

Zorica Mladenović

Faculty of Economics,
University of Belgrade,
Serbia

✉ zorima@eunet.rs

**Aleksandra
Nojković**

Faculty of Economics,
University of Belgrade,
Serbia

✉ nojkovic@gmail.com

Inflation Persistence in Central and Southeastern Europe: Evidence from Univariate and Structural Time Series Approaches

Summary: The purpose of this paper is to measure inflation persistence in the following countries of Central and Southeastern Europe: Slovakia, the Czech Republic, Poland, Hungary, Romania and Serbia. The study sample covers monthly data from January, 1995 to May, 2010 for Poland, Hungary and Slovakia, from January 1994 to May, 2010 for the Czech Republic, and from January, 2002 to June 2010 for Romania. The shortest sample used, from January, 2003 to September, 2010, was for Serbia and is due to the late start in the transition process. The results of this study enriched the existing ones on this topic by extending the sample period to cover even the recent years of relatively higher inflation rates and by including Romania and Serbia, which were not previously considered. The study led to two main findings: first, inflation of moderate to high magnitude persistence in Hungary, Poland, Romania and Serbia, and inflation of smaller order persistence in Slovakia and the Czech Republic was detected within the Markov switching model approach. In addition, the changes in inflation persistence often correspond to changes in variability and mean of inflation. Second, New Keynesian Phillips Curve represents a valid structural approach to describe the inflation dynamics in this region. In all the six cases studied, weights on backward and forward looking behaviors were significant, while the impact of the driving variable was insignificant only once. It is found that significant influence of the economic driving variable can be captured by real gross wage inflation and real broad money growth. The estimates show that the backward-looking term plays an important role in determining the inflation dynamics. Similar conclusions are drawn by using quarterly data in econometric estimations for the selected countries.

Key words: Inflation persistence, Inflation targeting, Markov switching models, The new hybrid Phillips curve model.

JEL: E31, E52.

Inflation persistence is one of the key parameters for adjusting monetary policy; accordingly, the significance of this parameter stems from the primary goal, adopted by most central banks, of stabilizing inflation at a low level. Given a specific shock, inflation persistence can be described as the tendency of the inflation rate to converge slowly toward its long-run value (Ivan Paya, Agustin Duarte, and Ken Holden 2007). Therefore, the magnitude of inflation persistence provides valuable information for central banks, enabling monetary policy to be framed such that it facilitates the achievement of the announced target level. The policy implications establishes the

relevance of understanding the path and determinants of inflation persistence to the Central and Southeastern European countries (CSEECs) that have adopted inflation targeting. More specifically, the difference in inflation persistence between current euro countries and CSEECs that are EU members could suggest the asymmetric impact of a common shock between two groups of countries; furthermore, the difference in persistence may indicate the likelihood of convergence, upon adopting the euro, of the CSEECs towards the level of persistence in the current euro area. In addition, the success of recently introduced inflation targeting in some CSEECs can be assessed by considering the experience of other economies from the same region.

It remains widely believed that the adoption of inflation targeting should be associated with a sharp fall in the persistence of inflation. The performance of inflation targeting has been extensively studied in major industrial countries (the U.S. and euro area) and, in recent years, has been extended to include emerging market economies. The survey based on descriptive and econometric evidence (Pierre L. Siklos 2008) confirms the prevailing empirical finding that inflation targeting diminishes inflation persistence in industrial countries; however, inflation targeting has had less success diminishing inflation persistence in emerging market economies.

Inflation persistence can be analyzed with both micro and macro data. Micro data analysis is based on the statistical properties of product-level consumer price indexes underlying the consumer basket, while macro data analysis is based on time series univariate and structural modeling. The univariate approach allows the application of different methods to estimate the extent to which inflation reacts to unexpected random shocks, while structural time series modeling facilitates the estimation of the hybrid New Keynesian Phillips Curve (NKPC).

For this study, the following CSEECs were considered: the Czech Republic, Poland, Hungary, Romania and Serbia. Among these, the Czech Republic, Poland and Hungary have utilized inflation targeting longer than Romania and Serbia, which only adopted it recently. The objective of this study has been to measure the magnitude of inflation persistence in these five economies. Likewise, a similar study was carried out for Slovakia as it was the only country in the euro zone from that region; thus, the estimation of inflation in Slovakia is utilized as the benchmark. Under the univariate time series approach, time varying estimation was carried out to take into account the switching of regimes over the period considered. Instead of relying completely on inflation dynamics, the analysis was extended by including relevant macroeconomic indicators and by taking into account the same indicators when formulating the NKPC to provide additional control for nominal rigidities. The NKPC function provides further insight into the inflation process through assessing whether the inflation is predominantly forward or backward-looking.

The study sample covers monthly data from January, 1995 to May, 2010 for Poland, Hungary and Slovakia, from January 1994 to May, 2010 for the Czech Republic, and from January, 2002 to June 2010 for Romania. The shortest sample used, from January, 2003 to September, 2010, was for Serbia and is due to the late start in the transition process. The empirical results of this study supplement those existing for CSEECs: the sample period was extended by covering the recent years of higher inflation rates and by including Romania and Serbia, which were not previously included within the scope of study. To facilitate comparison with the results drawn from the monthly data of this inquiry, as well as the results found in the literature, quarterly data was utilized for the Czech Republic, Poland, Hungary and Slovakia.

The quarterly data covered the period from the first quarter, 1995 to the fourth quarter, 2010.

The remainder of the paper is structured as follows. Section 1 presents the basic macroeconomic indicators for selected countries and Section 2 provides a short overview of the empirical literature on inflation persistence in CSEECs. Some methodological issues are discussed in Section 3, while the findings derived from the univariate methods are presented in Section 4 and the results based on estimating the NKPC in Section 5. Section 6 summarizes the work carried out and the conclusions drawn therefrom. Additionally, the data used for this study is presented in the Appendix.

1. Basic Macroeconomic Indicators for Selected Countries

This study covers six CSEE countries which adopted inflation targeting as their monetary policy framework. In 2009 Slovakia entered the euro zone, such that with the exception of Serbia, all of the selected countries are EU members (Romania joined the EU in 2007 and others in 2004). Details of these countries in terms of area, population, GDP per capita, average monthly wages, average annual inflation rates, etc. are presented in Table 1.1. Among these, Poland is the largest country with a population of 38.2 million, while Slovakia is at the other end with a population of 5.4 million. In terms of GDP per capita (adjusted for purchasing power parity) and average monthly wages, Romania and Serbia's numbers are sufficiently lower than the other countries of this study (e.g., Serbia's GDP per capita level is less than half that of Poland and the same also holds true for average wages). Additionally, Romania and Serbia have experienced much higher annual inflation rates since the mid 1990s. It is also important to note that Slovakia, the Czech Republic, and Hungary are highly integrated countries in terms of international trade (the trade openness ratio exceeds 100 percent). During the period of interest, all of the countries were characterized by exchange rate flexibility.

Table 1.1 Summary Indicators of the CSEE Countries Covered by this Study

	Unit	Period	Czech Republic	Hungary	Poland	Romania	Serbia	Slovakia
Population	in millions	2009	10.49	10.01	38.17	21.48	7.32	5.42
Area	000' sq km		78.9	93.0	313.9	238.0	102.0	49.0
GDP in euros	in billions	2009	137.16	92.94	310.49	115.87	31.45	63.33
GDP per capita	EUR at PPP	2009	18900	14800	14300	10700	8700	16900
Growth in real GDP, ave.	in %	1995-1999	2.0	3.3	6.0	-0.2	1.6	4.3
		2000-2004	3.2	4.6	3.2	5.3	5.3	3.9
		2005-2009	3.5	0.6	4.7	3.7	4.2	5.4
Gross monthly wages, ave.	EUR	2009	889	713	717	435	470	745
Annual inflation, ave.	in %	1995-1999	7.9	18.9	16.3	66.2	52.5	7.8
		2000-2004	2.7	7.1	4.4	26.0	40.6	7.7
		2005-2009	4.2	5.1	2.8	6.8	11.2	3.0

Trade openness (import+export)/GDP, ave.									
	of goods	in %	2000-2007	118.6	119.6	60.0	65.3	52.1	136.4
	of goods and services	in %	2000-2007	137.1	141.8	70.8	76.0	62.5	155.4
ERM-II entry date (planned or actual)				No date	No date	No date	No date	No date	November 2005
Date of IT adoption				January 1998	Jun 2001	October 1998	August 2005	September 2006	January 2005
Exchange-rate system				Managed float	Fixed with band to euro	Free float	Managed float	Managed float	in EMR II intr. of euro in 2009

Source: European Bank for Reconstruction and Development - EBRD (2010); The Vienna Institute for International Economic Studies - WIIW (2010).

2. Previous Literature on Inflation Persistence in CSEECs

Inflation persistence in CSEECs has been studied by several scholars whose results are summarized in Table 2.1.

Table 2.1 Overview of Selected Studies on CSEECs

Literature	Countries	Sample	Approach	Results
Babetski et al. (2008)	Czech Republic	94m1-05m12	Micro analysis	Inflation seems to be less persistent after adoption of IT.
Konieczny and Skrzypacz (2005)	Poland	90m1-96m12	Micro analysis	There is a high degree of rationality among price setters.
Coricelli and Horvath (2006)	Slovakia	97m1-01m12	Micro analysis	The price dispersion is higher while persistence is lower in the non-tradable sectors.
Darvas and Varga (2007)	Hungary, US and Euro area	76q02-05q04	Univariate time series	Inflation persistence tends to be higher in times of high inflation. It is higher in Hungary than in US and EU.
Menyhert (2008)	Hungary	95q01-06q02	Structural time series	Inflation is determined equally by past inflation and forward-looking expectations.
Franta et al. (2007)	Czech Republic, Hungary, Poland, Slovakia and EU12	93q01-06q01	Univariate and structural time series	Inflation persistence is comparable to that in the current euro area.
Hondroyannis et al. (2008)	Euro area and 7 new EU members: Czech Rep., Hungary, Latvia, Lithuania, Poland Slovakia and Slovenia	95q01-05q03	Structural time series	For both, the Euro area and new member inflation has a backward-looking element, but it is more important component in new member states.
Franta et al. (2010)	New EU member states and euro zone countries	93q1-98q1	Univariate time series	Inflation persistence is comparable to that in euro zone in one group of the new member states, while it remains high in another group.

