameter in 2007 suggests that the Hungarian National Bank exempted temporary inflationary shocks and aimed to avoid reacting aggressively to the rapid inflation increase (see Figure 2), opting instead for a higher preference for interest rate smoothing. Conversely, in 2015, the Bank appeared more inclined to prevent deflation, even if it meant less smooth interest rate adjustments (see Figure 1). Unlike Feldkircher et al. (2016), who estimated the time-invariable smoothing parameter slightly higher than in the Czech case during 2004 – 2015, we found that the preference was quite similar during

[26] Baksa et al. (2013) estimated real policy-neutral rate. To compare our values with their estimates, one must deduct the inflation from the policy-neutral rate.

this period. This implies that both central banks had comparable interest rate smoothing preferences. Mackiewicz-Łyziak (2016) estimated the constant Taylor rule for 1998 – 2014, finding a smoothing parameter above 0.9, which is subject to Rudebusch's (2005) critique. Despite the unreliability of estimates post the Covid-19 pandemic due to high filter prediction errors, the results suggest the Bank adopted a more hawkish stance in response to strong inflationary pressures.

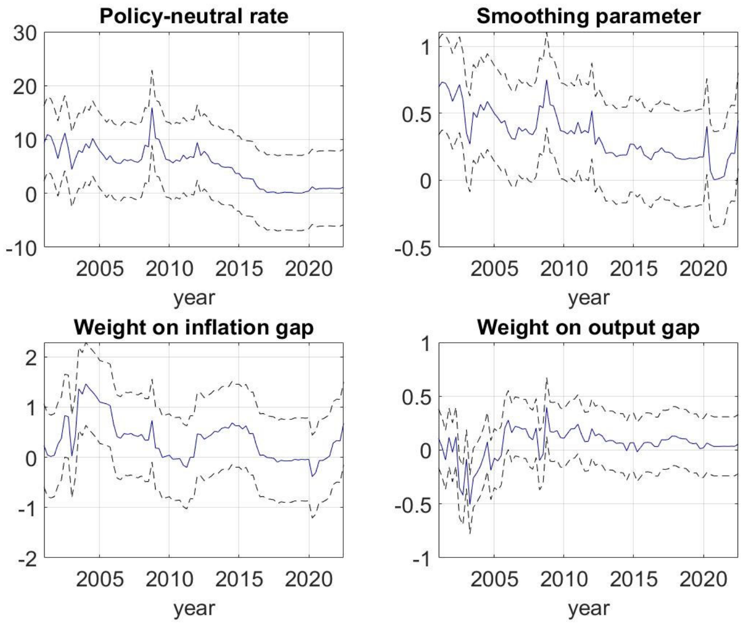
The changes in the Hungarian National Bank's response to the inflation gap are illustrated through the weight parameter on the inflation gap ($\xi_{\pi,t}$). The response exhibited considerable volatility during 2001 – 2004. The behavior of the Hungarian National Bank in this period proved challenging to capture using the selected Taylor rule (see Figure 7), implying a less strict adherence to the rule. Despite unanchored inflation expectations, the Bank decreased its response to the inflation gap between 2005 and 2011. This behavior led to inflation persistently hovering above the target during most of the quarters from 2005 to 2013, eventually converging back to the target when the Bank enhanced its weight on the inflation gap in response to mounting deflationary pressures between 2011 and 2014 (see Figure 2). Post-2014, the Bank's response to the inflation gap waned as interest rates neared the zero lower bound, preventing further decreases (see Figure 1). To counter robust inflationary pressures and re-anchor inflation expectations after the Covid-19 pandemic, the Bank needed to heighten its response to the inflation gap. The estimated trajectory of the weight parameter on the inflation gap ($\xi_{\pi,t}$) aligns with related empirical literature in the following way. Feldkircher et al. (2016) use a similar specification and identify a similar decreasing trend for the parameter during 2004 – 2015. Abaligeti et al. (2018) adopt a different specification without a smoothing parameter, yet their results mirror the parameter's dynamics, particularly during 2000 – 2014. Conversely, Mackiewicz-Łyziak (2016) finds a near-zero response to the inflation gap for a similar specification, contrasting our findings.

