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ARTICLES

# THE REACTION FUNCTION OF THREE CENTRAL BANKS OF VISEGRAD GROUP

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**Abstract**:

The aim of our paper is to formulate and empirically verify the simple backward looking econometric model of the monetary policy rule, which would be able to describe the development of monetary policy rate, namely only on the basis of statistically measured and at the given time available information. We focus on the Czech National Bank, the National Bank of Poland and the Magyar Nemzeti Bank in the period of January 1999 to April 2012. In the present paper we discuss some methodological problems associated with the *ex-post* empirical verification of the central bank's monetary policy rule. We construct an empirical model of the monetary policy rule, justify the choice and the inclusion of explanatory variables, analyse the statistical properties of time series, and verify the alternative forms of econometric models. Our analysis showed that the development of monetary policy rate in the reporting period can be explained by the past and present development of four explanatory variables: yearly inflation rate, exchange rate, ECB main refinancing rate and growth rate of M2. The annualized inflation rate proved to be statistically insignificant in the model. We find interesting that the statistical quality of the estimated model was further increased after a six-month lag of the annual inflation rate added to the model.

**Keywords**: repo rate, monetary policy rules, annual inflation rate, econometric model, cointegration, Visegrad Group.

**JEL Classification**: E43, E47, E52, C12, C22

## 1. Introduction

In a monetary policy system called inflation targeting the interest rate transmission mechanism is based on a short-term interest rate of central bank. From the aspect of commercial banks, investment funds and other portfolio investors, successful prediction

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of changes in the short-term interest rate of central bank plays a crucial role for the appropriate setting of the term structure of bond portfolio. From central bank's aspect, the final decision on a change in monetary policy rate is the completion of a complex multistage management process that is institutionalized by voting of the bank board members. Inflation targeting is a monetary policy system that accentuates maximum transparency. Significant elements of monetary policy transparency are also the openness and intelligibility of central bank's prognostic apparatus to the professional community. A macroeconomic analyst in a commercial bank would apparently consider such a situation as ideal when they have an opportunity for parallelly working with the same model as central bank. Unfortunately, macroeconomic analytical teams of medium-sized commercial banks are relatively little numerous in comparison with research and analytical departments of central banks. The use of large macroeconomic models of the forward looking type exceeds their abilities from the aspect of time capacity. Hence there arises a question whether the movement of monetary policy interest rates of central banks cannot be estimated on the basis of a "simple" econometric backward looking model with a small number of explanatory variables.

In the present paper the reaction function of the three central banks of the Visegrad Group with independent monetary policy is analysed: the Czech National Bank (CNB), the National Bank of Poland (NBP) and the Magyar Nemzeti Bank (MNB). As shown in Table 1, central banks have implemented the inflation targeting policy in all studied countries for more than a decade. All three central banks work in conditions of excess liquidity in the banking sector that is withdrawn through open market operations. The specific technical forms of open market operations are different at a glance but in all cases these are operations with short-term (one-week or two-week) monetary policy interest rate. All three banks use highly sophisticated (multi-equation) models that are based on the New-Keynesian approach. However, specific methodological approaches are different. The CNB uses a model of DSGE type while these are predictions of "non-conditional" nature that already take into account the assumed trajectory of interest rate development. Parameters are calibrated on the basis of expert estimation. The NBP works with a large macroeconometric model where an emphasis is laid on the econometric estimation of model parameters. Predictions are of "conditional" nature and assume a constant interest rate. Since 2011 the MNB has used a New-Keynesian macro model where the variables are defined in the gap form. Parameters of the model are calibrated on the basis of theory and empirical econometric analysis.

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Table 1
**Basic Characteristics of Monetary Policy and Prediction Models in the Analysed Countries**

| | Czech National Bank | National Bank of Poland | Magyar Nemzeti Bank |
|---|---|---|---|
| **Year of inflation targeting implementation** | 1998 | 1999 | 2001 |
| **Present inflation target** | CPI (2%) | CPI (2.5%, +/- 1 p.p.) | CPI (3%) |
| **Monetary policy interest rate** | CNB repo rate (two-week repo rate) | NBP reference rate (yield on 7-day NBP money market bills) | MNB base rate (two-week MNB bill rate) |
| **Name of prediction model** | G3 (since 2007) | NECMOD (since 2008) | MPM (since 2011) |
| **Type of model** | DSGE | Macroeconometric model (forward looking, error-correction form) | New-Keynesian macro model (parameters calibrated, forward looking, gap form) |
| **Type of prognosis** | Non-conditional forecast | Conditional forecast | Non-conditional forecast |
| **Prediction horizon** | 6 quarters | 12 quarters | 8 quarters |

Source: Beneš, Hlédik, Kumhof and Vávra (2005), Budnik, Kolasa, Hulej, Greszta, Murawski, Rot, Tarnicka and Rybaczyk (2008), Horváth, Kober and Szilagyi (2011)

The objective of the present paper is to formulate and empirically verify a "simple" econometric backward looking model of the monetary policy rule that would be able to describe the development of central bank's monetary policy interest rate only on the basis of statistically measured information available at a given time. For this purpose the panel cointegration analysis is used.

The paper consists of two basic parts. In Part 2 an overview of hitherto research that has been aimed at the countries of Central Europe is presented. Subsequently, some methodological problems connected with the *ex-post* empirical verification of central bank's monetary policy rule are discussed. Part 3 is an empirical one. It contains the construction of an econometric model of the monetary policy rule, reasons for the selection and inclusion of explanatory variables in this model, analysis of statistical properties of time series and econometric verification of alternative forms of the model. Finally, the knowledge of our research is summarized.

## 2. Overview of Hitherto Research and Some Methodological Considerations

Among the works in which Taylor's backward looking approach to the monetary policy rule (Taylor, 1993) or somewhat later formulated forward looking approach (Mehra, 1999) is applied to the countries of Central Europe, it is possible to cite these contributions. For small open economies, Mohanty and Klau (2004) introduced a change in the real exchange rate among explanatory variables in the Taylor rule. They worked with output gap and quantified the potential product by means of HP filter. They also tested a possibility of

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the existence of a non-linear monetary policy rule. Maria-Dolores (2005) concluded that parameters estimated in a model for the Taylor backward looking rule showed a higher statistical significance than parameters in the selected forward looking model where the future values of inflation and production gap were estimated from instrumental variables. Frommel and Schobert (2006) also reformulated the original Taylor rule for a forward looking one. They approximated the inflation forecast with time horizon $t$+12 by actual future inflation. Paez-Farrell (2007) summarized the research conducted until then and tested various alternatives of backward looking and forward looking monetary policy rules, linear and non-linear monetary policy rule, while exchange rate, unemployment rate and production gap in the first differences were included.

Horváth (2008) and Vašíček (2011) tested a possibility of the existence of forward looking non-linear monetary policy rule on an example of the Czech Republic and other countries. However, they used a three-month rate of the interbank market (3M PRIBOR), not the repo rate of central bank. The monetary policy rule for CNB repo rate using the methods of artificial intelligence (neural network) was studied by Kukal and Van Quang (2011). But the partial conclusions drawn from their introductory econometric analysis for CNB repo rate (linear model, eight explanatory variables) are pessimistic because parameters are statistically insignificant or they have signs in contradiction with formulated hypotheses.

