Inflation persistence in Central and Southeastern Europe: Evidence from time series approach

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Abstract: The purpose of the paper is to measure inflation persistence in the following countries from the Central and Southeastern Europe: Slovakia, Czech Republic, Poland, Hungary, Romania and Serbia. Sample covers monthly data from January, 1995 to May, 2010 for Poland, Hungary and Slovakia, and a previous year, 1994, in the case of Czech Republic. Inflation in Romania is considered for period January, 2002 - June 2010. The shortest sample is used for Serbia, given a late start of transition process (January, 2003 - September, 2010. Our results contribute to the existing empirical results on this topic in following way: sample period is extended by including recent years of relatively higher inflation rates, and Romania and Serbia, not being previously considered, are included. Paper offers two set of results. First, correlation pattern in inflation dynamics is estimated within the Markov switching model approach that allows for random changes in economic regimes. It is revealed that inflation persistence is still of moderate to high magnitude in Hungary, Poland, Romania and Serbia. In Slovakia and Czech Republic inflation persistence is estimated to be of smaller order. It is documented that changes in inflation persistence often correspond to changes in variability and mean of inflation. Second, we provide evidence that New Keynesian Phillips Curve represents a valid structural approach to describe inflation dynamics in this region. In all six cases weights on backward and forward looking behavior are significant, while the impact of driving variable was insignificant only once. We found that significant influence of economic driving variable is captured by real gross wage inflation and real broad money growth. Our estimates show that in general forward-looking element has slightly higher magnitude than backward-looking term in determining inflation persistence.

1. Introduction

Inflation persistence is one of the key issues in adjusting monetary policy since the stabilization of inflation at low level represents main goal of most central banks. Given a specific shock, inflation persistence can be described as a tendency of the inflation rate to converge slowly toward its long-run value (Paya et al., 2007). Therefore, the magnitude of inflation persistence provides valuable information for central banks to conduct such a policy that would meet announced target level. Understanding the path and determinants of inflation persistence is for the same reason relevant for those Central and Southeastern European countries (CSEECs) that have adopted inflation targeting. More specific, difference between inflation persistence in current euro area and in CSEECs that are EU members could suggest asymmetric impact of common shock between two groups of countries and indicate extent to which their convergence towards the current euro area might decline upon the euro is adopted. In addition, the successfulness of recently introduced inflation targeting in some CSEECs can be assessed by taking into account experience of other economies from the same region.

It is widely noted that the adoption of inflation targeting should be associated with the sharp decrease in persistence. The identification of this fall in persistence has been extensively investigated for major industrial countries (the U.S. and euro area). In recent years performances of inflation targeting in emerging market economies have also been analyzed. Survey based on descriptive and econometric evidence (Siklos, 2008) confirms prevailing empirical findings that inflation targeting has diminished inflation persistence in industrial countries, but only in a handful of emerging market economies.

Inflation persistence can be measured on micro and macro data. Micro data analysis considers statistical properties of product-level consumer prices indexes underlying the consumer basket. Macroeconometric investigation is based on time series univariate and structural modelling. Within univariate approach different methods can be applied to estimate the extent to which inflation reacts to unexpected random shocks. Structural time series modeling is usually based on the estimation of the hybrid New Keynesian Phillips Curve (NKPC).

In this paper we consider the following CSEECs: Czech Republic, Poland, Hungary, Romania and Serbia. Czech Republic, Poland and Hungary have relatively long experience in inflation targeting, while Romania and Serbia introduced inflation targeting recently. The objective of the paper is to measure the magnitude of inflation persistence in these five economies. In addition, inflation persistence will be measured in Slovakia, as the only country from this region being currently in the euro zone. Estimation for Slovakian inflation will be used as a benchmark case. Within the univariate time series approach time varying estimation is carried out to take account of switching in regimes over period considered. Structural modelling is based on formulating NKPC to additionally control for nominal rigidities. Estimation of NKPC function provides further insight into inflation process to assess whether inflation is mostly forward-looking phenomenon or instead backward-lookingness is predominant characteristic.