In contrast to the results of Feldkircher et al. (2016) and Abaligeti et al. (2018), our estimates of the weight on the output gap ($\xi_{x,t}$) – representing the Hungarian National Bank's response to business cycles – differ. While these studies found this parameter insignificant, Mackiewicz-Łyziak (2016) yielded significance in both active and passive monetary policy regimes with a similar specification. This disparity might stem from variations in output gap specifications. Unlike Feldkircher et al. (2016), Mackiewicz-Łyziak (2016), and Abaligeti et al. (2018), we employ the forward-looking output gap, following Clarida et al. (2000), Kim and Nelson (2006), Horváth (2006; 2009), and Baxa et al. (2014), offering better alignment with central banks' targeting of key variables at the monetary policy horizon.

The significance of this parameter underscores the Hungarian National Bank's active role in maintaining the economy near its potential output, although this preference has gradually diminished to very low values, signifying a shift towards the primary objective of price stability maintenance. Note that early periods of the range are challenging to interpret due to high prediction error in the estimates (see Figure 7). Our findings of a significant output gap parameter accord with the core MPM model of the Hungarian National Bank described in Szilágyi et al. (2013), where a positive value for the weight on the output gap is calibrated. They also echo the observations of Yilmazkuday (2008), who posits that the Hungarian National Bank did not fully implement inflation targeting.

F i g u r e 6

Evolution of the Taylor Rule Parameters of the Hungarian National Bank



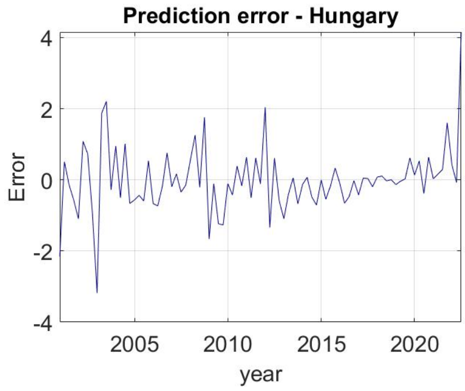
Note: The estimation results reveal notable time-variability in the monetary policy parameters of the Hungarian National Bank. All parameters are statistically significant at the 5% level. Each graph includes the 95% confidence bands, represented by dashed lines.

Source: Own calculations.

Displayed in Figure 7 are the prediction errors of the estimated time-varying parameters for Hungary. Similar to the Polish case, the Kalman filter takes time to reduce prediction errors, and its performance diminishes after the 2007 crisis and the Russian-Ukrainian conflict. The elevated prediction error implies that the behavior of the Hungarian Central Bank is challenging to capture using the standard Taylor rule specification.

F i g u r e 7

Prediction Error of Estimation of the Hungarian National Bank's Taylor Rule Parameters



Note: The graph shows the prediction error of the estimated time-varying parameters for Hungary.

Source: Own calculations.

4.3. Poland

Figure 8 illustrates the outcomes of estimating the time-varying monetary policy parameters of the National Bank of Poland. Mirroring the findings for the Czech Republic and Hungary, the parameters of the National Bank of Poland exhibit significant time variation. Notably, all parameters, including the weight on the output gap ($\xi_{x,t}$), hold significance at the 5% level.