### 2.1 Some Methodological Problems of Hitherto Research

#### *Choice of explained variable*

In the majority of research works central bank's monetary policy interest rate is not directly chosen as an explained variable, but it is the interest rate of interbank money market that expressed in the "old" terminology of transmission monetary mechanism. It plays a role of operative criterion. It is possible to demonstrate some pitfalls of this approach. The interbank market rates are influenced by central bank's monetary policy interest rate on the one hand, on the other hand they are affected by market expectations. This problem was analysed by Brada and Brůna (2004), who accentuated the influence of expectation concerning the future behaviour of central bank. *E.g.* growing inflation expectations in the market logically lead to an increase in interest rates in the interbank market. However, it does not automatically mean that central bank performs restrictive monetary policy. The aim of *ex-post* empirical verification of the monetary policy rule is not to test the behaviour of market interest rate in the interbank market but, on the contrary, the behaviour of central bank. Hence we are convinced that the monetary policy rule should be formulated directly for the monetary policy interest rate of central bank.[1]

1 There are another reasons, too. Namely, the liquidity stress on the interbank market, which was apparent during the recent financial turmoil. Also changes in the riskiness of uncollateralized debt and oligopolistic structure of the interbank markets exert strong influence on the interbank rates.

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### *Small open economy*

The Taylor monetary policy rule (Taylor, 1993) was verified empirically on the basis of U.S. economy facts in 1987–1992, *i.e.* in the case of a large closed economy. In a small open economy, the difference between inflation and inflation target and the production gap will not apparently be the only explanatory variable. This is the reason why in many works cited above the authors included also various forms of exchange rate in the monetary policy rule (nominal or real exchange rate, level data or first differences). We believe, that in the case of open economy and existence uncovered interest rate parity, it is necessary to test, besides exchange rate, the sensitivity of central bank's interest rate policy to foreign interest rates (Chinn, Meredith, 2004; Durčáková, Mandel, Tomšík, 2005). The relation to ECB repo rate seems logical in the studied countries.

### *Potential product and production gap*

The calculation of potential product as the starting point to determine production gap is one of the crucial empirical problems of the Taylor monetary policy rule. The authors (*e.g*. Mohanty, Klau, 2004 and Paez-Farrell, 2007) estimated the potential product by means of HP filter. The use of HP filter for the estimation of potential product development is based on an implicit assumption that aggregate demand adapts itself relatively quickly to aggregate supply development represented by the potential product. However, if an opposite relation holds good, *i.e.* aggregate demand does not adapt itself to aggregate supply and/or aggregate supply adapts itself to aggregate demand, the "filtration" of the time series of gross domestic product does not result in the potential product estimate. Let us imagine an extreme case that long-term aggregate supply (*i.e*. potential product) is constant and short-term aggregate supply shows certain elasticity in relation to changes in the price level. In this case the cyclic movement in aggregate demand will cause changes both in the product and in the price level. If the reached values of the product are smoothed with HP filter, the result will not be the potential product estimate but the "smoothed" movement of aggregate demand in the economic cycle.

### *Inclusion of past, present and future inflation*

Central banks use forward looking models on which the so called conditional or non--conditional inflation forecasts are based. In *ex-post* empirical verification of the forward looking monetary policy rule, some authors (*e.g.* Mehra, 1999; Frommel, Schobert, 2006) solved the problem by including actual (*i.e*. actually measured) inflation at the future time $t + h$ (monthly time frequency) in the model. We consider this approach as unsuitable. The problem can be elucidated by the following example. Let us assume that inflation at time $t$ is on the level of inflation target. At the same time, forecasted inflation in a conditional forecast for $t + 12$ is above the inflation target. Therefore central bank will increase its interest rate. Due to this measure actual inflation at time $t + 12$

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will remain on the level of inflation target and/or will return to this level. The increase in interest rate at time *t* cannot be explained by means of actually reached inflation at time $t + 12$.

This problem could be solved by the application of the so called conditional inflation forecast. More exactly, a difference between inflation target and forecasted inflation in the framework of conditional forecast can be used as explanatory variable in an econometric model. Currently (or since 2002), when CNB works with the so called non-conditional forecast, data on conditional forecast are not available to the public. The problem of conditional and non-conditional inflation forecast in greater detail (see Dědek, 2004).

### *Backward looking monetary policy rule*

Is it reasonable to test backward looking monetary rules in conditions when central banks present only forward looking models? It is possible to react by asking the following question. How is it possible that J. B. Taylor could successfully empirically verify the backward looking monetary policy rule when the Fed operated models "looking" into the future? There are three potential explanations of this contradiction. Firstly, if repo rate is to be changed, the management board does not take ultimate decisions on the basis of the results and recommendations of forward looking models but they implement backward looking policy based on their own individual decision models. Secondly, forward looking models are only a means of communication between central bank and the public while the analysts respect the backward looking approach of the superior management board through the *ad hoc* manipulation of basic parameters. Thirdly, the management board really takes decisions on the basis of the results and recommendations of forward looking models but logically these models work with exogenous and predetermined variables whose values are known at time *t*. In all cases some inherent statistical information on the future development of central bank's repo rate is considered in the past data.

## 3. Formulation and Empirical Verification of the Model

The formulation of an econometric model was subject to the requirement that all explanatory variables should be observable and statistically available. It was our intention not to work with variables that are acquired (estimated) on the basis of theory and hence on the basis of a number of explicit and implicit assumptions. It applies mainly to different types of equilibrium variables (*e.g*. potential product, natural unemployment rate, natural interest rate, *etc*.) or expected variables (expected inflation, expected GDP, expected exchange rate, *etc*.). This requirement is not only pragmatic, it is to recall traditional approaches of post-Keynesian economics that basically rejects the use of unobserved variables in the analysis of economic processes and accentuates the postulation of critical realism (*e.g*. Jespersen, 2001).

Especially the Czech Republic and Hungary are distinctly open economies. The development of exchange rate basically influences future inflation through a number of

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transmission channels (*e.g.* Hlédik, 2004). In the inflation targeting system central bank will tend to increase repo rate if domestic currency is depreciated in relation to EUR and *vice versa*.

High international mobility of financial capital is characteristic of the considered countries. Therefore it is to expect that monetary policy will be influenced by ECB repo rate development. It is so because a distinct difference of domestic rates in the money market from equivalent foreign interest rates would cause destabilization flows of short-term speculative capital that would consequently increase the exchange rate volatility and threaten the achievement of inflation target.

Even in conditions of forward looking monetary policy operating on the basis of a comparison of inflation forecast and inflation target it can be expected that also "past" inflation will play its role in the setting of repo rate. The last known value of inflation and/or its past development is probably a part of the information portfolio for the definition of a rational forward looking approach. The importance of past inflation will continue to increase when the management loses confidence in the results of their prognostic apparatus based on a number of theoretical variables, speculatively set values of parameters and subsequent dynamic interactive processes.