Source: The authors.

Early research on this topic was based on micro data analyses for the Czech Republic (Jan Babetski, Fabrizio Coricelli, and Roman Horvath 2008), Poland (Jerzy D. Konieczny and Andrzej Skrzypacz 2005) and Slovakia (Coricelli and Horvath

2006). Monthly data covering the period up until to 2005 was used for the sample of the aforementioned studies. Persistence among different price indices was measured for the overall CPI inflation in the Czech Republic, which revealed that inflation was less persistent after the adoption of inflation targeting. The magnitude of inflation persistence in Hungary was estimated by the univariate time series method for the period ending in 2005 and by the structural time series method for the period ending in mid-2006 (Zsolt Darvas and Balazs Varga 2007; Balint Menyhert 2008). It is found that during the study period, inflation persistence in Hungary was higher than that in the US and euro area and that inflation dynamics was determined as much by past inflation as by forward-looking expectations. Contrary to individual case studies, Michal Franta, Branislav Saxa, and Katerina Šmidkova (2007) covered several countries from CSEE along with the euro area based on a sample that ended in the first quarter of 2006, as did George Hondroyannis, P.A.V.B. Swamy, and George S. Tavlas (2008) based on a sample that ended in the third quarter of 2005. While Franta, Saxa, and Šmidkova (2007) estimated different univariate time series methods and the NKPC model; Hondroyannis, Swamy, and Tavlas (2008) considered only the NKPC model. A similar univariate analysis was performed by Franta, Saxa, and Šmidkova (2010) for an extended set of countries with a sample that ended in the first quarter of 2008; the main results indicate that inflation persistence was generally comparable to that in the euro area where backward behavior was dominant in inflation dynamics. However, there are countries among those recently admitted to EU that were characterized by relatively high inflation persistence. In addition, when time-varying specification of the NKPC was estimated, the role of lagged inflation diminished significantly.

The results reviewed here were obtained from the analysis of quarterly data of CPI and/or core inflation. However, it was the annualized quarterly inflation rate that was actually used in the empirical studies to measure the persistence. To clarify the point, if P_t represents the price index at quarter t , then the annualized inflation rate is derived in one of the following ways:

$$\pi_t = \begin{cases} I & 100(\log P_t - \log P_{t-4}) \\ II & 400(\log P_t - \log P_{t-1}) \end{cases}$$

Inflation rate I can be described as an aggregate derived from the quarterly data, thus:

$$I(\log P_t - \log P_{t-4}) = \log\left(\frac{P_t}{P_{t-1}}\right) + \log\left(\frac{P_{t-1}}{P_{t-2}}\right) + \log\left(\frac{P_{t-2}}{P_{t-3}}\right) + \log\left(\frac{P_{t-3}}{P_{t-4}}\right)$$

This determination of the inflation rate might make the measurement of persistence invalid when univariate time series methods are used (Paya, Duarte, and Holden 2007). More specifically, the inflation rate calculated as temporal aggregates from the actual highest frequency data available (quarterly or monthly) might cause upward bias in estimating inflation persistence.

Inflation rate II provides an adequate measure of annualized inflation rate only if the ratios shown below are approximately equal:

$$\left(\frac{P_t}{P_{t-1}} \right), \left(\frac{P_{t-1}}{P_{t-2}} \right), \left(\frac{P_{t-2}}{P_{t-3}} \right), \left(\frac{P_{t-3}}{P_{t-4}} \right)$$

In the economies that were characterized by significant regime changes, such as that of CSEE region, it is highly unlikely that quarterly inflation rates have equal ratios of the types shown above, thus the appropriateness of calculating the inflation rate this way is questionable. While such a computation does not cause bias in measuring persistence with the univariate methods, it can induce problems in the structural approach which evaluates the contribution of explanatory variables on inflation dynamics.

3. Short Overview of Methodology Applied

The results of this study are based on the seasonally adjusted monthly CPI index, denoted as P_t , where the monthly inflation rate was calculated as a percent, such that $\pi_t = 100(\log P_t - \log P_{t-1})$. In the baseline method of time series analysis, the persistence is measured by the sum of autoregressive coefficients, $\sum_{i=1}^p \phi_i$ from the autoregressive model of order p :

$$\pi_t = \alpha_0 + \sum_{i=1}^p \phi_i \pi_{t-i} + e_t,$$

which can be rewritten as

$$\pi_t = \alpha_0 + \rho \pi_{t-1} + \sum_{i=1}^{(p-1)} \delta_i \Delta \pi_{t-i} + e_t \quad (3.1)$$

such that the parameter $\rho = \sum_{i=1}^p \phi_i$ contains information about the sum of autoregressive parameters and thus provides the measure of inflation persistence. The error term, e_t , is Gaussian white noise and $\delta_i = -\sum_{j=1+i}^p \phi_j$.

This specification can be modified in a number of ways to provide for possible regime changes and nonlinearity in the inflation rate. In fact, several empirical papers on inflation persistence in transition economies (for example, Darvas and Varga 2007; Franta, Saxa, and Šmidkova 2007; Franta, Saxa, and Šmidkova 2010) argue that linear specification is not rich enough to capture the true dynamics in the inflation rate and that models with time-varying mean are preferable to describe the changes in mean due to switches in monetary policy regime and regulations of price administration. For the same reason, models with time-varying persistence and variance emerge as the natural candidates to estimate inflation persistence in CSEECs. For this study, the authors employed the Markov-switching autoregressive model (MS model), which assumes that mean, variability, and persistence of inflation randomly change between two regimes. The relevant specification is of the following form (James D. Hamilton 1989; Hamilton 1990):

$$\pi_t = (\alpha_0 + \alpha_1 S_t) + (\rho + \rho_1 S_t) \pi_{t-1} + \delta_1 \Delta \pi_{t-1} + \dots + \delta_{p-1} \Delta \pi_{t-p+1} + e_t \tag{3.2}$$

The switching mechanism of the economy is controlled by the state variable, S_t , which is an unobserved random variable, following a first order Markov chain defined by transition probabilities between two states. The full matrix of transition probabilities for two states reads as follows:

State at $t+1$	Condition at t	
	$S_t=0$	$S_t=1$
$S_{t+1}=0$	$q=p_{0/0}$	$f=p_{0/1}$
$S_{t+1}=1$	$p_{1/0}$	$p_{1/1}$

Note: $p_{ij}=P(\text{Regime } i \text{ at } t+1/\text{Regime } j \text{ at } t)$.
 Probabilities in each column aggregate to one.

Source: The authors.

Two states of economy, one as state 0 and the other as state 1, are denoted. The economy is in state (regime) 0 if $S_t=0$ and in state (regime) 1 if $S_t=1$.

The parameters α_1 and ρ_1 capture the changes respectively in the mean of inflation and in the persistence of a shock to inflation during regime 1 relative to regime 0. It is also assumed that uncertainty of inflation, measured by the error variance, changes across two regimes. In regime 0, the variability is $\text{var}(e_t) = h_0^2$, and in regime 1 $\text{var}(e_t) = (h_0 + h_1)^2$. Parameter h_1 describes the changes of standard deviation in regime 1 relative to regime 0.

The parameters of the Markov-switching model can be estimated by the method of maximum likelihood. The likelihood of the model can be evaluated efficiently using the filtering procedure of Hamilton (1990).

The magnitude of inflation persistence can be measured from structural time series modeling, i.e. from the modified structural Phillips curve estimated in the manner suggested by Jordi Gali and Mark Gertler (1999). The closed form version of the New Keynesian Phillips Curve (NKPC) reads as follows:

$$\pi_t = \alpha_b \pi_{t-1} + \alpha_f E(\pi_{t+1}) + \lambda mc_t + \text{error term} \tag{3.3}$$

indicating that inflation, π_t , is related to lagged inflation, π_{t-1} , expectations about future inflation, $E(\pi_{t+1})$, and deviation from average real marginal cost, mc_t . Assuming rational expectations, future actual inflation rate, π_{t+1} was used for expected future inflation. As shown by Gali and Gertler (1999), this specification can be rewritten, subject to the expected path of real marginal cost:

$$\pi_t = g_1 \pi_{t-1} + \frac{\lambda}{g_2 \alpha_f} \sum_{k=0}^{\infty} \left(\frac{1}{g_2} \right)^k E\{mc_{t+k}\} + \text{error term} \tag{3.4}$$

where g_1 and g_2 denote, respectively, the stable and unstable roots of the corresponding second order difference equations associated with (3.3) that are equal to

$$g_1 = \frac{1 - \sqrt{1 - 4\alpha_b\alpha_f}}{2\alpha_f}, g_2 = \frac{1 + \sqrt{1 - 4\alpha_b\alpha_f}}{2\alpha_f} \quad (3.5)$$

The parameter on real marginal cost, λ , can be represented, under certain restrictions, as:

$$\lambda = \frac{(1-\theta)(1-\alpha_f\theta)}{\theta} \quad (3.6)$$

where parameter $(1-\theta)$ measures the probability that each firm can reset its price in the given period. Based on the estimated values of λ and α_f , one can determine two different values of θ and hence of $(1-\theta)$ as a result of the quadratic equation solution. One of the values of $(1-\theta)$ lies between zero and one; this value represents the probability that prices will be adjusted over a certain period of time.