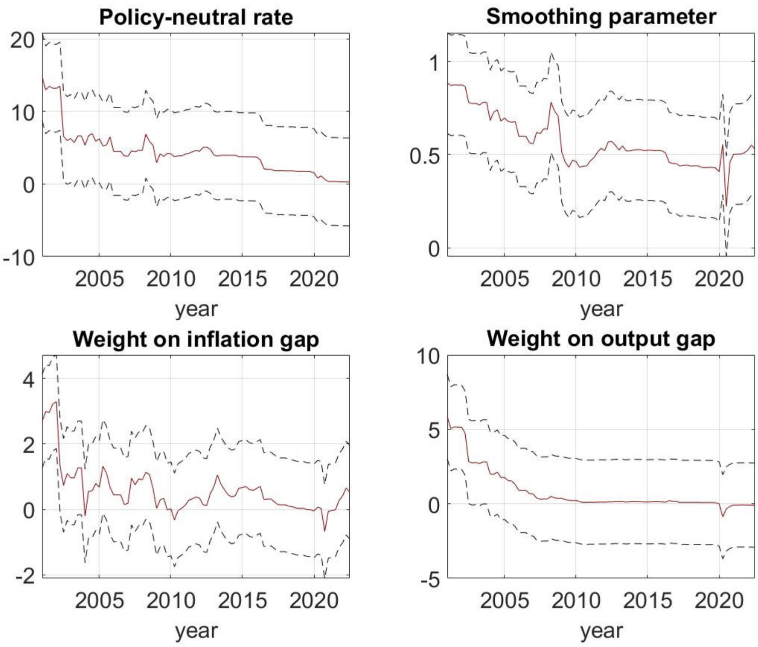
The gradual decline of the policy-neutral rate (r_t^{nat}) from above 10 % level in 2001 to low levels around 0 % is driven partly by the decrease in inflation and by the decrease in the interest rate for which the loanable funds were traded on the money market. The introduction of inflation targeting contributed to the drop in inflation. Unlike the Czech and Hungarian cases, Poland's policy-neutral rate didn't descend to near-zero levels as rapidly. As a result, the Bank's policy rate reductions were not as forceful as those of the Czech and Hungarian counterparts. Refer to Figure 1 for the dynamics of Poland's short-term interest rate, which only approached zero levels by 2019. Brzoza-Brzezina (2006) estimated the real policy-neutral rate for 1998 – 2004 and found substantial volatility, aligning with our estimates. Our findings also concur with Michalek et al. (2012), who estimated the real policy-neutral rate for 1998 – 2011.

Similarly to our estimates for the Czech Republic and Hungary, the smoothing parameter (ρ_t) for Poland demonstrates a decreasing trend. The values of the smoothing parameter are in line with Rudebusch's (2005) critique of high monetary policy inertia. This dynamics indicates that the National Bank of Poland is adjusting policy rates with less smoothness than in the past. A disruption in this trend occurred in 2007 when the Bank responded more assertively to inflationary pressures. The estimated average of this parameter is markedly lower than Feldkircher et al.'s (2016) time-invariable estimate, which nearly reaches 1. Such a value contradicts Rudebusch's (2005) criticism of unrealistically high smoothing preferences in empirical literature. Mackiewicz-Łyziak (2016) estimates this parameter as 0.87 for passive and 0.85 for active monetary policy regimes in a related Taylor rule specification. Contrary to our predictions, her estimates are higher, potentially influenced by the usage of censored policy rates as endogenous variables, a practice critiqued by Podpiera (2008). Mackiewicz-Łyziak (2017) finds that elevated smoothing parameter values only apply to the periods of 1998 Q2 – 2003 Q4 and 2004 Q1 2008 Q3. In her third sample split 2008 Q4 – 2014 Q2, her estimate of the smoothing parameter in the related specification is 0.46, corroborating our observation of decreasing monetary policy inertia in Poland to around 0.5 values.

The weight on the inflation gap ($\xi_{\pi,t}$) in Poland mirrors inflation expectations (see Lyziak (2014), page 10 and Łyziak et al. (2016), page 37). When inflation expectations rose, the National Bank of Poland increased its interest rate response to the inflation gap and *vice versa*. In the face of significant deflationary pressures in 2013 (see Figure 2), the Bank responded with more pronounced policy rate changes to stimulate inflation (see Figure 1). As deflationary pressures waned in 2016 and inflation expectations returned to the target level, the Bank moderated its reaction to the inflation gap, leading to a decline in the weight on the inflation gap. The post-pandemic estimates are unreliable due to high prediction error (see Figure 9). Our estimate of the weight on the inflation gap is slightly lower than Feldkircher et al.'s (2016) but shares a similar decreasing trend. Mackiewicz-Łyziak (2016), estimating a constant Taylor rule, also suggests a lower parameter value compared to Feldkircher et al. (2016). Mackiewicz-Łyziak (2017) reaffirms these findings through sample split analysis, yielding values consistent with her prior study.