Sample covers monthly data from January, 1995 to May, 2010 for Poland, Hungary and Slovakia, and a previous year, 1994, in the case of Czech Republic. Inflation in Romania is considered for period January, 2002 - June 2010. The shortest sample is used for Serbia, given a late start of transition process (January, 2003 - September, 2010). Our empirical results contribute to the existing empirical results for CSEECs in the following way: sample period is extended by including recent years of relatively higher inflation rates, and Romania and Serbia, not being previously considered, are included.

Paper is structured as follows. Basic macroeconomic indicators for selected countries are given in Section 2. Section 3 shortly overviews empirical literature on inflation persistence in CSEESs. Some methodological issues are discussed in Section 4. Findings derived from univariate methods are presented in Section 5, while Section 6 contains results based on estimating NKPC. Section 7 concludes. Data used in the paper are described in Appendix 1.

2. Basic macroeconomic indicators for selected countries

Our study covers six selected CSEE countries. As noted above, all of them adopted inflation targeting as monetary policy framework, but only Slovakia has recently (in 2009) entered the euro zone. With the exception of Serbia, all selected countries are EU members (Romania joined the EU in 2007, while the others in 2004). As reported in Table 2.1, these countries diverse in terms of area and population. Poland is the largest one, with the population of 38.2 million, while Slovakia is on the other end with population of 5.4 million. In terms of GDP per capita (adjusted for purchasing power parity) and average monthly wages, Romania and Serbia are sufficiently below the others (as illustration, Serbia is less than half from the Poland GDP per capita level, and the same stands in terms of average wages for both countries). Romania and Serbia have experienced much higher annual inflation rates since the mid 1990s. It is also important to notice that Slovakia, Czech Republic and Hungary are highly integrated countries in terms of international trade (trade openness ratio exceeds 100 percent). In the period of interest, all of the countries are characterized by the exchange rate flexibility.

Table 2.1 Summary indicators for selected CSEE countries

	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
Exchanege-rate system			Managed float	Fixed with band to euro	Free float	Managed float	Managed float	in EMR II intr. of euro in 2009

Source: WIIW (2010), EBRD (2010).

3. Previous literature on inflation persistence in CSEECs

The issue of inflation persistence in CSEECs has been considered in a number of papers with results summarized in Table 3.1.

Table 3.1 Overview of selected studies

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 higer 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.
Hondroyiannis 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.

Early researches on this topic are based on micro data analyses for Czech Republic (Babetski et al., 2008), Poland (Konieczny and Skrzypacz, 2005) and Slovakia (Coricelli and Horvath, 2006). Monthly data are used for the sample that does not exceed year 2005. Persistence among different price indices is measured with the only finding for the overall CPI inflation in Czech Republic suggesting that inflation seems to be less persistent after adoption of inflation targeting. Magnitude of inflation persistence in Hungary is estimated by univariate and structural time series methods for the sample that ends in 2005 and mid 2006 (Darvas and Varga, 2007 and Menyhert, 2008). It is found that for that period inflation persistence in Hungary was higher that in US and euro area and that inflation dynamics was determined equally by past inflation and forward-looking expectations. Contrary to individual case studies, Franta et al. (2007) and Hondroyiannis et al. (2008) cover several countries from CSEE along with euro area based on sample that ends in the first quarter of 2006 and the third quarter of 2005 respectively. While estimation of different univariate time series methods and the NKPC model is performed in Franta et al. (2007), only the NKPC model is considered in Hondroyiannis et al. (2008). The main results indicate that inflation persistence was comparable to that in the euro area with dominant role of backward behaviour in inflation dynamics. In addition, when time-varying specification of NKPC is estimated the role of lagged inflation significantly diminished.

Results overviewed are obtained from the analysis of quarterly data of CPI and/or core inflation. However, annualized quarterly inflation rate is actually used in empirical studies to measure the persistence. To clarify, let us assume that P_t represents the log price index at quarter t. Then annualized inflation rate is derived in one of the following ways:

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

Inflation rate I can be described as an aggregate one derived from quarterly data in the following way:

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

Such inflation rate might make measurement of persistence invalid when univariate time series methods are used (cf. Paya et al., 2007). More precisely, inflation rate calculated as temporal aggregates from the actual highest frequency available data (quarterly or monthly) might cause upward bias in estimating inflation persistence. One could also argue that this transformation might induce additional autocorrelation when structural econometric model is estimated.