Several issues emerge in estimating the NKPC (3.3), of which the two most important ones are highlighted here briefly. First, there is no agreement over the choice of the estimation method because the specification assumes endogeneity of a variable and an AR(1) structure for an error term. Gali and Gertler (1999) advocate the general method of moments (GMM) for robustness of empirical results (Gali, Gertler, and David J. Lopez-Salido 2005) given the ongoing debate on the appropriateness of this method. Apart from the GMM, parameters of the NKPC were estimated in different empirical studies by using either the full or limited information maximum likelihood-LIML (for example, Jesper Linde 2005). Alternatively, the baseline specification can be slightly modified to resolve issues of autocorrelation, allowing the two stage least squares (TSLS) method to be employed (for example, Gunnar Bardsen, Eilev S. Jansen, and Ragnar Nymoen 2004; Chengsi Zhang and Joel Clovis 2010). Second, it remains unclear as to which variables can be used as proxies for real marginal costs, considering that the variables are often measured by real unit labor costs which captures the impact of wages and productivity on inflation. Real output gap was also used in many empirical studies, although Gali and Gertler (1999) argue against it being a valid real driving force for inflation. Some recent studies suggest that, for small and open economies, the terms of trade should play an important role in driving inflation.

The set-up of the NKPC appears as a demanding task in the CSEECs as many relevant time series are either incomplete or unavailable. Thus, most of the existing empirical results are based on samples which included the turbulent episodes at the beginning of the transition process, during which price liberalization dominated the inflation dynamics over the whole period under consideration. In addition, in comparison to real labor unit costs, wage inflation appears as a natural candidate for the inflation driving force, especially given the well-known sensitivity of inflation in these countries to wage shocks. Alternatively, real broad money growth could be used because in some countries it encompasses the effects of aggregate demand, mostly driven by capital inflows, on inflation.

4. Empirical Findings from Univariate Approach

In this section the MS models were estimated. The superiority of this specification to the standard linear autoregressive model was confirmed for each economy by the linearity test that was based on the likelihood ratio statistic between the estimated MS model and the derived linear model.¹ Therefore, Markov-switching representation provides an adequate set-up for obtaining the statistical measure of inflation persistence. In addition, it enables distinction between the regimes characterized by different levels of inflation persistence as well as an assessment of how inflation persistence is associated with its mean and uncertainty.

Hungary

The two-state MS model fits well with the dynamics of the monthly inflation rate in Hungary for the period January, 1995-May, 2010. As shown in Table 4.1, two different inflation persistence regimes have been detected. Regime 0 has lower inflation persistence characterized by an estimated magnitude of $\hat{\rho} = 0.38$. This regime also contains a lower expected average inflation rate, $0.20\% / (1-0.375) = 0.32\%$ and a lower variability (measured by standard deviation), $\hat{h}_0 = 0.07$. Regime 1 has higher inflation persistence characterized by an estimated magnitude of $\hat{\rho} + \hat{\rho}_1 = 0.82$. During regime 1, the inflation rate exhibited a higher mean value, $0.10\% / (1-0.823) = 0.56\%$ and a higher variability, 0.30. Statistically, there is significant difference between the two parameters of inflation persistence.

The probability, q , of remaining in the regime of lower persistence, while already in that regime, is 0.55. The probability of remaining in the regime of higher persistence, while already in that regime, is 0.92, thus implying that the probability, f , of switching from the regime of higher to the regime of lower persistence is small and equal to 0.08.

The estimates are depicted in Figure 4.1. It is noticed that the economy was predominantly (85% of the time) in regime 1, a regime of relatively high inflation persistence. This also holds for the last years of the sample which is characterized by a relatively high inflation rate.

The simple linear autoregressive model suggests that the inflation persistence was of moderate size, 0.78. As this specification was improved significantly by the estimates obtained herein, it is concluded that inflation persistence remained relatively high in Hungary with the prevailing estimate being 0.82. Furthermore, it is closely related to sub-periods of higher inflation mean and variability.

The study sample (since mid 2001) covers inflation targeting policy in Hungary. The occurrence of one or another inflation persistence regime was compared with the dates on which the interest rates were changed by the Hungarian Central Bank for that period.² In the long-run, the interest rate significantly declined, although positive changes were made over the short-run. The lower inflation persis-

¹ Oxmetrics 6.1 package (Jurgen Doornik 2009) was used in estimating Markov switching models. Estimations of NKPC were derived later on from EvIEWS 7.0 package.

² Information on interest rate changes are available at central banks' websites.

tence regime, identified by the present authors' analysis, was associated predominantly with the sub-periods denoted by a decrease in the interest rate. Such a trend in the interest rate signals the economic agents of a likely fall in the inflation rate. This may explain why inflation persistence remains at lower level in the sub-intervals found. On the other side, the higher inflation persistence regime is accompanied by intervals during which, on the short-run, the interest rate either remained stable or changed in both directions. However, a detailed assessment of the relationship between inflation dynamics and monetary policy actions is beyond the scope of the paper.

Table 4.1 Estimated Inflation Model for Hungary (January, 1995-May, 2010)

Parameter	Estimate	t-ratio
α_0	0.204	4.51
$(\alpha_0 + \alpha_1)$	0.101	2.47
ρ	0.375	7.04
$(\rho + \rho_1)$	0.823	19.30
h_0	0.070	2.56
$(h_0 + h_1)$	0.300	16.60
q	0.550	2.05
f	0.080	2.52
δ_1	-0.385	-7.41
δ_2	-0.196	4.71

Linearity test : $\chi^2_5 = 16.53(0.005)$; *Box-Ljung* $Q(36) \chi^2_{36} = 34.22(0.55)$,
ARCH 1 $F(1,165) = 2.07(0.15)$; *Normality* $\chi^2_2 = 5.80(0.06)$

Five impulse dummy variables were included that have non-zero value 1 for the following months: 1997:6, 2000:7, 2004:1, 2006:9 and 2009:7.

Source: Authors' estimations.

Note to Tables 4.1-4.6, 4.8 and 4.9.

Linearity test stands for the likelihood ratio statistic of this model and its implied autoregressive specification. Under the null, it has χ^2 distribution with the number of degrees of freedom equal to the number of parameters estimated in addition to autoregressive model. Box-Ljung Q (36) and Q (12) tests were used to test for the serial correlation up to order 36 and 12 respectively in scaled residuals. Under the validity of the null hypothesis, the statistic has asymptotically χ^2 distribution. ARCH 1 test was applied to test for the autoregressive heteroskedasticity of order one in scaled residuals. Under the null hypothesis that heteroskedasticity is absent, this statistic has F distribution. Normality stands for the Doornik-Hansen residual normality test that

under the null has χ_2^2 distribution. It was also applied on scaled residuals. Number in parentheses is the corresponding p-value.

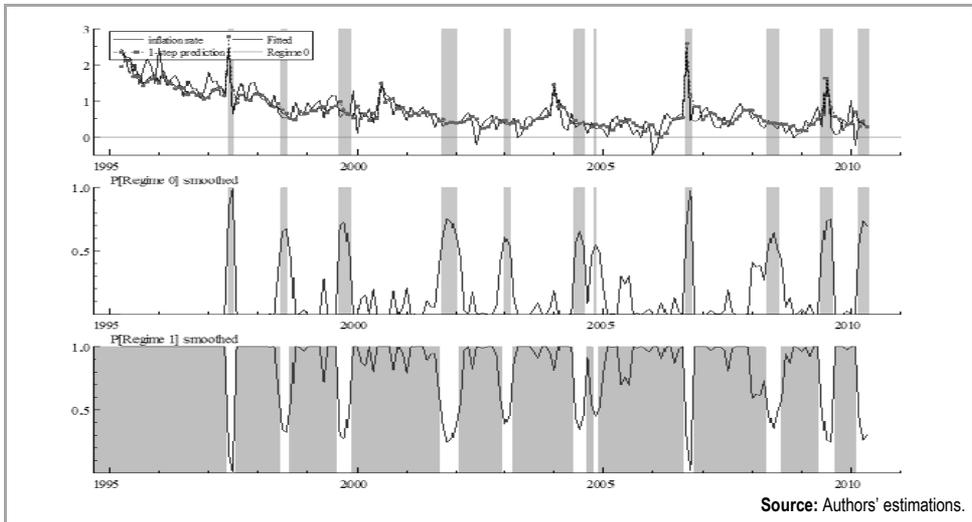


Figure 4.1 Two Regimes of Inflation Persistence in Hungary

Czech Republic

The two-state MS model was also used to estimate the monthly inflation rate in the Czech Republic for the period January, 1994-May, 2010 (Table 4.2). Regime 0 had higher inflation persistence, which was estimated to be 0.57. During this regime, the inflation rate exhibits a higher mean, $0.14\% / (1-0.570) = 0.33\%$ and variability, 0.39. Regime 1 can be considered a regime of lower inflation persistence, which was estimated to be 0.46. During regime 1, the inflation rate had a lower mean value, $0.14\% / (1-0.460) = 0.26\%$, and variability, 0.15. Although the two estimates of persistence do not differ significantly, the two regime analysis was continued because of significant differences in the inflation mean and variability.

In the Czech Republic, the inflation targeting policy has been in place since January 1998. The two inflation regimes were compared with the dates on which the interest rate was changed. Because of the sharp decreases in the interest rate occurring in 1998-1999, the period from 2000 onward was considered. It appears that the regime of lower inflation persistence, mean, and variability prevails predominantly during intervals of either stable or decreasing interest rates. On the other side, the regime of higher inflation persistence, mean, and variability is associated with intervals when the interest rate exhibited either a downward or upward trend in the short-run, although sub-intervals of positive changes in interest rate are more frequent. It seems that the declining short-term trend in the interest rate is a potential indicator against a rise in inflation, thus ultimately reducing inflation persistence.

Visual inspection of Figure 4.2 shows that the two regimes are equally split with 49.75% months in regime 0 and 50.25% months in regime 1. However, when the data of the last two years (25 months) was considered, it is observed that this economy was predominantly (20 months) in the regime of lower inflation persis-

tence, mean, and variability.