The weight parameter on the output gap ($\xi_{x,t}$) exhibits significance at the 5% level; however, this significance primarily stems from the initial sample periods. A t-test performed on the predicted values of the weight on the output gap, excluding the 2001 – 2007 period during which the National Bank of Poland aimed for stable economic growth, reveals its insignificance afterward as it converges to nearly zero values. These results suggest a commitment to pure inflation targeting by the National Bank of Poland post-2007. Although we are not aware of any official statement from the National Bank of Poland that specifies its reaction to the output gap, Brzoza-Brzezina, Kotłowski and Miśkowiec (2013) offer valuable empirical insights when analyzing the Bank's forecasts. Their study, covering 2004 – 2011, concludes that the weight on the output gap is insignificant. In contrast, Feldkircher et al. (2016) find a significant parameter that evolves around 0.17. A related specification by Mackiewicz-Łyziak (2016) reports a significant parameter at approximately 0.8 for the active regime and 0.14 for the passive regime. Mackiewicz-Łyziak (2017) further employs sample splits analysis, suggesting a decreased response between the 1998 – 2003 and 2004 – 2008 subsamples. Comparing the outcomes of Feldkircher et al. (2016) and Mackiewicz-Łyziak (2016; 2017) requires considering that our use of a forward-looking output gap, following Clarida et al. (2000), Kim and Nelson (2006), Horváth (2006; 2009), and Baxa et al. (2014), might contribute to divergent results from the mentioned authors.[27]

F i g u r e 8
Evolution of the Taylor Rule Parameters of the National Bank of Poland



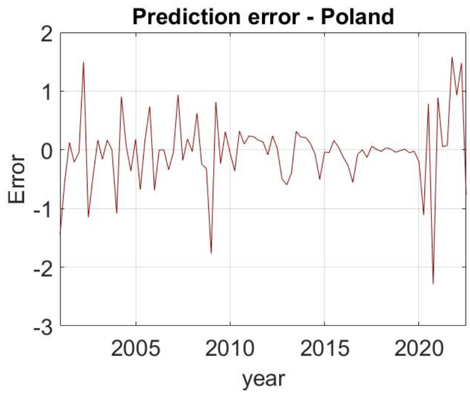
Note: The estimation outcomes reveal substantial time-variation in the monetary policy parameters of the Hungarian National Bank. All parameters exhibit significance at the 5% level. The accompanying graphs feature 95% confidence bands represented by dashed lines.

Source: Own calculations.

[27] Adopting a backward-looking output gap assumption does not align with the actual efficacy of monetary policy. The monetary policy horizon at which the transmission mechanism significantly impacts the economy spans 4 – 6 quarters. Employing a shorter horizon could introduce bias into estimates. However, it might also mitigate prediction errors, as elaborated in Baxa et al. (2014).

Displayed in Figure 9 is the prediction error of the estimated time-varying parameters for Poland. It's important to note that the post-pandemic results are marred by considerable prediction error, rendering them challenging to interpret.

Figure 9
Prediction Error of Estimation of the National Bank of Poland's Taylor Rule Parameters



Note: The graph shows the prediction error of the estimated time-varying parameters for Poland.
Source: Own calculations.

Concluding Remarks

This study has delved into the evolution of time-variable Taylor rule parameters for the Czech National Bank, the Hungarian National Bank, and the National Bank of Poland. Typically, sources of parameter time-variability stem from changes in the preferences of the central bank's governing body as discussed in Judd et al. (1998), Clarida et al. (2000), Orphanides (2004), and Kim and Nelson (2006); in the changes to the central bank's prediction models; in the changes to the transmission mechanism in the economy, as explored by Baxa et al. (2014), Benati (2008), and Zhang et al. (2008); and also in the utilization of unconventional monetary policy and exemptions to temporary shocks. Our analysis is motivated by the recent Covid-19 pandemic and the energy crisis caused by the Russian-Ukrainian war, both of which could potentially prompt adjustments in interest rate responses to targeted variables.

Monetary policy consistency remains a cornerstone for stable economic growth (Kydland and Prescott, 1977), as it enables agents to anticipate policy changes when forming inflation expectations and making critical economic decisions. While the variability of monetary policy parameters has been extensively explored in literature (see section 1), analysis specific to the Central and Eastern European

(CEE) region remains limited. This paper confronts estimation results from related empirical literature in which the authors use different estimators and methodology. Moreover, we provide more current results, as the last analyses of time-varying Taylor rules on CEE region were performed in 2016 (see Feldkircher et al., 2016; Mackiewicz-Łyziak, 2016).

This paper adheres to the mainstream methodology, employing the Kalman (1960) filter and full information maximum likelihood estimation to derive time-varying Taylor rules that capture shifts in interest rate responses. Our focus is on the data range from 1996 Q1 to 2020 Q4, encompassing the post-inflation targeting period in all three countries. The pre-inflation targeting phase is used solely for initialization purposes, thus limiting the analyzed data range to 2001 Q2 – 2020 Q4.