Pursuant to Act No. 6/1993 on the Czech National Bank the main objective of CNB's monetary policy is to ensure price stability. If this main target is not affected, the CNB has to "support the government's general economic policy leading to sustainable economic growth". The main objective and support of the government's economic policy by the NBP and the MNB are defined in a similar way. Hence it cannot be excluded *a priori* that in certain circumstances and phases of economic development the bank board can take into account the growth rate of GDP when setting the monetary policy interest rate. In our case GDP growth (only quarterly frequency) is approximated by the growth rate in industrial production.

The growth rate of money supply is the last considered explanatory variable (Hu, Phillips, 2004; Senbet, 2010). It is a variable that was an alternative monetary policy objective in the historically preceding monetary transmission mechanism. In the theoretical and model concept of inflation targeting money supply is considered as endogenous, *i.e.* passively adapting itself to real economic processes. Monetary aggregates may play a role as additional monetary indicators containing information on the future (long-term) development of aggregate demand. In this sense they are complementary to (not competing with) the main prognostic model that is conceived as a medium-term one.

In line with standardly considered economic relations we assume that the bank boards of the CNB, the NBP and the MNB will tend to increase monetary policy rates if domestic currency is weakening (*i.e.* the exchange rate increases quantitatively), if the ECB increases repo rate, domestic inflation is increasing (and/or increases above the inflation target), the real growth rate of domestic economy and growth rate of domestic money supply are rising. In all cases these are directly proportional relations.

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### 3.1 Data

The empirical analysis is based on the panels of monthly time series for the period from January 1999 to April 2012. The panels are created by time series of the Czech Republic, Hungary and Poland, and are symbolically marked as "*series_cz, series_hu, series_pl*". The time series are in the log transformation.

*lrepo_cz,_hu,_pl* – monetary policy interest rates (source: national banks),

*lrepo_ecb* – Eurozone repo rate (source: ECB),

*lir_yearly_cz,_hu,_pl* – yearly inflation rate computed as the consumer price index growth rate,

$cpi_t/cpi_{t\text{-}12}$ (*cpi_cz* - Czech Statistical Office, *cpi_hu* – Eurostat, *cpi_pl* – Eurostat),

*lir_annual_cz _hu,_pl* – annualized monthly inflation rate, computed as $(cpi_t/cpi_{t\text{-}1})^{12}$,

*ler_cz,_hu,_pl* – exchange rates of the national currencies to EUR (source ECB),

*lm2_yearly_cz,_hu,_pl* – yearly growth rate of the money stock computed as $M2_t/M2_{t\text{-}12}$ (source of *M2* – national banks)

*lip_yearly_cz_hu_pl* – yearly growth rate of the industrial production computed as $IP_t/IP_{t\text{-}12}$ (source Eurostat).

The above-mentioned time series panels were tested for the presence of a panel unit root. Levin, Lin and Chu (2002), Breitung (2000), Im, Pesaran and Shin (2003), Hadri (2000) and Fisher-type tests using ADF and PP tests [Maddala, Wu (1999) and Choi (2001)] were used. The results are contained in Table 2. In all-time series, except the panel *lip_yearly_cz_hu_pl* of industrial production, most of tests failed to reject the null hypothesis of the unit root presence, so the panels are of I(1) type. In the panel *lip_yearly_cz_hu_pl* the tests rejected the hypothesis of a unit root, so the hypotheses that this panel is of I(0) type was accepted.

The relationship between the yearly inflation rates and the annualized monthly inflation rates is interesting. The annualized inflation rates are volatile and contain significant seasonal components. After their adjustment by Hoedrick-Prescott filter with smoothing parameter 140 (*lir_annhp_cz, _hu, _pl*), it is revealed that the shapes of these time series are very similar to their yearly inflation rate counterparts. It is interesting (and it is clearly seen on the figures of these time series) that the yearly inflation rates lag behind the price level by 6 months. This situation can be interpreted so that the yearly inflation rate gives analogous information about the dynamics of the price level at time *t* as the annualized inflation rate at time *t*-6, *i.e.* about six months earlier. In other words, in the case of the analysis of the yearly inflation rate we waive the information about the development of prices in the last 6 months and we work with six-month old information. In our opinion, this is a serious problem, which was analysed and justified in detail in Arlt and Basta (2008, 2010).

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Table 2
**Panel Unit Root Tests**

| ***lrepo?*** | ***Method*** | ***N*** | ***C*** | ***C, T*** |
|---|---|---|---|---|
| Null: Unit root | LLCH | 0.0140 | **0.4883** | **0.8902** |
| | BREITUNG | x | x | **0.7287** |
| | IPS | x | **0.8123** | **0.8476** |
| | ADF-FISCHER | **0.0792** | **0.7667** | **0.8100** |
| | PR-FISCHER | 0.0305 | **0.7379** | **0.7213** |
| Null: Stationarity | HADRI Z | x | **0.0000** | **0.0000** |
| | HADRI HC Z | x | **0.0000** | **0.0001** |

| ***lir_yearly?*** | ***Method*** | ***N*** | ***C*** | ***C, T*** |
|---|---|---|---|---|
| Null: Unit root | LLCH | 0.0321 | **0.6449** | **0.9191** |
| | BREITUNG | x | x | **0.1070** |
| | IPS | x | **0.1107** | **0.4790** |
| | ADF-FISCHER | **0.1736** | **0.1558** | **0.5793** |
| | PR-FISCHER | **0.1112** | **0.1415** | **0.5714** |
| Null: Stationarity | HADRI Z | x | **0.0003** | **0.0000** |
| | HADRI HC Z | x | **0.0032** | **0.0003** |

| ***lir_annhp?*** | ***Method*** | ***N*** | ***C*** | ***C, T*** |
|---|---|---|---|---|
| Null: Unit root | LLCH | 0.0408 | **0.0910** | **0.5357** |
| | BREITUNG | x | x | **0.4590** |
| | IPS | x | 0.0005 | **0.0818** |
| | ADF-FISCHER | **0.1113** | 0.0006 | **0.0532** |
| | PR-FISCHER | **0.2976** | **0.1531** | **0.7863** |
| Null: Stationarity | HADRI Z | x | **0.0331** | **0.0000** |
| | HADRI HC Z | x | **0.0805** | **0.0001** |

| ***lm2_yearly?*** | ***Method*** | ***N*** | ***C*** | ***C, T*** |
|---|---|---|---|---|
| Null: Unit root | LLCH | 0.0038 | **0.1513** | **0.2364** |
| | BREITUNG | x | x | **0.0689** |
| | IPS | x | **0.0622** | **0.1319** |
| | ADF-FISCHER | 0.0378 | **0.1018** | **0.1860** |
| | PR-FISCHER | **0.0523** | **0.1012** | **0.1503** |
| Null: Stationarity | HADRI Z | x | **0.0000** | **0.0000** |
| | HADRI HC Z | x | **0.0000** | **0.0002** |