Table 4.2 Estimated Inflation Model for Czech Republic (January, 1994-May, 2010)

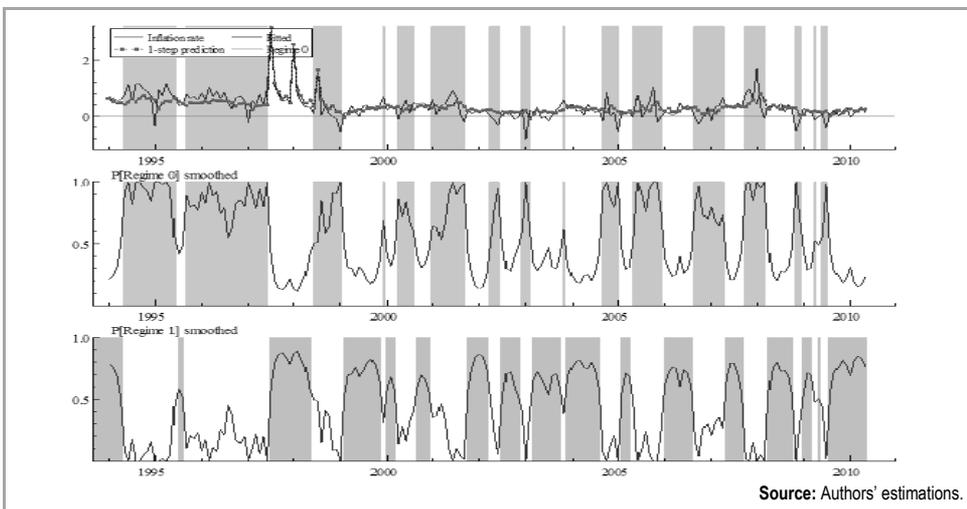
Parameter	Estimate	t-ratio
α_0	0.135	2.22
$(\alpha_0 + \alpha_1)$	0.138	4.03
ρ	0.570	5.54
$(\rho + \rho_1)$	0.460	6.73
h_0	0.390	9.30
$(h_0 + h_1)$	0.150	4.26
q	0.800	6.28
f	0.260	2.09
δ_1	-0.259	-4.33
δ_2	-0.123	-2.48
δ_3	-0.110	-2.54

Linearity test: $\chi^2_5 = 17.21(0.00)$; *Box-Ljung* $Q(36) \chi^2_{36} = 47.54(0.10)$,

ARCH 1 $F(1,181) = 0.212(0.64)$; *Normality* $\chi^2_2 = 1.15(0.56)$

Three impulse dummy variables were included that have non-zero value 1 for the following months: 1997:7, 1998:1 and 1998:7.

Source: Authors' estimations.



Source: Authors' estimations.

Figure 4.2 Two Regimes of Inflation Mean and Variability in Czech Republic
Poland

Based on the sample for January, 1995 - May, 2010, the monthly inflation rate was modeled within the two-state MS specification and the results are summarized in Table 4.3. It is found that regime 0 has higher inflation persistence (estimate 0.89) and regime 1 lower inflation persistence (estimate is 0.61). The parameters differ significantly as suggested by the Wald test: $\chi_1^2 = 3.70(0.04)$.

The probability, q , of staying in the regime of higher persistence, while already in that regime, is 0.85. The probability, f , of switching from the regime of lower persistence to that of higher persistence is 0.40.

The results are depicted in Figure 4.3. It is found that, for almost 80% of the time, the economy was in regime 0, characterized by a higher level persistence. The remaining 20% was in regime 1, ranging from periods of one to four months, characterized predominantly by non-standard values. Owing to these transitory shocks, inflation uncertainty is estimated to be higher in the regime of lower persistence. Thus, there is no regularity observed between the switches in persistence with the changes in central bank policy.

A relatively high degree of persistence is associated with the last two years of estimation, which, as in the case of Hungary, emphasizes the potential problem with the ongoing economic instability at the level of inflation persistence.

Table 4.3 Estimated Inflation Model for Poland (January, 1995 -May, 2010)

Parameter	Estimate	t-ratio
α_0	0.007	0.38
$(\alpha_0 + \alpha_1)$	0.325	2.29
ρ	0.893	28.90
$(\rho + \rho_1)$	0.611	4.41
h_0	0.130	9.95
$(h_0 + h_1)$	0.510	8.68
q	0.850	17.60
f	0.400	4.12
δ_1	-0.225	-3.60
δ_2	-0.132	-2.79
δ_3	-0.092	-2.19
Linearity test : $\chi_5^2 = 83.91(0.00)$; Box - Ljung $Q(36) \chi_{36}^2 = 40.57(0.28)$, ARCH 1 $F(1,168) = 3.021(0.09)$; Normality $\chi_2^2 = 1.94(0.38)$		

Source: Authors' estimations.

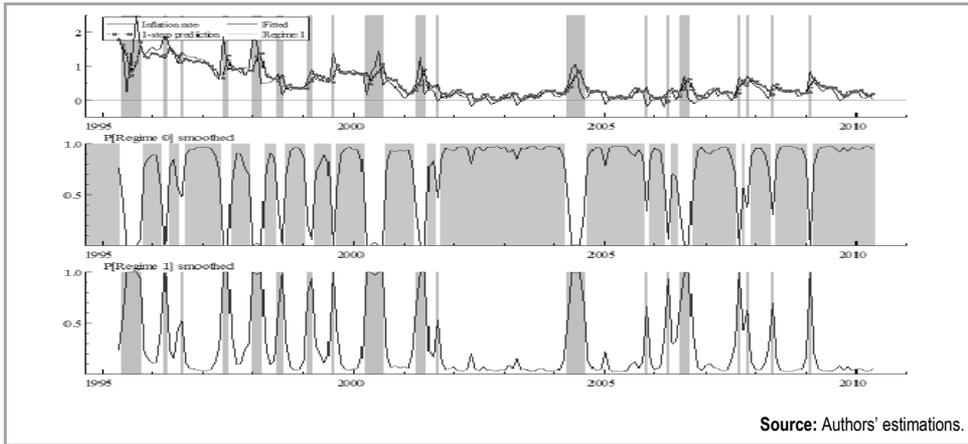


Figure 4.3 Two Regimes of Inflation Persistence in Poland

Romania

By using the sample for January, 2002-June, 2010 (Table 4.4), it is found that the two-state MS model performs well with regard to the monthly inflation rate in Romania. It is further found that regime 0 had higher inflation persistence (estimate: 0.73) and regime 1 lower inflation persistence (estimate: 0.52), with the measures of persistence differing significantly. The average duration of regime 0 is 1.79 months and occurs through 43.43% of the sample. The remaining 56.57% belongs to regime 1, which lasts on average 2.33 months (Figure 4.4). Furthermore, the economy moves from one regime to another frequently. In this case, higher inflation persistence episodes were associated with an inflation rate that exhibited a lower mean and less variability.

The probability of staying in the regime of lower persistence while already in that regime is $1-f=0.66$, while the probability of switching to the other regime is 0.34.

Table 4.4 Estimated Inflation Model for Romania (January, 2002-June, 2010)

Parameter	Estimate	t-ratio
α_0	0.092	2.49
$(\alpha_0 + \alpha_1)$	0.308	3.22
ρ	0.733	18.9
$(\rho + \rho_1)$	0.522	4.46
h_0	0.060	4.44
$(h_0 + h_1)$	0.400	10.20
q	0.410	2.80
f	0.340	3.69

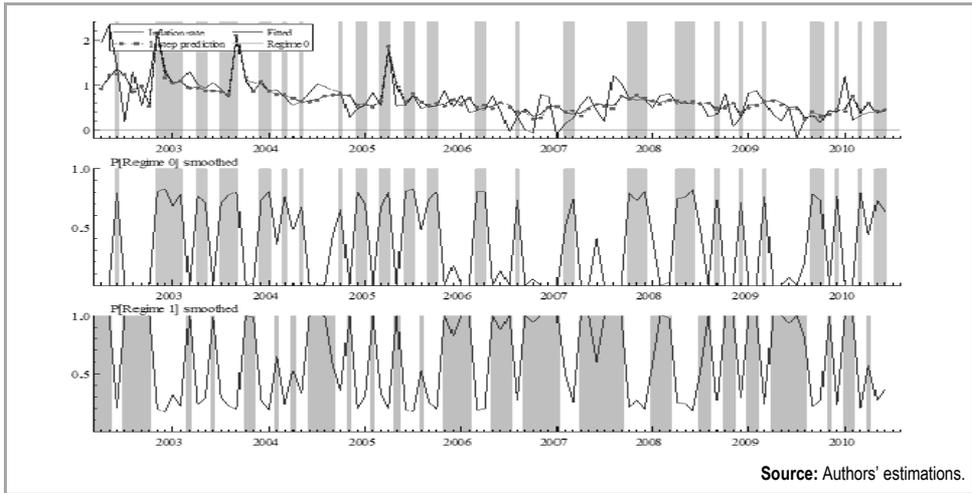
δ_1	-0.253	-5.66
δ_2	-0.204	-7.43

Linearity test : $\chi^2_5 = 31.93(0.00)$; Box – Ljung $Q(36) \chi^2_{36} = 26.04(0.89)$,

ARCH 1 $F(1,84) = 3.22(0.08)$; Normality $\chi^2_2 = 1.30(0.53)$

Three impulse dummy variables were included that had non-zero value 1 for the following months: 2002:11, 2003:9 and 2005:4.

Source: Authors' estimations.



Source: Authors' estimations.

Figure 4.4 Two Regimes of Inflation Persistence in Romania

Serbia

Inflation dynamics in Serbia were examined within the MS model. However, since Serbia entered the transition process at the end of 2000, the sample covered a shorter time interval than was employed for the other countries. To avoid shocks due to price liberalization and to ensure the adjustment of several price indices undertaken at the beginning of this process, the study sample covered the period starting in January 2003 and ending in September 2010.

Models with three regimes fit the Serbian inflation rate satisfactorily well. The estimated model is presented in Table 4.5 and the corresponding transition probabilities in Table 4.5a. Regime 0 occurred during 22.47% of the sample and is characterized by high inflation persistence (estimate: 0.81) lasting for an average duration of 1 month. Regime 1 was also associated with relatively high inflation persistence (estimate: 0.72). Identical to regime 0, it also took 22.47% of the sample with an average duration of 1.25 months. Inflation persistence was lowest (estimate: 0.48) in regime 2, and lasted, on the average, 1.58 months. Of the three regimes, inflation variability was the highest for regime 2 (see Figure 4.5).