Firstly, our analysis reveals a consistent trend: the policy-neutral rate declined to nearly zero across all three countries. This trend emerged due to decreases in both inflation and the long-term real equilibrium interest rate, which aligns money demand with money supply in the money market.[28]

Secondly, the weight parameter on the inflation gap, reflecting the central bank's reaction to this gap, also exhibited a decline towards near-zero values, followed by an increase during the current high inflation period. While the Czech Republic and Poland justified the decrease in the inflation gap due to well-anchored inflation expectations that reinforce the inflation target, a similar decrease in Hungary in an environment of unanchored inflation expectations explains its high inflation volatility. These findings suggest that the central bank should not be weaker in its response to the inflation gap when the inflation expectations in the economy are not anchored. Additionally, our findings indicate that central banks intensify their response to the inflation gap in the face of deflationary pressures at the policy horizon, underscoring their aversion to deflation. This aligns with their positive inflation targets. In 2013, all three central banks confronted deflationary pressures and swiftly reduced their policy rates in response.

Thirdly, the interest rate smoothing preference of all three analyzed central banks exhibited a downward trend, reflecting a promotion of less smooth policy rate changes. This suggests a shifting emphasis towards embracing more immediate adjustments in central banking.

Lastly, in terms of responding to business cycles as depicted by the weight on the output gap, only Hungary demonstrated a significant response throughout the whole analyzed data range. This indicates that the Czech National Bank strictly adheres to its primary constitutional duty, while the National Bank of Poland maintains its focus on upholding price stability based on legal obligations.

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[28] The lower real equilibrium interest rate is caused mainly by the long-term surplus of liquidity in this market.

The primary contribution of this paper lies in illustrating the evolving monetary policy response of the Czech National Bank, the Hungarian National Bank, and the National Bank of Poland. These dynamic changes shed light on the performance of these central banks' monetary policies, underscored by their public credibility reflected in inflation expectations. This analysis emphasizes that central banks still retain a degree of discretion in their monetary policy parameters.

Based on our findings, we recommend that central banks monitor their credibility through vigilant observation of inflation expectations. Additionally, we suggest that central banks should refrain from reducing their policy rate response to the inflation gap when confronted with unanchored inflation expectations, as this could exacerbate instability. Following Kydland's and Prescott's (1977) critique of unexpected discretion, central banks should communicate any unexpected changes in preferences well in advance to ensure predictability.

Further research avenues could encompass extending this analysis to other Central and Eastern European countries with independent monetary policies. Exploring the benefits of employing the Kalman smoother for improved estimation accuracy is another potential area of study.[29]

Additionally, considering a more realistic policy horizon of 4 – 6 periods, as opposed to the current 2-period horizon, could yield insightful results, even though it may lead to a rise in prediction error.

The main limitation of this research lies in the initialization of the Kalman filter. Due to no prior information being available about the parameters' values and short dataset, we decided to calibrate the parameters and use those values that minimize the prediction error. Furthermore, we excluded initial periods from the analysis, utilizing them solely for filter initialization.[30]

Another limitation is the dataset concluding during a high inflation period. The Kalman filter's performance would improve with data that reflects a return to long-term averages.

This data, however, was not available at the time this research was conducted. We encourage researchers to revisit this study when adequate observations beyond the present high inflation period become available.

[29] Theoretically, the Kalman smoother is expected to yield superior performance owing to its utilization of backward prediction, thereby leveraging a greater amount of information. Nonetheless, certain studies present evidence that the filter, in fact, exhibits enhanced efficacy, as demonstrated by Evensen and Van Leeuwen (2000).

[30] The first periods in the estimation process are the most sesitive to the inicialization of the filter. Therefore, we believe, that omitting 1996 The initial periods within the estimation procedure are particularly susceptible to the effects of filter initialization. Consequently, we hold the perspective that the exclusion of 1996 Q1 – 2001 Q1 should significantly contribute to bolstering the credibility of the obtained outcomes.