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Table 2 – continuation

| ***ler?*** | ***Method*** | ***N*** | ***C*** | ***C, T*** |
|---|---|---|---|---|
| Null: Unit root | LLCH | **0.1626** | **0.1268** | **0.3346** |
| | BREITUNG | x | x | 0.0009 |
| | IPS | x | 0.0312 | 0.0074 |
| | ADF-FISCHER | **0.2367** | 0.0317 | 0.0147 |
| | PR-FISCHER | **0.3640** | **0.2750** | **0.1984** |
| Null: Stationarity | HADRI Z | x | **0.0000** | 0.0430 |
| | HADRI HC Z | x | **0.0000** | 0.0044 |

| ***lip_yearly?*** | ***Method*** | ***N*** | ***C*** | ***C, T*** |
|---|---|---|---|---|
| Null: Unit root | LLCH | 0.0000 | 0.0056 | 0.0004 |
| | BREITUNG | x | x | 0.0056 |
| | IPS | x | 0.0004 | 0.0048 |
| | ADF-FISCHER | 0.0008 | 0.0008 | 0.0095 |
| | PR-FISCHER | 0.0000 | 0.0000 | 0.0000 |
| Null: Stationarity | HADRI Z | x | 0.4059 | 0.5093 |
| | HADRI HC Z | x | 0.5694 | 0.4766 |

| ***lrepo_ecb*** | ***Method*** | ***N*** | ***C*** | ***C, T*** |
|---|---|---|---|---|
| Null: Unit root | ADF | **0.2037** | **0.6281** | **0.4851** |

Note: Exogenous variable – *N*-none, *C*-individual effects, *C, T*-individual effects, individual linear trends

Source: Own computations

## 3.2 Model

### *Cointegration tests*

As the time series panels are nonstationary of I(1) type, it is appropriate to test the cointegration of these panels. The Pedroni (1999, 2004) and Kao (1999) panel cointegration tests follow the classical Engle-Granger (1987) two-step (residual-based) cointegration test. The Engle-Granger (1987) cointegration test is based on the analysis of residuals of regression performed using I(1) variables. If the variables are cointegrated, then the residuals must be I(0). On the other hand, if the variables are not cointegrated, then the residuals are I(1).

Pedroni proposed a test for panel cointegration that allows for heterogeneous intercepts and trend coefficients across cross-sections. Consider the following regression

$$Y_{it} = \alpha + \boldsymbol{X}_{it}'\boldsymbol{\beta}_i + \delta_i + \varepsilon_{it} \quad (1)$$

for $t = 1, 2, ..., T$; $i = 1, 2, ..., N$ (in our case, $N = 3$ countries – cz, hu, pl); where $Y_{it}$ is individual “cross-sectional” time series and $\boldsymbol{X}_{it}$ is a vector of the “cross-sectional” time

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series, the panels of $N$ time series are assumed to be integrated of order one, *e.g.* I(1). Parameter $\alpha$ represents the overall constant in the model, parameters $\delta_i$ are individual effects, which may be set to zero if desired, the parameter vector $\boldsymbol{\beta}_i$ contains cross-section specific regression parameters, $\varepsilon_{it}$ are the error terms. Under the null hypothesis of no cointegration, the residuals will be I(1). The general approach is to obtain residuals from the above-mentioned equation and then to test whether they are I(1) by running the auxiliary regression,

$$\varepsilon_{it} = \rho_i \varepsilon_{it-1} + u_{it}$$

or

$$\varepsilon_{it} = \rho_i \varepsilon_{it-1} + \sum_{j=1}^{p} \Psi_{ij} \Delta\varepsilon_{it-j} + v_{it}$$

for each cross-section, where it is supposed that $u_{it}$ and $v_{it}$ are independent and identically distributed. Pedroni described various statistics for testing the null hypothesis of no cointegration, $\rho_i = 1$. There are two alternative hypotheses: the homogeneous alternative, $\rho_i = \rho < 1$, for all $i$ and the heterogeneous alternative, $\rho_i < 1$, for all $i$.

The Kao test follows the same basic approach as the Pedroni tests, but specifies cross-section specific intercepts and cross-section homogeneous regression parameters, *i.e.*

$$Y_{it} = \alpha + \boldsymbol{X}_{it}'\boldsymbol{\beta} + \delta_i + \varepsilon_{it}. \qquad (2)$$

The residuals obtained from the above-mentioned equation are tested for unit root through the auxiliary regression

$$\varepsilon_{it} = \rho \varepsilon_{it-1} + u_{it}$$

or

$$\varepsilon_{it} = \rho \varepsilon_{it-1} + \sum_{j=1}^{p} \Psi_{ij} \Delta\varepsilon_{it-j} + v_{it}.$$

The hypothesis of no cointegration is $\rho = 1$, the alternative hypothesis of cointegration is $\rho < 1$. He suggested different test statistics.

Fisher (1932) derived a combined test that uses the results of particular independent tests. Maddala and Wu (1999) used Fisher's result to propose an alternative approach to testing for cointegration in panel data by combining Johansen tests from individual cross-sections. This test is called Johansen Fisher Panel Cointegration Test.

### *Model*

In the case of panel cointegration, $\varepsilon_{it}$ in model (1) creates the stationary time series panel. The applied time series are frequently in the log form, the main reason for this transformation is usually the reduction of heteroscedasticity of error term (for example Verbeek, 2004). The cross-section specific regression parameters $\boldsymbol{\beta}_i$ of such model are called elasticities and are interpreted as relative changes in dependent variables *per* unit relative changes in the respective explanatory variables. They express the long-run relationships of the time series.

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The one-equation model follows from the above formulated economic hypothesis, which is verified in this paper. The board of the national bank decides on the level of repo rate on the basis of available information and knowledge of the problem. It implies that repo rate is an endogenous variable in this case and other variables are exogenous. In this situation the one-equation model is sufficient for describing the relationship between variables.

In time series model (1) the problem of autocorrelation occurs very frequently. When estimating the model parameters by the least-squares method, the autocorrelation leads *inter alia* to underestimation of the parameter standard error estimates and thus to overestimation of *t*-tests, which can lead to erroneous conclusions when testing the model parameters. In this situation it is suitable to use the *Period SUR* (Period Seemingly Unrelated Regressions) (*PCSC* – "Panel Corrected Standard Error") method which provides the standard error estimates which are robust to the residual autocorrelation for a given cross-section. This method leads also to the robust results to heteroscedasticity across the cross-sections or periods (Beck and Katz, 1995).

### 3.3 Empirical Analysis and Results

Tables 3, 4, 5 contain the correlation coefficients of pairs of time series in the Czech Republic, Hungary and Poland. The correlation coefficients with repo rates are relatively high in all countries. The exception is the yearly growth rate of the industrial production (*ip_yearly*) in all three countries and exchange rate in Poland (*er_pl*) (in industrial production in the Czech Republic and Poland and in the exchange rate in Poland, the correlation coefficients are negative). This is the reason why these time series are not included in the cointegration analysis. Very high correlation coefficients of the yearly and annualized inflation rates are logical and expected. It is certainly interesting that the correlation coefficients of the yearly inflation rates with monetary policy interest rates are higher in all countries than the correlation coefficients of annualized inflation rates with monetary policy interest rates.