Adequacy of the estimated model was confirmed by several specification tests. While simple AR model shows that inflation persistence is just the value found for regime two (0.48), MS modeling approach enables one to distinguish different episodes in inflation behavior. The Serbian monthly inflation rate exhibits, similar to Romania's results, frequent—almost monthly—changes of regimes. Such a behavior of inflation dynamics disables any reliable assessment of association between the regimes identified and the dates when the central banks of these countries changed the interest rate under the inflation targeting policy.

One can only conclude that these inflation rates are highly sensitive to unexpected random shocks. Furthermore, this is confirmed by the observation that inflation uncertainty was higher when the persistence was lower. In addition, given that regime 2 had lower persistence, there was almost equal probability (about 0.40) that the economy would stay in that regime as there was to the economy switching to regime 0 of higher persistence.

Table 4.5 Estimated Inflation Model for Serbia (January, 2003-September, 2010)

Parameter	Estimate	t-ratio
α_0 regime 0	-0.448	-1.67
α_0 regime 1	0.081	7.96
α_0 regime 2	0.533	5.27
ρ regime 0	0.815	5.46
ρ regime 1	0.719	63.50
ρ regime 2	0.479	6.15
h_0 regime 0	0.221	1.93
h_0 regime 1	0.014	5.10
h_0 regime 2	0.350	9.20
δ_1	-0.179	-14.10
δ_2	-0.191	-20.60
δ_3	-0.032	-4.48

Linearity test : $\chi_{10}^2 = 43.63(0.0)$; Box – Ljung $Q(36) \chi_{36}^2 = 26.50(0.88)$,

ARCH1 $F(1,165) = 0.02(0.90)$; Normality $\chi_2^2 = 4.12(0.13)$

Five impulse dummy variables were included to have non-zero value 1 for the following months: 2005:1, 2005:10, 2007:8, 2009:1 and 2010:8. In addition, transitory dummy with only non-zero values 1 and -1 for 2008:11 and 2008:10 respectively, was also introduced.

Source: Authors' estimations.

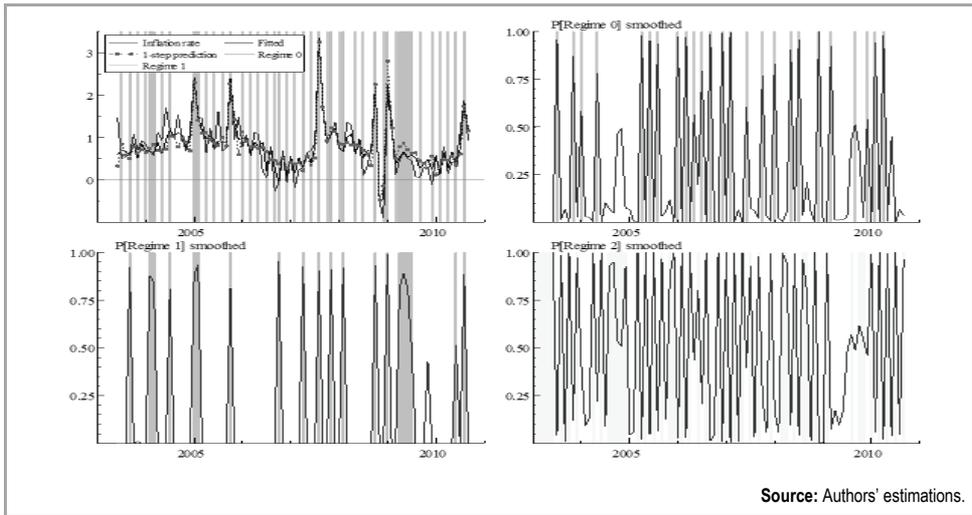


Figure 4.5 Three Regimes of Inflation Persistence in Serbia

Table 4.5a Estimated Transition Probabilities from Table 4.5

	$S_t=0$	$S_t=1$	$S_t=2$
$S_{t+1}=0$	0.00	0.00	0.40
$S_{t+1}=1$	0.19	0.18	0.22
$S_{t+1}=2$	0.81	0.82	0.38

Source: Authors' estimations.

Slovakia

As the final phase of univariate modeling, Slovakian monthly inflation was considered for the period January, 1995-May 2010. The estimated model shows (Table 4.6) one regime of inflation persistence and two regimes of inflation mean and variability. Estimated inflation persistence was 0.28. The regime of lower inflation mean covered 96% of the sample and had an average duration of 21.75 months. The remaining 4% had higher a mean inflation rate that lasted on average 1 month. As can be seen in Figure 4.6, the Slovakian economy has been in the regime of lower mean inflation since November 2005, which coincides with the start of inflation targeting.

Table 4.6 Estimated Inflation Model for Slovakia (January, 1995-May, 2010)

Parameter	Estimate	t-ratio
α_0	0.206	7.11
$(\alpha_0 + \alpha_1)$	0.914	14.90
ρ	0.283	5.81
h_0	0.201	16.50

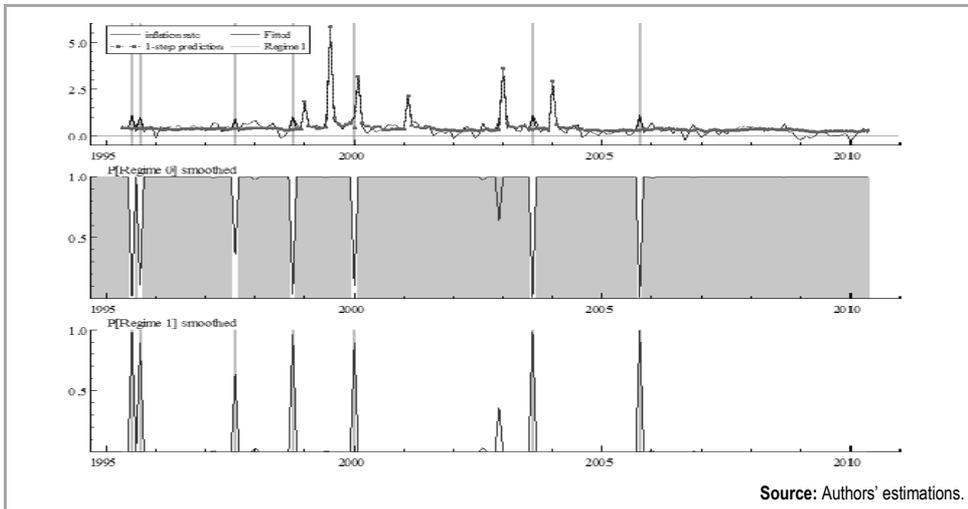
h_1	0.100	2.40
δ_1	-0.174	-3.87
δ_2	-0.130	-3.30
δ_3	-0.109	-3.21
δ_4	-0.064	-2.41

Linearity test : $\chi_3^2 = 16.70 (0.0)$; *Box - Ljung* $Q(36) \chi_{35}^2 = 42.590 (0.18)$,

ARCH $1 F(1,163) = 0.37 (0.54)$; *Normality* $\chi_2^2 = 0.85 (0.65)$

Four impulse dummy variables were included to have non-zero value 1 for the following months: 1999:1, 1999:7, 2000:2 and 2003:1. Also, two composite dummies were introduced such that the first one took non-zero value 1 for 2001:2 and 2000:2, and the second one for 2004:1 and 2003:1.

Source: Authors' estimations.



Source: Authors' estimations.

Figure 4.6 Two Regimes of Inflation Mean and Variance in Slovakia

The findings for the different countries are summarized in Table 4.7. The level of inflation persistence was by far the lowest for Slovakia and was relatively modest for the Czech Republic. It is only in these two economies that there was no significant difference in inflation persistence across regimes.

Inflation persistence remained at such a high level in Hungary and Poland that the regimes of higher inflation persistence dominated the sample considered. Similarly, the regimes of higher inflation persistence occurred more frequently in Romania and Serbia as well. In these two economies, the switching between the regimes was so frequent that the relatively lower inflation persistence was characterized by higher inflation uncertainty. This could also be a sign of a high sensitivity of the inflation rate to unexpected random shocks. The reaction of inflation to impulses may

manifest itself as a lower level persistence in one of the detected regimes, but at the cost of higher variability in the inflation rate.

In regards the actual rates of inflation, inflation persistence was estimated to be at the highest level in the countries that experienced the highest rates of inflation over the period considered (Serbia and Romania). Likewise, the lowest inflation persistence was detected for the economies that experienced the lowest inflation rates (Slovakia and the Czech Republic). Slightly higher inflation rates were experienced by Poland and Hungary, which were found to be characterized by relatively high inflation persistence.

Table 4.7 Estimated Inflation Persistence across Countries

Country	Estimated inflation persistence	
	Regime of lower persistence	Regime of higher persistence
Hungary	0.38	0.82
Czech Republic	0.46	0.57
Poland	0.61	0.89
Romania	0.52	0.73
Serbia	0.48	0.72/0.81
Slovakia		0.28

Source: Authors' estimations.

To check for the robustness of these results, MS model estimation was performed with the quarterly data of Hungary, the Czech Republic, Poland, and Slovakia. The sample covered the period from the first quarter, 1995 to the fourth quarter, 2010. The inflation rate represents the first difference between the log values of the quarterly CPI index (a seasonal adjustment was made and inflation is given in percentages). The results, reported in Tables 4.8 and 4.9, are again in favor of MS specification.