Inflation rate II provides adequate measure of annualized inflation rate only if the ratios 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 economies that were characterized by significant regime changes such as CEES region it is highly unexpected for quarterly inflation rates to have equal ratios of type above, which questions the appropriateness of calculating inflation rate in this way. Such a computation, however, does not cause bias in measuring persistence from univariate methods, but can induce problems within structural approach that evaluates contribution of the set of explanatory variables on inflation dynamics.

4. Short overview of methodology applied

Our results are based on seasonally adjusted monthly CPI index, denoted as P_t, based on which monthly inflation rate is calculated as: $\pi_t = 100(\log P_t - \log P_{t-1})$. Baseline method in time series analysis to measure the persistence is 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

$$\pi_t = \alpha_0 + \rho\pi_{t-1} + \sum_{i=1}^{(p-1)} \delta_i \Delta \pi_{t-i} + e_t \qquad (4.1)$$

such that parameter $\rho = \sum_{i=1}^{p} \phi_i$ contains information about the sum of autoregressive parameters and thus provides measure of inflation persistence. This specification can be modified in a number of different ways to take account of possible regime changes and nonlinearity in inflation rate. In fact, several empirical papers on inflation persistence in transition economies (for example, Franta et al., 2007 and Darvas and Varga, 2007) argued that linear specification is not rich enough to capture true dynamics in inflation rate. To allow for changes in some parameters we employ the Markov-switching model assuming that mean, variability and persistence differ among two regimes. The relevant specification is of the following form (Hamilton, 1989, 1990):

$$\pi_t = \left(\alpha_0 + \alpha_1 S_t\right) + \left(\rho + \rho_1 S_t\right)\pi_{t-1} + \delta_1 \Delta\pi_{t-1} + \ldots + \delta_{p-1}\Delta\pi_{t-p+1} + \left(h_0 + h_1 S_t\right)e_t \qquad (4.2)$$

S_t is the unobserved random variable that follows a Markov chain defined by transition probabilities between two states. The full matrix 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: $p_{i/j}$ =P(Regime *i* at *t+1*/Regime *j* at *t*)

Switches of economy from state 0 to state 1 is governed by introduced random variable S_t. Under this specification we have two different regimes: regime 0 (i.e. S_t =0) and regime 1 (i.e. S_t =1). The parameters α_1, ρ_1, h_1 capture the changes in the mean of inflation, persistence of a shock to inflation and the variance during regime 1 relative to regime 0. Positive value of parameter implies a shift from low to high inflation persistence and vice versa.

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

$$\pi_t = \alpha_b \pi_{t-1} + \alpha_f E(\pi_{t+1}) + \lambda mc_t + \text{error term} \qquad (4.3)$$

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} is used for expected future inflation. As shown in Gali and Gertler (1999), this specification can be rewritten conditional on the expected path of real marginal cost:

$$\pi_t = g_1\pi_{t-1} + \frac{\lambda}{g_2\alpha_f}\sum_{k=0}^{\infty}\left(\frac{1}{g_2}\right)^k E\{mc_{t+k}\} + \text{error term} \qquad (4.4)$$

g_1 and g_2 denote respectively the stable and unstable roots of the corresponding second order difference equation associated with (4.3) that are equal to:

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

Several issues emerge in estimating NKPC (4.3) and we will briefly highlight two most important ones. First, there is no agreement over the choice of estimation method since the specification assumes endogeneity of a variable and an AR(1) structure of an error term. Gali and Gertler (1999) advocated general method of moments (GMM) showing robustness of their empirical results (Gali et al., 2005) given ongoing debate on the appropriateness of this method. Apart from GMM, parameters of NKPC are estimated in different empirical studies by full or limited information maximum likelihood (Linde, 2005 for example). Alternatively, the baseline specification can be slightly modified to take care of autocorrelation such that TSLS can be employed (for example, Zhang and Clovis, 2010 and Bardsen et al., 2004). Second, there is no clear answer to the question what variables to use as proxies for real marginal costs. They are often measured by real unit labor costs that capture impact of wages and productivity on inflation. Real output gap is also used in a number of empirical studies, although Gali and Gertler (1999) argued that it was not a valid real driving force for inflation. Some recent studies suggest that for small and open economies terms of trade should play important role in driving inflation.