Table 3
**Correlation Matrix of Log Time Series of the Czech Republic**

| | ***lrepo_cz*** | ***lmi_yearly_cz*** | ***lmi_annhp_cz*** | ***ler_cz*** | ***lrepo_ecb*** | ***lm2_yearly_cz*** | ***lip_yearly_cz*** |
|---|---|---|---|---|---|---|---|
| ***lrepo_cz*** | 1.0000 | | | | | | |
| ***lir_yearly_cz*** | **0.3942** | 1.0000 | | | | | |
| ***lir_annhp_cz*** | **0.2516** | **0.7120** | 1.0000 | | | | |
| ***ler_cz*** | **0.7661** | -0.0204 | 0.0444 | 1.0000 | | | |
| ***lrepo_ecb*** | **0.8556** | 0.5370 | 0.4350 | 0.5132 | 1.0000 | | |
| ***lm2_yearly_cz*** | **0.6075** | 0.5355 | 0.4217 | 0.3034 | 0.6126 | 1.0000 | |
| ***lip_yearly_cz*** | -0.0823 | 0.1347 | 0.3490 | 0.0381 | 0.1610 | -0.0023 | 1.0000 |

Source: Own computations

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Table 4

**Correlation Matrix of Log Time Series of Hungary**

| | *lrepo_hu* | *lmi_yearly_hu* | *lmi_annhp_hu* | *ler_hu* | *lrepo_ecb* | *lm2_yearly_hu* | *lip_yearly_hu* |
|---|---|---|---|---|---|---|---|
| *lrepo_hu* | **1.0000** | | | | | | |
| *lir_yearly_hu* | **0.6886** | 1.0000 | | | | | |
| *lir_annhp_hu* | **0.5332** | ***0.7406*** | 1.0000 | | | | |
| *ler_hu* | **-0.3732** | -0.2616 | 0.1467 | 1.0000 | | | |
| *lrepo_ecb* | **0.5324** | 0.5778 | 0.3577 | -0.5900 | 1.0000 | | |
| *lm2_yearly_hu* | **0.6852** | 0.3029 | 0.3636 | -0.4734 | 0.5999 | 1.0000 | |
| *lip_yearly_hu* | 0.0335 | 0.3697 | 0.3019 | -0.2475 | 0.2572 | 0.0776 | 1.0000 |

Source: Own computations

Table 5

**Correlation Matrix of Log Time Series of Poland**

| | *lrepo_pl* | *lmi_yearly_pl* | *lmi_annhp_pl* | *ler_pl* | *lrepo_ecb* | *lm2_yearly_pl* | *lip_yearly_pl* |
|---|---|---|---|---|---|---|---|
| *lrepo_pl* | **1.0000** | | | | | | |
| *lir_yearly_pl* | **0.7249** | 1.0000 | | | | | |
| *lir_annhp_pl* | **0.5545** | ***0.8546*** | 1.0000 | | | | |
| *ler_pl* | -0.1625 | -0.0623 | 0.1599 | 1.0000 | | | |
| *lrepo_ecb* | **0.6441** | 0.2686 | 0.1446 | -0.5064 | 1.0000 | | |
| *lm2_yearly_pl* | **0.2889** | 0.5413 | 0.6349 | -0.1831 | 0.2919 | 1.0000 | |
| *lip_yearly_pl* | -0.1645 | -0.0992 | 0.0689 | 0.1759 | -0.0695 | -0.2006 | 1.0000 |

Source: Own computations

Kao and Johansen Fisher tests from Table 6 and Table 7 indicate that the above-mentioned panels, *i. e. lrepo?*, *lir_yearly?*, *lir_annhp?*, *ler?*, *lm2_yearly?*, and time series *lrepo_ecb* are cointegrated ("?" indicates "*_cz*, *_hu*, *_pl*").

Table 6

**The Panel Cointegration Tests - lrepo? lir_yearly? ler? lm2_yearly? lrepo_ecb**

| **Pedroni Residual Cointegration Test**<br>Series: *lrepo? lir_yearly? ler? lm2_yearly? lrepo_ecb*<br>Null Hypothesis: No cointegration<br>Trend assumption: No deterministic trend | | | | |
|---|---|---|---|---|
| **Alternative hypothesis: common AR coefs. (within-dimension)** | | | | |
| | **Statistic** | **Prob.** | **Weighted Statistic** | **Prob.** |
| Panel v-Statistic | -0.215750 | 0.5854 | -0.712041 | 0.7618 |
| Panel rho-Statistic | 0.195587 | 0.5775 | 0.478101 | 0.6837 |
| Panel PP-Statistic | -0.152201 | 0.4395 | 0.138205 | 0.5550 |
| Panel ADF-Statistic | -0.125459 | 0.4501 | -0.179314 | 0.4288 |

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Table 6 – continuation

| Alternative hypothesis: individual AR coefs. (between-dimension) | | | | |
|---|---|---|---|---|
| | **Statistic** | **Prob.** | | |
| Group rho-Statistic | 0.618464 | 0.7319 | | |
| Group PP-Statistic | 0.120424 | 0.5479 | | |
| Group ADF-Statistic | 0.231780 | 0.5916 | | |

| **Kao Residual Cointegration Test**<br>Series: *lrepo? lir_yearly? ler? lm2_yearly? lrepo_ecb*<br>Null Hypothesis: No cointegration<br>Trend assumption: No deterministic trend | | |
|---|---|---|
| ADF | **t-Statistic** | **Prob.** |
| | -2.331336 | **0.0099** |

| **Johansen Fisher Panel Cointegration Test**<br>Series: *lrepo? lir_yearly? ler? lm2_yearly? lrepo_ecb*<br>Trend assumption: Linear deterministic trend | | | | |
|---|---|---|---|---|
| **Hypothesized No. of CE(s)** | **Fisher Stat.* (from trace test)** | **Prob.** | **Fisher Stat.* (from max-eigen test)** | **Prob.** |
| **None** | **39.81** | **0.0000** | **34.80** | **0.0000** |
| ***At most 1*** | ***12.62*** | ***0.0495*** | **16.03** | **0.0136** |
| At most 2 | 2.876 | 0.8242 | 4.101 | 0.6631 |
| At most 3 | 1.101 | 0.9815 | 0.985 | 0.9862 |
| At most 4 | 5.042 | 0.5385 | 5.042 | 0.5385 |

*Probabilities are computed using asymptotic Chi-square

Source: Own computations

Table 7

**The Panel Cointegration Tests - lrepo? lir_annhp? ler? lm2_yearly? lrepo_ecb**

| **Pedroni Residual Cointegration Test**<br>Series: *lrepo? lir_annhp? ler? lm2_yearly? lrepo_ecb*<br>Null Hypothesis: No cointegration<br>Trend assumption: No deterministic trend | | | | |
|---|---|---|---|---|
| **Alternative hypothesis: common AR coefs. (within-dimension)** | | | | |
| | **Statistic** | **Prob.** | **Weighted Statistic** | **Prob.** |
| Panel v-Statistic | -0.536709 | 0.7043 | -0.930981 | 0.8241 |
| Panel rho-Statistic | 0.853037 | 0.8032 | 1.090629 | 0.8623 |
| Panel PP-Statistic | 0.611290 | 0.7295 | 0.872656 | 0.8086 |
| Panel ADF-Statistic | 1.068661 | 0.8574 | 1.308763 | 0.9047 |