Table 4.8 Estimated Inflation Model for Hungary and Poland Based on Quarterly Data (1st quarter, 1995-4th quarter 2010)

Parameter	Estimate for Hungary	Estimate for Poland
α_0	-0.16 (-0.85)	0.09 (0.19)
$(\alpha_0 + \alpha_1)$	0.97 (8.63)	1.25 (42.80)
ρ	0.85 (9.64)	0.73 (14.50)
$(\rho + \rho_1)$	0.73 (16.90)	0.64 (50.90)
h_0	0.36 (4.79)	0.29 (8.80)
$(h_0 + h_1)$	0.20 (4.24)	0.05 (4.27)
δ_1	-0.14 (-2.07)	-
δ_2	-0.17 (-2.13)	-

δ_3	-0.28 (-3.15)	0.16 (7.35)
δ_4	-	-0.19 (-7.64)
δ_5	-	-0.16 (-6.80)
Time in regime of higher inflation persistence	60%	81%

Hungary

Linearity test : $\chi_5^2 = 10.95(0.05)$;

Box – Ljung Q(12) $\chi_{12}^2 = 13.72(0.32)$,

ARCH 1 F(1,44) = 0.88(0.35); Normality $\chi_2^2 = 2.32(0.31)$

One impulse dummy variable was included that had non-zero value 1 for the third quarter of 2000.

Poland

Linearity test : $\chi_5^2 = 34.02(0.00)$;

Box – Ljung Q(12) $\chi_{12}^2 = 8.68(0.73)$,

ARCH 1 F(1,43) = 0.23(0.63); Normality $\chi_2^2 = 1.33(0.51)$

Composite impulse dummy variable was included that had two non-zero values: 1 for the second quarter of 1998 and -1 for the first quarter of 1998.

Note to Tables 4.8 and 4.9: t-ratios are reported in parentheses below estimates.

Source: Authors' estimations.

Table 4.9 Estimated Inflation Model for Czech Republic and Slovakia Based on Quarterly Data (1st quarter, 1995-4th quarter, 2010)

Parameter	Estimate for Czech Republic	Estimate for Slovakia
α_0	-0.30 (-3.71)	0.59 (4.80)
$(\alpha_0 + \alpha_1)$	0.53 (7.01)	1.24 (5.98)
ρ	0.38 (8.23)	0.22 (2.31)
h_0	0.15 (3.24)	0.48 (8.34)
$(h_0 + h_1)$	0.42 (7.79)	0.45 (5.66)

Czech Republic

Linearity test : $\chi_3^2 = 6.40(0.09)$;

Box – Ljung Q(12) $\chi_{11}^2 = 15.43(0.16)$,

ARCH 1 F(1,45) = 0.25(0.62); Normality $\chi_2^2 = 1.77(0.41)$

Three impulse dummy variables were included that had non-zero value 1 for the following three quarters: third quarter of 1997, first quarter of 1998 and first quarter of 2008.

Slovakia

Linearity test : $\chi_3^2 = 9.51(0.02)$;

Box – Ljung Q(12) $\chi_{11}^2 = 7.67(0.74)$,

ARCH 1 F(1,45) = 0.97(0.33); Normality $\chi_2^2 = 3.36(0.19)$

Two impulse dummy variables were included that took non-zero value 1 for the following two quarters: the third quarter of 1999 and the first quarter of 2003. In addition, two composite dummies were present. The first one had two non-zero values 1 for the first quarter of 2000 and the second quarter of 2001. The second one took non-zero value 1 for the first quarters of 2003 and 2004.

Source: Authors' estimations.

The two-state MS model reveals once again high inflation persistence for Hungary and Poland. As in the case of the monthly data results, the higher inflation persistence regime (estimated at 0.85 for Hungary and 0.73 for Poland) occurred throughout a substantial portion of the sample. The level of inflation persistence was estimated to be much smaller in the Czech Republic and Slovakia. This inference was drawn from a statistically valid specification of one regime of inflation persistence (estimates are respectively 0.38 and 0.22), but two regimes of inflation mean and variability. This is also in accordance with the results that were obtained with the monthly data. Both of the data frequencies identify similar inflation episodes throughout the sample period. Additional evidence in favor of low inflation persistence in the latter two economies is the absence of short-run dynamics in the estimated model, suggesting the existence of mild inertia in the data. On the other hand, the models for Hungary and Poland needed relatively rich, short-run dynamics for valid statistical performance.

These findings are now comparable to the existing empirical results that were based on the quarterly data. The research data published recently by Franta, Saxa, and Šmidkova (2010) shows that the new EU member states do not have the same level of inflation persistence; rather, relatively higher inflation persistence was estimated for Hungary and Poland and smaller inflation persistence for the Czech Republic and Slovakia. Therefore, the present results, albeit derived from different frameworks, led to similar conclusions.

5. Empirical Findings from Structural Approach

The analysis of inflation dynamics in Section 4 requires further investigation in regard to the determination of its main sources. Therefore, the discussion now turns to the structural modeling for estimating inflation persistence within the NKPC framework. By this, one compares the results obtained from the univariate set-up with the ones obtained by including other macroeconomic variables in the model.

Like most previous empirical studies, the GMM was applied in the NKPC set-up. To check for the robustness of the results, the LIML method, yielding similar results, was also employed. The estimates of the standard errors were adjusted for serial correlation and heteroscedasticity by using the commonly applied HAC correction.

The estimation of the NKPC depends on the selection of the instrumental variables as well as the inflation forcing variable. For the CSEECs, this selection was to some extent influenced by poor data-availability, especially at the level of monthly frequency. This influence on selection was particularly more pronounced for Slovakia and Romania. Moreover, some relevant time series were only available for certain subsamples. Because of this limitation, a set of instrumental variables was chosen according to the suggestion of Gali and Gertler (1999). Additionally, the peculiar features of economies in the CSEECs were also taken into account.

After trying different combinations of instruments and proxies for the real sector forcing variables, the most satisfactory regression estimates are presented for each country in Tables 5.1-5.6. From the econometric results, it is concluded that the inflation dynamics in the CSEECs are well captured by the NKPC.

Several diagnostic tests were performed to evaluate the NKPC specification. To check for the potential weaknesses of the instruments, the F-statistic was presented from the first stage regression of inflation (π_t) against the instruments set. The rule of thumb advocated by Douglas Staiger and James H. Stock (1997) is that a value bigger than 10 indicates “no weak instruments” in the case of one endogenous regressor. Additional evidence of the GMM validity is the Hansen J-test of over-identifying restrictions, which examines whether the orthogonality conditions for instruments have been met. The p-value of the Hansen J-statistic is denoted as p-over in the tables below. It should be noted that J-test had deficient small-sample properties and could prove deceptive when the estimated residuals manifest serial correlation (Bardsen, Jansen, and Nymoen 2004; Zhang and Clovis 2010).

By using Polish and Slovakian data, previous empirical studies could not find empirical support for the NKPC. Moreover, from the New Keynesian perspective, this is the first analysis of inflation dynamics in Romania and Serbia.

Hungary

The NKPC model for Hungary was estimated with real gross wage inflation as the real marginal cost variable (Table 5.1). This variable has a positive and significant impact on the inflation dynamics. Likewise, weights on past and future inflation are highly significant. The estimates reached are robust to the change of estimation method. Parameter α_b was estimated to be in the range 0.41-0.45 and parameter α_f in interval 0.48-0.51. Expectations about future inflation contribute slightly more to inflation persistence than the backward-looking behavior. As reported below in the case of Poland, both of the parameters aggregate to 1.

The present estimates are not completely comparable to those of Menyhert (2008) where the annualized quarterly core inflation rate was applied for the sample that ended in 2006. It was found that the portions of backward and forward looking behaviors are about the same and close to 0.5. Based on this finding, one may argue that the backward-looking element of inflation dynamics has been diminished recently, while the forward-looking element has remained at about the same level.

Table 5.1 Estimation of NKPC for Hungary Using Real Gross Wage Inflation (April, 1995-July, 2010)

	$\hat{\alpha}_b$	$\hat{\alpha}_f$	$\hat{\lambda}$	F-stat	p-over
GMM	0.446 (0.065)	0.479 (0.081)	0.010 (0.039)	17.667	0.199
LIML	0.413 (0.082)	0.508 (0.092)	0.074 (0.100)	5.431	

Note to Tables 5.1-5.6 and 5.8: GMM results were obtained by the TSLs estimator; HAC-robust standard errors are shown in parenthesis; p-over denotes p-value of Hansen J-test; the F-statistic is a measure of the instruments strength (the joint significance test of the instruments in the first stage of regression).

Source: Authors' estimations.

Note 1: The full set of instruments includes lagged inflation of order 2 and 11, two lags of real gross wage inflation (growth), output gap and real broad money growth. The output gap is based on detrended log GDP. Monthly data of GDP was obtained by the disaggregation of quarterly measured GDP (seasonally adjusted). For a definition of these instruments, see Appendix.

Poland

Real broad money growth was employed in the NKPC specification for Poland (Table 5.2) and it was found to have a positive and significant impact on the inflation dynamics. Parameters on past and future inflation are also highly significant. Parameter α_b was estimated to be 0.41 by GMM and 0.35 by LIML, while the estimate for α_f is around 0.48 in both of the equations. Accordingly, the forward-looking behavior seems to have a slightly larger impact than the backward-looking activity.

Table 5.2 Estimation of NKPC for Poland Using Real Broad Money Growth (April, 1995-July, 2010)

	$\hat{\alpha}_b$	$\hat{\alpha}_f$	$\hat{\lambda}$	F-stat	p-over
GMM	0.412 (0.071)	0.484 (0.071)	0.055 (0.024)	22.235	0.524
LIML	0.354 (0.094)	0.487 (0.086)	0.118 (0.069)	11.032	

Note: The full set of instruments includes lagged inflation of order 2 and 11, two lags of real gross wage inflation (growth) and real broad money growth.

Source: Authors' estimations.

Czech Republic

The inflation driving variable in the Czech Republic is real broad money. All the variables included in the NKPC were found to be significant with a positive impact (the only exception is real broad money within LIML estimation, Table 5.3). The estimate for past inflation is in the interval 0.24-0.36, while the estimate for the forward-looking weight ranges from 0.21 to 0.31 depending on the method employed. While the GMM suggests a slightly higher impact of the forward-looking element, the LIML indicates just the opposite. The authors' estimate of the forward-looking behavior is close to what Franta, Saxa, and Šmidkova (2007) estimated for the impact of backward-inflation, which is high (0.42-0.47). It is possible that during recent years there was reduction in the backward-looking weight.