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Appendix

A.1: Inflation Targets

Table 1

Inflation Targeting of the CNB, NBP and HNB (in %)

	Inflation target:		
Year	**CNB**	**NBP**	**HNB**
1998	5.5 – 6.5		
1999	4 – 5	*6.6 – 7.8	
2000	3.5 – 5.5	5.4 – 6.8	
2001	3 – 5	6 – 8	†7
2002	3 – 5	6 – 8	4.5
2003	3 – 5	#below 4	3.5
2004	3 – 5	2.5	3.5
2005	2 – 4	2.5	4
2006	3	2.5	3.5
2007	3	2.5	3
2008	3	2.5	3
2009	3	2.5	3
2010 – ?	2	2.5	3

Note: *Initially set to 8 – 8.5%. #For simplicity assumed to be 4%. †Starting from June 2001.

Source: Czech National Bank, Hungarian National Bank, National Bank of Poland websites.

The Czech National Bank adopted inflation targeting in 1998, initially employing an announced interval as its targeted inflation level. From 2006 onwards, the inflation target has remained set as constant. Similarly, the National Bank of Poland embraced inflation targeting in 1999, initially targeting an interval for inflation. Since 2004, it has aimed to achieve an inflation rate of 2.5%. In the case of the Hungarian National Bank, inflation targeting was introduced in June 2001, with a subsequent shift to targeting a fixed inflation level. From 2007, its inflation target has been set at 3%.

To calculate an inflation gap, we consider a constant inflation target for each year. When a central bank specifies an interval as its inflation target, we take the mean value of the interval bounds as the inflation target. Notably, in 1999, the National Bank of Poland modified its initial target from 8 – 8.5% to 6.6 – 7.8%; here, we simplify by assuming the latter target as the inflation level. Similarly, in 2003, when the National Bank of Poland aimed to maintain inflation below 4%, we treat it as if the target were set at 4%. Additionally, the Hungarian National Bank introduced its inflation target in June 2001. For simplicity, we treat this target as in effect since January 2001.

Lastly, it's important to highlight that the period before the adoption of inflation targeting is solely utilized for initializing the Kalman filter. Given the higher inflation levels prevalent in the pre-inflation targeting era, our aim is to minimize prediction errors during initialization. To achieve this, we employ the earliest inflation target announced by a central bank and treat it as if it applied throughout the pre-inflation targeting period. This approach enhances the precision of estimates during the initial phase of the observed inflation targeting period, as the Kalman filter receives a quicker boost.

A.2: Hodrick-Prescott Filter

The Hodrick-Prescott filter assumes that a time series can be decomposed into a growth (trend) component g_t and a cyclical component c_t

$$y_t = g_t + c_t \tag{A.10}$$

The filter estimates the growth component g_t by solving the following optimization problem:

$$\min_{\{g_t\}_{t-1}^{T}} \sum\nolimits_{t-1}^{T} c_t^2 + \lambda \sum\nolimits_{t-1}^{T} \left[\left(g_t - g_{t-1} \right) - \left(g_{t-1} - g_{t-2} \right) \right]^2 \tag{A.11}$$

The parameter λ penalizes the variability in the growth component g_t and is specified as follows

$$\lambda = f^a \times 1600 \tag{A.12}$$

where f differs according to the data frequency. For the quarterly data $f = 1$. In the literature, there is no consensus about the exponent *a*. Hodrick and Prescott (1997) suggest $a = 2$, whereas Ravn and Uhlig (2002) shows on the US data that $a = 4$ is more precise. Nevertheless, as we use quarterly data, we do not have to examine these two concepts further.

The Hodrick-Prescott filter has faced criticism, notably from J. D. Hamilton (2018), who argues that it can produce spurious dynamic relationships that do not align with the true data generating process. Hamilton proposed an alternative H filter. In response, Hodrick (2020) compared the estimates of the H filter, the HP filter, and other filtering techniques. He found that while the H filter performs better in simple time-series models like random walk or ARIMA(2, 1, 2), the Hodrick-Prescott filter is superior for complex time-series models. Given the intricate nature of output dynamics, we deem it more appropriate to use the Hodrick-Prescott filter.

A.3: Descriptive Statistics

The following Table 2, Table 3 and Table 4 provide descriptive statistics of the data for each country.