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Table 7 – continuation

| **Alternative hypothesis: individual AR coefs. (between-dimension)** | | | | |
|---|---|---|---|---|
| | **Statistic** | **Prob.** | | |
| Group rho-Statistic | 1.389144 | 0.9176 | | |
| Group PP-Statistic | 1.050298 | 0.8532 | | |
| Group ADF-Statistic | 1.597743 | 0.9449 | | |
| **Kao Residual Cointegration Test**<br>Series: *lrepo? lir_annhp? ler? lm2_yearly? lrepo_ecb*<br>Null Hypothesis: No cointegration<br>Trend assumption: No deterministic trend | | | | |
| ADF | | | **t-Statistic** | **Prob.** |
| | | | -2.054722 | **0.0200** |

| **Johansen Fisher Cointegration Test**<br>Series: *lrepo? lir_annhp? ler? lm2_yearly? lrepo_ecb* | | | | |
|---|---|---|---|---|
| **Trend assumption: Linear deterministic trend** | | | | |
| **Hypothesized No. of CE(s)** | **Fisher Stat.* (from trace test)** | **Prob.** | **Fisher Stat.* (from max-eigen test)** | **Prob.** |
| **None** | **131.6** | **0.0000** | **107.9** | **0.0000** |
| **At most 1** | **37.59** | **0.0000** | **35.34** | **0.0000** |
| At most 2 | 10.40 | 0.1088 | 11.58 | 0.0720 |
| At most 3 | 3.745 | 0.7111 | 3.532 | 0.7398 |
| At most 4 | 5.647 | 0.4638 | 5.647 | 0.4638 |

*Probabilities are computed using asymptotic Chi-square

Source: Own computations

From Granger cointegration theory it follows that in the case of cointegration the static model is able to catch the long-run behaviour of time series. Table 8 contains the parameters estimate of the static panel model, in which it is assumed that the monetary policy interest rates of the Czech Republic, Hungary and Poland are explained by the yearly inflation rates, exchange rates of the currencies of the above-mentioned states to EUR, ECB repo rate and by the yearly growth rate of M2. As the panel of the yearly growth rate of the industrial production is of I(0) type [the other time series panels are of I(1) type] and the correlations of these time series with monetary policy interest rates are very low, this panel was not included in the model.

The time series of the yearly inflation rates are correlated with the yearly growth rate of M2 and with ECB repo rate, the exchange rates are correlated with the growth rate of M2 and with ECB repo rate, which is correlated with the growth rate of M2. This situation could lead to multicollinearity. But the empirical analysis does not seem to show that the possible multicollinearity would overestimate the parameter standard errors. Therefore, all above-mentioned data panels are included in the models.

With respect to the results of the cointegration tests, the estimated parameters of the models can be interpreted as cointegration parameters.

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Table 8

**The Model with the Yearly Inflation Rate**

| | | | | |
|---|---|---|---|---|
| Dependent Variable: *lrepo?*<br>Method: Pooled Least Squares<br>Sample (adjusted): 1999M01 2012M04<br>Cross-sections included: 3<br>Period SUR (PCSE) standard errors & covariance (d.f. corrected) | | | | |
| **Variable** | **Coefficient** | **Std. Error** | **t-statistic** | **Prob.** |
| *C* | -1.626396 | 1.739201 | -0.935140 | 0.3502 |
| *lir_yearly_cz* | 2.489609 | 0.324316 | 7.676488 | 0.0000 |
| *lir_yearly_hu* | 8.591064 | 1.702734 | 5.045453 | 0.0000 |
| *lir_yearly_pl* | 15.61207 | 0.613811 | 25.43465 | 0.0000 |
| *ler_cz* | 2.353280 | 0.694060 | 3.390601 | 0.0008 |
| *ler_hu* | -0.366563 | 0.845107 | -0.433747 | 0.6647 |
| *ler_pl* | 0.945507 | 0.563574 | 1.677699 | 0.0941 |
| *lrepo_ecb* | 0.673775 | 0.107799 | 6.250299 | 0.0000 |
| *lrepo_ecb* | -0.137616 | 0.191161 | -0.719898 | 0.4719 |
| *lrepo_ecb* | 0.644877 | 0.064131 | 10.05563 | 0.0000 |
| *lm2_yearly_cz* | 2.926710 | 0.585694 | 4.996999 | 0.0000 |
| *lm2_yearly_hu* | 3.770998 | 1.516362 | 2.486872 | 0.0132 |
| *lm2_yearly_pl* | -2.268891 | 0.779282 | -2.911513 | 0.0038 |
| Fixed Effects (Cross) | | | | |
| *_cz—c* | -6.301367 | | | |
| *_hu—c* | 5.027000 | | | |
| *_pl—c* | 1.274367 | | | |

$R^2$=0.923517 *s*=0.212650 *F*=401.0562(0.0000) *DW*=0.154749 *AIC*=-0.227589

*Source:* Own computations

The Durbin-Watson statistic indicates the presence of autocorrelation, which leads to the overestimation of the index of determination $R^2$, *F*-test and *t*-test. Deformation of *t*-test was eliminated by using the *Period SUR* method of the standard error correction. The estimates of parameters for the explanatory variables are in line with the economic theory. Exceptions are the estimates of parameters for exchange rate in Hungary, ECB repo rate as an explanatory variable for repo rate in Hungary and growth rate of M2 in Hungary. The estimates of the first two parameters are statistically non-significant. The parameter estimate for the exchange rate in Poland is significant at 10%.

In the model in Table 9, instead of the yearly inflation rate, the annualized monthly inflation rate smoothed with HP filter is included. It is seen that from the aspect of the index of determination, residual standard deviation *s* and from the aspect of *AIC* criterion (Akaike Information Criterion) this model is worse than the first one. The parameter estimate for the annualized inflation rate in the Czech Republic is statistically non-significant (for other countries these estimates are statistically significant), which confirms the reality

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that when deciding on the monetary policy interest rate in the Czech National Bank this variable is not considered (also the correlation coefficients of the annualized inflation rates with the monetary policy interest rates in all countries are lower than the correlation coefficients of the yearly inflation rates with the monetary policy interest rates, see Tables 1, 2 and 3). This conclusion can have also another interpretation. In accordance with what was stated in Part 3.1, there is an approximate relationship $lir\_yearly_t \approx lir\_annhp_{t\text{-}6}$ and therefore $lir\_yearly_{t+6} \approx lir\_annhp_t$, which means that the annualized smoothed monthly inflation rate at time *t* can be considered as a very good forecast of the yearly inflation rate at time *t*+6. The statistically non-significant parameter estimate for the annualized inflation rate can be understood as the acceptance of the hypothesis that when deciding on the monetary policy interest rate, the forecast of the future yearly inflation rate development with 6-month horizon is not considered. Similarly, like in the first model, the parameter estimates for the exchange rate in Poland and ECB repo rate to explain the Hungary repo rate are statistically non-significant. The parameter estimate for the growth rate of M2 in Hungary is significant at 10%.