Table 5.3 Estimation of NKPC for Czech Republic Using Real Broad Money (March, 1996-July, 2010)

	$\hat{\alpha}_b$	$\hat{\alpha}_f$	$\hat{\lambda}$	F-stat	p-over
GMM	0.240 (0.081)	0.310 (0.069)	0.005 (0.003)	10.065	0.106
LIML	0.358 (0.125)	0.209 (0.121)	0.002 (0.585)	7.200	

Note: The full set of instruments includes lagged inflation of order 2,3,5 and 11, lagged real broad money of order 1 and 3 lags of broad money, two lags of real gross wages and real exchange rate to euro, and four lags of the output gap.

Source: Authors' estimations.

Coefficients α_b and α_f do not aggregate to 1. As the value of the F-statistic is just above 10, the Hansen J-test was carried out for over-identifying restrictions. The null hypothesis that the over-identified restrictions hold is accepted at an 11% significance level.

Slovakia

The NKPC model for Slovakia was obtained by employing the output gap as a measure of the economic driving variable (Table 5.4), however, it was not significant. On the other side, both parameters, α_b and α_f , are significant with estimates of 0.27 and 0.45 respectively. According to the results obtained, inflation dynamics are impacted to a greater extent by the expectations about future inflation than the backward-looking behavior. The estimated coefficients α_b and α_f do not aggregate to near unity.

The value of the F-statistic is a bit higher than 10 (11.06), while the corresponding p-value of the Hansen J-test is 0.06. It is possible that some of the instruments play an endogenous role; however, utilizing the instruments did not lead to the development of a better quality model.

Table 5.4 Estimation of NKPC for Slovakia Using Output Gap (January, 1997-July, 2010)

	$\hat{\alpha}_b$	$\hat{\alpha}_f$	$\hat{\lambda}$	F-stat	p-over
GMM	0.268 (0.060)	0.448 (0.086)	0.018 (0.018)	11.062	0.060
LIML	0.271 (0.123)	0.468 (0.026)	0.026 (0.038)	5.534	

Note: The full set of instruments includes lagged inflation of order 2 and 11, lagged real gross wage inflation (growth) of order 2, 3 and 4, three lags of real exchange rate to euro and two lags of the output gap.

Source: Authors' estimations.

Romania

For estimating the NKPC in Romania, real gross wage inflation was taken as the inflation forcing variable; furthermore it was found to have a significant and positive influence on inflation dynamics (Table 5.5). The estimate for the backward-looking element is in the range of 0.31-0.40 and that for the forward-looking element is 0.59-0.67. Though the estimates depend to some extent on the method applied, they uniformly suggest that the expectations about future inflation serve a more important role in determining the inflation dynamics.

Table 5.5 Estimation of NKPC for Romania Using Real Gross Wage Inflation (January, 2001-Jun, 2010)

	$\hat{\alpha}_b$	$\hat{\alpha}_f$	$\hat{\lambda}$	F-stat	p-over
GMM	0.314 (0.095)	0.668 (0.105)	0.031 (0.009)	10.118	0.665
LIML	0.399 (0.130)	0.586 (0.138)	0.039 (0.013)	3.375	

Note: The full set of instruments includes lagged inflation of order 2,4,5 and 11, three lags of real gross wage inflation (growth) and four lags of unemployment rate (seasonally adjusted).

Source: Authors' estimations.

The value of the F-statistic for the overall relevance of excluded instruments is marginally greater than 10, but the p-value for the J-statistic is relatively high, 0.665. Thus, the specification reported can be considered adequate.

Serbia

In estimating the NKPC for Serbia, real broad money growth was taken as the inflation driving variable (Table 5.6). It appears as a significant regressor, but only if the GMM was applied. However, both tests that check the appropriateness of instruments clearly accept their validity and thus, the implementation of the GMM. Parameters on past and future inflation are highly significant in both versions: parameter α_b was estimated to be around 0.35-0.36 and parameter α_f around 0.51-0.55; yet, the forward-looking weight is higher than the weight on the backward behavior.

Table 5.6 Estimation of NKPC for Serbia Using Real Broad Money Growth (January, 2003-Jun, 2010)

	$\hat{\alpha}_b$	$\hat{\alpha}_f$	$\hat{\lambda}$	F-stat	p-over
GMM	0.353 (0.061)	0.513 (0.082)	0.046 (0.027)	10.071	0.371
LIML	0.361 (0.156)	0.552 (0.142)	0.053 (0.367)	2.570	

Note: The full set of instruments includes lagged inflation of order 2 and 11, five lags of real gross wage inflation (growth), four lags of real broad money growth and three lags of real exchange rate to euro.

Source: Authors' estimations.

In summary, it is evident from Table 5.7 that inflation persistence was driven both by expectations about future inflation and lagged inflation. However, the contribution of the backward term is relatively less important. The effect of lagged inflation was further assessed by the value of the root g_1 which appears in moderate size across the sampled countries. The influence of lagged inflation is the highest (0.65) for Romania and the lowest for the Czech Republic (0.26) and Slovakia (0.32). For the remaining three countries, it is in the range of 0.42-0.49. Finally, in five out of the six NKPC models (Slovakia is the exception), the impact of the inflation driving variable is found to be significant; thus, the results obtained confirm the validity of the NKPC for the CSEECs.

Table 5.7 Summary from Estimated NKPC Models

Country	$\hat{\alpha}_b$	$\hat{\alpha}_f$	\hat{g}_1	$\hat{\lambda}$	$p = (1 - \hat{\theta})$
Hungary	0.41-0.45	0.48-0.51	0.42	Significant	0.15
Czech Republic	0.24-0.36	0.21-0.31	0.26	Significant	0.01
Poland	0.35-0.45	0.48	0.49	Significant	0.09
Romania	0.31-0.40	0.59-0.67	0.65	Significant	0.08
Serbia	0.36	0.51-0.55	0.45	Significant	0.08
Slovakia	0.27	0.45	0.32	Insignificant	-

Source: Authors' estimations.

Using GMM estimates, the probability of a price change over a given month was calculated for all of the countries (the last column of Table 5.7). The probability for Hungary is the highest (0.15), suggesting that firms need about 7 months to adjust their prices. The probability found for Poland (0.09) indicates that full adjustment of prices requires about 11 months. The frequency of price change (0.08) in Romania and Serbia shows that price adjustment required approximately 12.5 months. The adjustment detected for the Czech Republic is extremely small. The present authors' calculation shows that, on the average, the price adjustment was faster than the 4.5 quarters (about 13.5 months) reported by Hondroyannis, Swamy, and Tavlas (2008) in their panel analysis for the new EU member states.

The present authors also performed the NKPC estimation for the quarterly data for Hungary, the Czech Republic, Poland and Slovakia. Again, the sample covered the period from the first quarter, 1995 to the fourth quarter, 2010 (except for Slovakia where the sample started in the first quarter of 1997). The GMM method was implemented and the results are summarized in Table 5.8.

Table 5.8 Estimation of NKPC for Selected CSEECs Using Quarterly Data (1st quarter 1995/1997- 4th quarter, 2010)

	$\hat{\alpha}_b$	$\hat{\alpha}_f$	$\hat{\lambda}$	Driving var.	F-stat	p-over
Hungary	0.428 (0.072)	0.557 (0.074)	0.010 (0.018)	Insignificant	12.897	0.018
Czech Republic	0.505 (0.079)	0.385 (0.114)	0.119 (0.024)	Significant	6.816	0.807
Poland	0.469 (0.076)	0.532 (0.097)	-0.002 (0.003)	Insignificant	10.214	0.119
Slovakia	0.429 (0.112)	0.502 (0.099)	0.116 (0.028)	Significant	6.261	0.451

Note: The real gross wage inflation (growth) was used as a driving variable for all countries, except Poland, where real broad money was used. Following are the compositions of the full sets of instruments for the different countries: Hungary – lagged inflation of orders 2 and 11, lagged real gross wage inflation (growth) and output gap of orders 1 and 3, and four lags of real broad money growth; the Czech Republic – lagged inflation of orders 2 and 11, two lags of real gross wage inflation (growth), output gap and real broad money growth; Poland – lagged inflation of orders 2, 3, 5, 7 and 11, lagged real gross wage inflation (growth) of orders 1 and 4, and lagged real broad money of orders 1 and 3; Slovakia – lagged inflation of orders 2 and 11, two lags of real gross wage inflation (growth) and output gap. The output gap is based on detrended logarithm of quarterly measured GDP (seasonally adjusted).

Source: Authors' estimations.

In all four economies, weights on backward and forward looking variables are significant. The role of the driving variable varies across countries and it is found to be significant only in the Czech Republic and Slovakia. However, the results for these two economies are not completely satisfactory given the low values of the F-statistics, indicating weak instruments. It is found that the forward-looking element has a slightly higher magnitude than the backward-looking term, except for the Czech Republic, in explaining inflation dynamics. However the dominant position of the forward-looking component over past inflation has not been supported by either the monthly or quarterly results. This absence leads to the additional conclusion that inflation persistence exists at a relatively high level. Similar findings, based on quarterly data, were reported by Franta, Saxa, and Šmidkova (2007) and Menyhert (2008). However, Hondroyiannis, Swamy, and Tavlas (2008) contend that past inflation plays much smaller role than the forward-looking component. The difference in the results can be explained as a result of the difference between the samples studied and the methodologies applied.

6. Conclusions

This study offers two sets of results concerning inflation persistence in the selected CSEECs. First, a detailed analysis of the correlation pattern in inflation dynamics was undertaken within the univariate approach, allowing for random changes of economic regimes. The analysis reveals that inflation persistence is still of a moderate to high magnitude in Hungary, Poland, Romania and Serbia, and of small magnitude in Slovakia and the Czech Republic. It is found that changes in inflation persistence

often correspond to changes in the variability and mean of inflation. Unlike the countries from Central Europe, the two Balkan countries: Romania and Serbia, exhibit more frequent switching of regimes with different levels of inflation persistence. This indicates that these economies have a higher sensitivity to random inflation shocks.