T a b l e 2

Descriptive Statistics of the Czech Data

Statistic	Mean	St. Dev.	Min	Pctl(25)	Pctl(75)	Max
Interest rate	3.615	4.240	0.280	0.883	4.485	19.673
Lagged interest rate	3.651	4.287	0.280	0.883	4.485	19.673
Inflation gap	0.575	3.530	−4.387	−1.421	1.137	15.578
Output gap	–0.025	2.053	−9.335	−1.598	1.035	4.396

Note: This table shows the descriptive statistics of the Czech data for the period of 1996 Q1 – 2022 Q3.
Source: Own calculations.

T a b l e 3

Descriptive Statistics of the Hungarian Data

Statistic	Mean	St. Dev.	Min	Pctl(25)	Pctl(75)	Max
Interest rate	7.744	6.222	0.036	1.944	10.883	26.367
Lagged interest rate	7.912	6.574	0.036	1.944	10.883	30.067
Inflation gap	2.337	4.605	−4.059	−0.275	3.413	22.409
Output gap	−0.024	2.146	−13.937	−0.749	0.985	3.933

Note: This table shows the descriptive statistics of the Hungarian data for the period of 1996 Q1 – 2022 Q3.
Source: Own calculations.

T a b l e 4

Descriptive Statistics of the Polish Data

Statistic	Mean	St. Dev.	Min	Pctl(25)	Pctl(75)	Max
Interest rate	7.185	6.852	0.210	1.800	7.680	24.310
Lagged interest rate	7.358	7.081	0.210	1.800	8.903	25.667
Inflation gap	0.819	4.017	−6.067	−1.417	1.967	14.883
Output gap	0.007	1.708	−8.598	−1.021	1.294	4.394

Note: This table shows the descriptive statistics of the Polish data for the period of 1996 Q1 – 2022 Q3.
Source: Own calculations.

A.4: Kalman Filter

The Kalman (1960) filter can be used to get the time series of unobserved variables or time-varying parameters. The filtering technique requires the model to be written in a state-space manner in order to specify the Kalman filter equations.

The filter needs to have specified the observation equation, which includes the observable variables in the following matrix representation

$$\boldsymbol{z}_t = \boldsymbol{H}\varrho_t + \boldsymbol{D}_t\boldsymbol{x}_t + \boldsymbol{v}_t, \ \forall t \in [1, 2, ..., T] \tag{A.13}$$

where $\boldsymbol{z}_t$ includes the observable endogenous variable values, ϱ_t is the vector of exogenous variables, $\boldsymbol{x}_t$ is the vector of unobservables and $\boldsymbol{v}_t$ is the vector of errors caused by statistical measurement mistakes. $\boldsymbol{D}_t$ and $\boldsymbol{H}$ are the matrices of parameters.

Further, the transition equation describes the evolution process of the unobservables:

$$\boldsymbol{x}_t = \boldsymbol{A}\boldsymbol{x}_{t-1} + \boldsymbol{u}_t, \quad \forall t \in [1, 2, ..., T] \tag{A.14}$$

where $\boldsymbol{x}_t$ is the vector of unobservable variables, $\boldsymbol{A}$ is the matrix of parameters and $\boldsymbol{u}_t$ is the vector of random errors.

In the first step, the Kalman filter predicts the vector of unobservable variables using the prediction equations

$$\boldsymbol{x}_{t+1|t} = \boldsymbol{A}\boldsymbol{x}_{t|t} \tag{A.15}$$

and

$$\boldsymbol{P}_{t+1|t} = \boldsymbol{A}\boldsymbol{P}_{t|t}\boldsymbol{A}' + \boldsymbol{Q} \tag{A.16}$$

where $\boldsymbol{x}_{t+1|t}$ is the prediction of unobservable variables computed as the projection of $\boldsymbol{x}_{t+1}$ on the matrix of disposable information in the last period $\boldsymbol{\Omega}_t$ (i.e. $\boldsymbol{x}_{t+1|t} = \hat{P}(\boldsymbol{x}_{t+1} | \boldsymbol{\Omega}_t)$). $\boldsymbol{P}_{t+1|t}$ is the prediction of the covariance matrix of $\boldsymbol{x}_{t+1|t}$ computed in the following way $\boldsymbol{P}_{t+1|t} = E\left[\left(\boldsymbol{x}_{t+1} - \boldsymbol{x}_{t+1|t}\right)\left(\boldsymbol{x}_{t+1} - \boldsymbol{x}_{t+1|t}\right)'\right]$. Finally, $\boldsymbol{Q}$ is the covariance matrix of the error terms $\boldsymbol{u}$.