Table 9

**The Model with the Annualized Inflation Rate**

Dependent Variable: *lrepo?*
Method: Pooled Least Squares
Sample (adjusted): 1999M01 2012M04
Cross-sections included: 3
Period SUR (PCSE) standard errors & covariance (d.f. corrected)

| **Variable** | **Coefficient** | **Std. Error** | **t-statistic** | **Prob.** |
|---|---|---|---|---|
| *c* | 1.001719 | 1.521935 | 0.658188 | 0.5107 |
| *lir_annhp_cz* | -3.182136 | 2.602205 | -1.222861 | 0.2220 |
| *lir_annhp_hu* | 5.314066 | 1.067078 | 4.980015 | 0.0000 |
| *lir_annhp_pl* | 14.39563 | 2.021364 | 7.121740 | 0.0000 |
| *ler_cz* | 2.115216 | 0.860414 | 2.458371 | 0.0143 |
| *ler_hu* | -1.361053 | 0.633276 | -2.149226 | 0.0321 |
| *ler_pl* | -0.005502 | 0.701421 | -0.007844 | 0.9937 |
| *lrepo_ecb* | 0.765119 | 0.089774 | 8.522702 | 0.0000 |
| *lrepo_ecb* | 0.001970 | 0.125042 | 0.015754 | 0.9874 |
| *lrepo_ecb* | 0.701749 | 0.076180 | 9.211739 | 0.0000 |
| *lm2_yearly_cz* | 4.035004 | 1.172141 | 3.442422 | 0.0006 |
| *lm2_yearly_hu* | 2.755437 | 1.616882 | 1.704167 | 0.0890 |
| *lm2_yearly_pl* | -3.001507 | 0.622580 | -4.821082 | 0.0000 |
| Fixed Effects (Cross) | | | | |
| *_cz—c* | -8.129895 | | | |
| *_hu—c* | 8.095036 | | | |
| *_pl—c* | 0.034859 | | | |

$R^2$=0.896106 *s*=0.247844 *F*=286.4991(0.0000) *DW*=0.107729 *AIC*=0.078716
Source: Own computations

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The model in Table 8 indicates that when setting the repo rates, the actual yearly inflation rate plays an important role. Unlike this model, the model in Table 10 includes the yearly inflation rate lagged by 6 months. The signs of the parameter estimates are the same as in the first model. However, one can easily find that this model is better than the first one from the aspect of the index of determination $R^2$, residual standard deviation $s$ and $AIC$ criterion. These results show that when setting the repo rate, the last level of the yearly inflation rate plays a more important role than its present level. The parameter estimate signs and significance are the same as in the first model.

Table 10

**The Model with the Lagged Yearly Inflation Rate**

Dependent Variable: *lrepo?*
Method: Pooled Least Squares
Sample (adjusted): 1999M01 2012M04
Cross-sections included: 3
Period SUR (PCSE) standard errors & covariance (d.f. corrected)

| Variable | Coefficient | Std. Error | t-statistic | Prob. |
|---|---|---|---|---|
| *c* | -2.570955 | 1.670669 | -1.538877 | 0.1245 |
| *lir_yearly_cz(-6)* | 6.995287 | 1.121889 | 6.235273 | 0.0000 |
| *lir_ yearly_hu(-6)* | 7.241018 | 1.632448 | 4.435681 | 0.0000 |
| *lir_ yearly_pl(-6)* | 15.38098 | 0.692643 | 22.20621 | 0.0000 |
| *ler_cz* | 2.334031 | 0.360305 | 6.477940 | 0.0000 |
| *ler_hu* | -0.010002 | 0.862057 | -0.011603 | 0.9907 |
| *ler_pl* | 1.616265 | 0.679173 | 2.379753 | 0.0177 |
| *lrepo_ecb* | 0.637934 | 0.066661 | 9.569779 | 0.0000 |
| *lrepo_ecb* | -0.094418 | 0.145583 | -0.648551 | 0.5169 |
| *lrepo_ecb* | 0.625904 | 0.052125 | 12.00765 | 0.0000 |
| *lm2_yearly_cz* | 1.807061 | 1.052650 | 1.716677 | 0.0867 |
| *lm2_yearly_hu* | 3.562914 | 0.680297 | 5.237291 | 0.0000 |
| *lm2_yearly_pl* | -2.156427 | 0.683558 | -3.154708 | 0.0017 |
| Fixed Effects (Cross) | | | | |
| *_cz—c* | -5.302107 | | | |
| *_hu—c* | 4.035888 | | | |
| *_pl—c* | 1.266219 | | | |

$R^2$=0.941897 $s$=0.185345 $F$=538.4345(0.0000) $DW$=0.227591 $AIC$=-502446
Source: Own computations

## 4. Conclusion

We ask a question in our analysis whether it is possible to verify successfully empirically the backward looking monetary policy rule when the central bank declares the implementation of monetary policy on the basis of the forward looking monetary policy

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rule. Our experience has shown that the analysts in the commercial sphere do not have an opportunity to verify central bank's behaviour when repo rates are changed on the basis of complicated multi-equation models used by central banks. Let us imagine the following potential situations. Firstly, if repo rate is changed, the management board does not take the ultimate decision on the basis of results and recommendations of forward looking models and implements backward looking policy based on their own individual decision models. Secondly, forward looking models are only a "means of communication" between the central bank and the public and through the *ad hoc* manipulation of basic parameters of the model the analysts respect the backward looking approach of the superior management board. Thirdly, the management boards really takes a decision on the basis of results and recommendations of forward looking models, but logically, these models also work with exogenous and predetermined variables whose values are known at the time of decision-making.

The empirical analysis is based on monthly time series for the period of January 1999 to April 2012 and involves these variables: central bank's monetary policy interest rate, spot exchange rate of the national currency for EUR, ECB repo rate, yearly inflation rate, annualized monthly inflation rate, yearly growth rate of M2, yearly growth rate of industrial production. Our empirical research demonstrates that the correlation coefficients of particular time series with central bank's monetary policy interest rate are relatively high in most cases. The time series of the yearly growth rate of industrial production are an exception. These variables appeared as statistically insignificant also in tested econometric models, therefore they have not been included in the resultant published form of models.

Our analysis revealed that the estimates of parameters in the explanatory variables of annual inflation rate, exchange rate, ECB repo rate and the yearly growth rate of M2 mostly have the signs in line with economic theory and are statistically significant. The results for Hungary were found to be worse (parameters in the variables of exchange rate and ECB repo rate). Instead of the yearly inflation rate the annualized monthly inflation rate smoothed with HP filter was included in the alternative model that can also be understood as a very good forecast of future development of the yearly inflation rate with six-month horizon. Taking into account the index of determination, the residual standard deviation and Akaike information criterion, this alternative model is generally worse than the basic model. The estimate of the parameter in annualized inflation rate was statistically insignificant for the Czech Republic, which confirms that the decision on CNB repo rate does not consider this inflation rate. It seems surprising that the quality of the estimated model further increased after the six-month lag of the yearly inflation rate. It is indicated by the values of index of determination, residual standard deviation, Akaike information criterion and standard error of the parameter estimates. Hence the obtained results document that in the studied central banks the past level of the yearly inflation rate plays a more important role in the setting of monetary policy interest rate than its present level.