Second, the study proves that the NKPC represents a valid structural approach to describe inflation dynamics in the studied region. In all of the six cases studied, the weights on backward and forward looking behavior are significant, while the impact of the driving variable was insignificant only once. It is found that significant influence of the economic driving variable is captured by real gross wage inflation and real broad money growth. The estimates of this study show that the forward-looking element does not play a dominant role over the backward-looking term in determining inflation dynamics.

Relatively high inflation persistence suggests that agents employ simple adaptive expectations in price setting, which is often the case when monetary policy has low credibility. In such a situation, the appropriate monetary policy should be one that is inertial, such that a reduction in inflation is associated with output loss. On the other side, as the NKPC suggests, expectations about future inflation cannot be ignored as the formation of expectations is not exclusively adaptive. Monetary policy could have been relatively effective had it changed the expectations about future inflation which are now found. It is fair to say that the monetary authorities in these countries should consider all of the instruments available for conducting credible policy, because considering only one or some of them may not be sufficient to reach the announced inflation target level.

The present analysis is based on monthly data which provides accurate insight into inflation dynamics without inducing false autocorrelation in either the univariate or structural set-up. The present authors opted to work with a switching regime time-varying framework in univariate models to discover specific features of the time-series structure in inflation. They did not follow the same framework in estimating the NKPC, because the included driving variable has the potential to capture information about the changes.

Given that most researches on this topic use quarterly or yearly data, quarterly data was considered for this research also in its final phase, for Poland, Hungary, Czech Republic and Slovakia. Within the Markov switching framework, similar conclusions were obtained as those obtained with monthly data; also, the findings concur with those of the recent study by Franta, Saxa and Šmidkova (2010). The quarterly data-based NKPC suggests that both backward and forward components explain significantly the inflation in each of the four economies. The role of backward component was considerably large. This is in accordance with some of the results obtained earlier from quarterly data-based NKPC (Franta, Saxa and Šmidkova 2007; Menyhert 2008).

Comparison between current euro zone countries and the economies studied here is relevant, because it enables one to assess the extent that inflation dynamics in CSEECs might converge with that of the euro zone. For this, Slovakia can be taken as a benchmark for which there is evidence to the effect that the persistence level was the lowest and that the weight on the forward-looking term is almost double that on

the lagged element, as inferred from the monthly data. It needs to be stressed here again that, except for the Czech Republic, all other economies in the region were characterized by a higher level of persistence and that their forward-looking element was not as dominant over the backward-looking one, as the forward-looking element of Slovakia. Not enough evidence was found to show that the inflation dynamics in the countries included in this study had a tendency to converge on the inflation dynamics of the current euro zone members. To strengthen the inferences drawn from this study, it is proposed to carry out similar econometric studies in additional euro zone countries.

References

- Angeloni, Ignazio, Luc Aucremanne, Jordi Gali, Andrew Levin, and Frank Smets.** 2006. "New Evidence on Inflation Persistence and Price Stickiness in the Euro Area: Implications for Macro Modeling." *Journal of the European Economic Association*, 4(2-3): 562-574.
- Babetski, Jan, Fabrizio Coricelli, and Roman Horvath.** 2008. "Assessing Inflation Persistence: Micro Evidence on an Inflation Targeting Economy." CERGIE-EI Working Paper 353.
- Bardsen, Gunnar, Eilev S. Jansen, and Ragnar Nymoen.** 2004. "Econometric Evaluation of the New Keynesian Phillips Curve." *Oxford Bulletin of Economics and Statistics*, 66(S1): 671-686.
- Coricelli, Fabrizio, and Roman Horvath.** 2006. "Price Setting Behaviour: Micro Evidence on Slovakia." Centre for Economic Policy Research Discussion Paper 5445.
- Darvas, Zsolt, and Balazs Varga.** 2007. "Inflation Persistence in the Euro area, US, and New Members of the EU: Evidence from Time-Varying Coefficient Models." Paper presented at the Money Macro and Finance Research Group Conference, Birmingham.
- Doornik, Jurgen.** 2009. *Oxmetrics6*. London: Timberlake Consultants Press.
- Franta, Michal, Branislav Saxa, and Katerina Šmidkova.** 2010. "The Role of Inflation Persistence in the New EU Member States." *Czech Journal of Economics and Finance (Finance a uver)*, 60(6): 480-500.
- Franta, Michal, Branislav Saxa, and Katerina Šmidkova.** 2009. "The Role of Inflation Persistence in the New EU Member States." In *What Drives Inflation in the New EU Member States*, 11-24. Brussels: European Commission.
- Gali, Jordi, and Mark Gertler.** 1999. "Inflation Dynamics: A Structural Econometric Approach." *Journal of Monetary Economics*, 44(2): 195-222.
- Gali, Jordi, Mark Gertler, and David J. Lopez-Salido.** 2005. "Robustness of the Estimates of the Hybrid New Keynesian Phillips Curve." *Journal of Monetary Economics*, 52(6): 1107-1118.
- Hamilton, James D.** 1989. "A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle." *Econometrica*, 53(5): 357-384.
- Hamilton, James D.** 1990. "Analysis of Time Series Subject to Changes in Regime." *Journal of Econometrics*, 45(1-2): 39-70.
- Hondroyannis, George, P.A.V.B. Swamy, and George S.Tavlas.** 2008. "Inflation Dynamics in the Euro Area and in New EU Members: Implications for Monetary Policy." *Economic Modelling*, 25(6): 1116-1127.
- Konieczny, Jerzy D., and Andrzej Skrzypacz.** 2005. "Inflation and Price Setting in a Natural Experiment." *Journal of Monetary Economics*, 52(3): 621-632.
- Linde, Jesper.** 2005. "Estimating New-Keynesian Phillips Curves: A Full Information Maximum Likelihood Approach." *Journal of Monetary Economics*, 52(6): 1135-1149.
- Menyhert, Balint.** 2008. "Estimating the Hungarian New-Keynesian Phillips Curve." *Acta Oeconomica*, 58(3): 295-318.
- Paya, Ivan, Agustín Duarte, and Ken Holden.** 2007. "On the Relationship between Inflation Persistence and Temporal Aggregation." *Journal of Money, Credit and Banking*, 39(6): 1522-1531.

- Siklos, Pierre L.** 2008. "Inflation Targeting Around the World." *Emerging Markets Finance and Trade*, 44(6): 17-37.
- Staiger, Douglas, and James H. Stock.** 1997. "Instrumental Variables Regression with Weak Instruments." *Econometrica*, 65(3): 557-586.
- Zhang, Chengsi, and Joel Clovis.** 2010. "The New Keynesian Phillips Curve of Rational Expectations: A Serial Correlation Extension." *Journal of Applied Economics*, 13(1): 159-179.

Appendix

Table A1. Data Definitions and Sources for Monthly Data

Variable	Definition	Sources
Inflation rate , based on CPI base monthly index (Pt).	transformation: $\pi_t = 100 (\log Pt - \log Pt-1)$, seasonal adjustment is made	OECD data: http://stats.oecd.org/Index.aspx data for Serbia: Statistical Office of the Republic of Serbia http://webrzs.stat.gov.rs/axd/en/index.php data for Romania: WIW Monthly Database on Eastern Europe: http://mdb.wiiw.ac.at/
Output gap , based on detrended log GDP.	Deviation from the long-term trend component of GDP monthly data. Smooths the GDP series are obtained by Hodrick-Prescott method with a smoothing parameter of 14400. Monthly data of GDP are obtained by the disaggregation of quarterly measured GDP (seasonally adjusted).	OECD dataset on Quarterly National Accounts: http://stats.oecd.org/Index.aspx data for Romania: National Institute of Statistics http://www.insse.ro/cms/rw/pages/index.n.do data for Serbia: Statistical Office of the Republic of Serbia http://webrzs.stat.gov.rs/axd/en/index.php
Unemployment rate	Unemployment rate, registered (in %, end of a period), seasonally adjusted monthly data.	WIW Monthly Database on Eastern Europe: http://mdb.wiiw.ac.at/
Real broad money , (BMRT)	transformation: broad money (base index) /CPI index, monthly data.	OECD data: http://stats.oecd.org/Index.aspx data for Serbia: National Bank of Serbi-bi- http://www.nbs.rs/export/internet/englis h/80/index.html
Growth of real broad money	transformation: $\Delta BMRT = BMRT - BMRT-1$	
Real gross wages , (WRT)	transformation: gross wages (total economy) /CPI index, monthly data.	OECD data: http://stats.oecd.org/Index.aspx data for Romania: WIW Monthly Database on Eastern Europe: http://mdb.wiiw.ac.at/ data for Serbia: Statistical Office of the Republic of Serbia http://webrzs.stat.gov.rs/axd/en/index.php
Real gross wage inflation	transformation: $\Delta WRT = WRT - WRT-1$	
Real exchange rate to euro	monthly averages, adjusted for domestic and EU inflation.	WIW Monthly Database on Eastern Europe: http://mdb.wiiw.ac.at/

Table A2. Data Definitions and Sources for Quarterly Data

Variable	Definition	Sources
Inflation rate , based on CPI base quarterly index (Pt).	transformation: $\pi_t = 100 (\log Pt - \log Pt-1)$, seasonal adjustment is made	OECD data: http://stats.oecd.org/Index.aspx
Output gap , based on detrended log GDP.	Deviation from the long-term trend component of GDP quarterly data (seasonally adjusted). Smooths the GDP series are obtained by Hodrick-Prescott method with a smoothing parameter of 1600.	OECD dataset on Quarterly National Accounts: http://stats.oecd.org/Index.aspx
Real broad money , (BMRT)	transformation: broad money (base index) /CPI index, quarterly data.	OECD data: http://stats.oecd.org/Index.aspx
Growth of real broad money	transformation: $\Delta BMRT = BMRT - BMRT-1$	
Real gross wages , (WRT)	transformation: gross wages (total economy) /CPI index, quarterly data.	OECD data: http://stats.oecd.org/Index.aspx
Real gross wage inflation	transformation: $\Delta WRT = WRT - WRT-1$	