The following steps include iteration of prediction equations (A.15) and (A.16). In each of these steps, the Kalman filter uses filtering equations:

$$\boldsymbol{x}_{t|t} = \boldsymbol{x}_{t|t-1} + \boldsymbol{P}_{t|t-1}\boldsymbol{D}_t'\left(\boldsymbol{D}_t\boldsymbol{P}_{t|t-1}\boldsymbol{D}_t' + \boldsymbol{R}\right)^{-1}\left(\boldsymbol{z}_t - \boldsymbol{z}_{t|t-1}\right) \tag{A.17}$$

and

$$\boldsymbol{P}_{t|t} = \boldsymbol{P}_{t|t-1} - \boldsymbol{P}_{t|t-1}\boldsymbol{D}_t'\left(\boldsymbol{D}_t\boldsymbol{P}_{t|t-1}\boldsymbol{D}_t' + \boldsymbol{R}\right)^{-1}\boldsymbol{D}_t\boldsymbol{P}_{t|t-1} \tag{A.18}$$

where $\boldsymbol{x}_{t|t}$ and $\boldsymbol{P}_{t|t}$ denote predictions of the unobservable variables and their covariance matrix made in time t using the disposable information set at period t computed as the following projection $\hat{P}(\boldsymbol{x}_t \mid \boldsymbol{z}_t, \varrho_t, \boldsymbol{\Omega}_{t-1})$. Covariance matrix $\boldsymbol{P}_{t|t}$ is obtained as follows $\boldsymbol{P}_{t|t} = E\left[\left(\boldsymbol{x}_t - \boldsymbol{x}_{t|t}\right)\left(\boldsymbol{x}_t - \boldsymbol{x}_{t|t}\right)'\right]$. Note that the prediction of $\boldsymbol{x}_{t|t}$ is improved by the difference between the actual values of observable variables $\boldsymbol{z}_t$ and its prediction of $\boldsymbol{z}_{t|t-1}$ based on the information set available at period $t-1$. $\boldsymbol{R}$ is the covariance matrix of the error terms $\boldsymbol{v}$.[31]

As the covariance matrices $\boldsymbol{Q} = \sigma_u^2 \boldsymbol{I}$ and $\boldsymbol{R} = \sigma_u^2$ are unknown, we use Fisher's (1922) full information maximum likelihood estimator which maximizes the following log-likelihood function:

$$\log L\left(\boldsymbol{\theta} | \boldsymbol{\Omega}_t\right) = -\frac{T \times k}{2} \log\left(2\pi\right) - \frac{1}{2} \sum_{t-1}^{T} \left[\log\left|F_{t|t-1}\right| + \left(\tilde{\boldsymbol{z}}_t' \boldsymbol{F}_{t|t-1}^{-1} \tilde{\boldsymbol{z}}_t\right)\right] \qquad \text{(A.19)}$$

where $\boldsymbol{\theta}$ is maximized according to likelihood, $\tilde{\boldsymbol{z}}_t = \boldsymbol{z}_t - \boldsymbol{z}_{t|t-1}$ and

$$\boldsymbol{F}_{t|t-1} = E_{t-1}\left\{\left[\boldsymbol{D}_t\left(\boldsymbol{x}_t - \boldsymbol{x}_{t|t-1}\right) + \boldsymbol{v}_t\right]\left[\boldsymbol{D}_t\left(\boldsymbol{x}_t - \boldsymbol{x}_{t|t-1}\right) + \boldsymbol{v}_t\right]'\right\}.$$[32]

A.5: Kalman Filter Initialization

In the following Table 5, we provide the calibrated initial values of the Kalman filter for all three analyzed central banks. Note that $\sigma_{\varepsilon,0}$ is the initial value for standard error of the error term ε_t described in the observation equation (8).

T a b l e 5

Kalman Filter Parameters' Initialization

Parameter	**Value**
$\sigma_{\varepsilon,0}$	1
r_0^{nat}	15
$\xi_{\pi,0}$	2
$\xi_{x,0}$	0.2
ρ_0	0.7

Note: The initial parameters of the Kalman filter are selected upon calibration.
Source: Own calculations.

[31] For details see Kalman (1960).

[32] For details see Fisher (1922).