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## References

**Andrews, D. W. K.** (1991), "Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimation." *Econometrica*, Vol. 59, No. 3, pp. 817–858.

**Andrews, D. W. K, Monahan, J. CH.** (1992), "An Improved Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimator." *Econometrica,* Vol. 60, No. 4, pp. 953–966.

**Arlt, J., Arltová, M.** (2009), *Ekonomické časové řady*. Praha: Professional Publishing.

**Arlt, J., Bašta, M.** (2008), "Time Series of Monthly and Yearly Inflation Rates and Their Properties". (in Czech) *Politická ekonomie*, Vol. 56, No. 4, pp. 536–556.

**Arlt, J., Bašta, M.** (2010), "The Problem of the Yearly Inflation Rate and Its Implications for the Monetary Policy of the Czech National Bank." *Prague Economic Papers*, Vol. 19, No. 2, pp. 99-117.

**Arlt, J., Mandel, M.** (2012), "Is It Possible to Predict the CNB Repo Rate on the Basis of the Backward-Looking Monetary Rule?" (in Czech) *Politická ekonomie*, Vol. 60, No. 4, pp. 484–504.

**Beneš, J., Hlédik, T., Kumhof, M., Vávra, D.** (2005), "An Economy in Transition and DSGE: What the Czech National Bank's New Projection Model Needs." CNB Working Paper, 2005, No. 12.

**Budnik, K., Kolasa, M., Hulej, M., Greszta, M., Murawski, K., Rot. M., Tarnicka, M., Rybaczyk, B.** (2008). NECMOD. National Bank of Poland, June 2008.

**Brada, J., Brůna, K.** (2004), "An Analysis of PRIBOR Interest Rates Sensitivity to Changes in Czech National Bank Repo Rate." (in Czech) *Politická ekonomie*, Vol. 52, No. 5, pp. 601–621.

**Chinn, M. D., Meredith, G.** (2004), "Monetary Policy and Long-Horizon Uncovered Interest Parity." IMF Staff Papers, Vol. 51, No. 3, 2004.

**Den Haan, W. J., Levin, A.** (1997), "A Practitioner's Guide to Robust Covariance Matrix Estimation." In Maddala, G. S., Rao, C. R. (eds.), *Handbook of Statistics Vol. 15, Robust Inference*, pp. 291–341. Amsterdam: North-Holland.

**Dědek, O.** (2004), "Four Reflections on Practising Inflation Targeting in the Czech Republic." *Politická ekonomie*, Vol. 52, No. 2, pp. 147–170.

**Durčáková, J., Mandel, M., Tomšík, V.** (2005), "Puzzle in the Theory of Uncovered Interest Rate Parity-Empirical Verification for Transitive Countries." *Finance India*, 2005, Vol. 19, No. 2, pp. 449–464.

**Frommel, M., Schobert, V.** (2006), "Monetary Policy Rules in Central and Eastern Europe." Deutsche Bundesbank Discussion Paper, 2006, No. 341.

**Hlédik, T.** (2004), "Quantifying the Second-Round Effects of Supply-Side Shocks on Inflation." *Prague Economic Papers*, Vol. 13, No. 2, pp. 125–141.

**Horváth, A., Kober, C., Szilagyi, K.** (2011), "The Magyar Nemzeti Bank´s Monetary Policy Model." *MNB Bulletin*, June 2011.

**Horváth, R.** (2008), "Asymmetric Monetary Policy in the Czech Republic?" *Czech Journal of Economics and Finance,* Vol. 58, No. 9–10, pp. 470–481.

**Hu, L., Phillips, P. C. B.** (2004), "Dynamics of the Federal Funds Target Rate: a Nonstationary Discrete Choice Approach." *Journal of Applied Econometrics*, October 2004, Vol. 19, No. 7, pp. 851-867.

**Jespersen, J.** (2011). *Macroeconomic Methodology: A Post-Keynesian Perspective*. Cheltenham: Edward Elgar Publishing, Ltd., 2011.

**Kukal, J., Van Quang, T.** (2011), "Modeling the CNB's Monetary Policy Interest Rate by Artificial Neural Networks." (in Czech). *Politická ekonomie*, Vol. 59, No. 6, pp. 810–829.

**Mandel, M., Kosmata, V.** (2000), "The Czech Embarrassment of Inflation Targeting in Transition." *Prague Economic Papers*, Vol. 9, No. 3, pp. 230–241.

**Maria-Dolores, R.** (2005), "Monetary Policy Rules in Accession Countries to EU: Is the Taylor Rule a Pattern?" *Economics Bulletin*, Vol. 5, No. 5, pp. 1–16.

DOI: 10.18267/j.pep.484

**Mehra, Y. P.** (1999), "A Forward–Looking Monetary Policy Reaction Function. *FRB of Richmond Economic Quarterly*, Vol. 85, No. 2, pp. 33–53.

**Mohanty, M. S., Klau, M.** (2004), "Monetary Policy Rules in Emerging Market Economies: Issues and Evidence." BIS Working Paper, 2004, No. 149.

**Newey, W. K., West, K. D.** (1987), "A Simple Positive Semi-Definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix." *Econometrica*, Vol. 55, No. 3, pp. 703–708.

**Paez-Farrell, J.** (2007), "Understanding Monetary Policy in Central European Countries Using Taylor-Type Rules: The Case of the Visegrad Four." *Economics Bulletin*, Vol. 5, No. 3, pp. 1–11.

**Senbet, D.** (2010), "Estimating Monetary Policy Reaction Function: A Factor-Augmented Sector Autoregressive (FAVAR) Approach." *Journal of Money, Investment and Banking*, 2010, Vol. 19, pp. 85–113.

**Svensson, L. E. O.** (1998), "Open-Economy Inflation Targeting." NBER Working Paper Series 6545, May 1998.

**Svensson, L. E. O.** (1999), "Inflation Targeting as a Monetary Policy Rule." *Journal of Monetary Economics*, Vol. 43, No. 3, pp. 607–654.

**Škop, J.** (2012). Central Bank Reaction Functions in CEE. Societe Generale, Cross Asset Research, March 2011, pp. 10–14.

**Taylor, J. B.** (1993), "Discretion versus Policy Rules in Practice." *Carnegie-Rochester Conference Series in Public Policy,* Vol. 39, December 1993, pp. 195–214.

**Taylor, J. B.** (2001), "The Role of the Exchange Rate in Monetary-Policy Rules." *American Economic Review, Papers and Proceedings*, Vol. 91, No. 2, pp. 263–267.

**Vašíček, B.** (2011), "Is Monetary Policy in the New EU Member States Asymmetric?" CNB Working Paper, 2011, No. 5.

**Verbeek, M.** (2004), *A Guide to Modern Econometrics*. Chichester, UK: John Wiley and Sons, Ltd., 2004.