The set-up of NKPC appears as demanding task in the CEECs. Many relevant time series are either incomplete or unavailable. Most existing empirical results are based on the samples that include turbulent episodes at the beginning of transition process with price liberalization that dominated inflation dynamics over the whole period considered. In addition, instead of using real labor unit costs, wage inflation appears as a natural candidate for inflation driving force 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.

5. Empirical findings from univariate approach

Hungary

Two-state Markov switching model fits well dynamics of monthly inflation rate in Hungary for period January, 1995-May, 2010. As reported in Table 5.1, two different inflation persistence regimes have been detected. Regime 0 has lower inflation persistence characterized by estimated magnitude $\hat{\rho} = 0.38$. This is also a regime of lower mean inflation rate, 0.204/(1-0.375)=0.33, and lower variability, $\hat{h}_0 = 0.07$. Regime 1 is found to have higher inflation persistence: estimate is 0.82. During regime 1 inflation rate exhibited higher mean value, 0.101/(1-0.823)= 0.57, and higher variability, 0.30. Statistically, there is a significant difference between two parameters of inflation persistence.

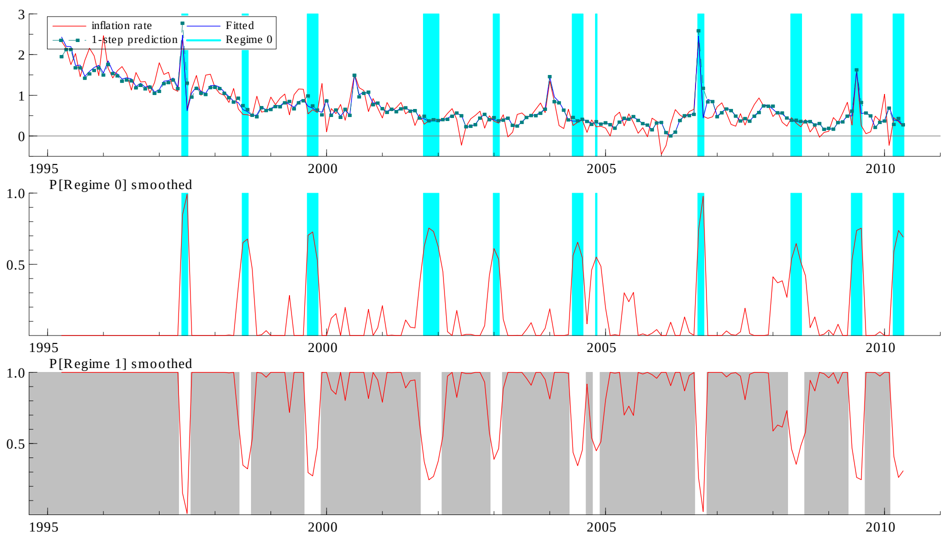
The probability q of remaining in the regime of lower persistence, while being in that regime is 0.55. The probability of remaining in regime of higher persistence, as already being in that regime is 0.92, implying that the probability f of switching from the regime of higher to regime of lower persistence is small and is equal to 0.08. Visual representation of estimates is given in Graph 5.1. We notice that in fact most of the time, 84.62%, economy is in the regime 1 of relatively high inflation persistence. This also holds for the last years of relatively high inflation rate.

Simple linear autoregressive model suggested that inflation persistence is of moderate size, 0.78. Since this specification is improved significantly by estimates obtained, our conclusion is that inflation persistence is relatively high in Hungary as prevailing estimate is 0.823.

Table 5.1 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	19.3
h_0	0.07	2.56
(h_0+h_1)	0.30	16.6
Q	0.55	2.05
F	0.08	2.52
δ_1	-0.385	-7.41
δ_2	-0.196	4.71
Linearity test : $\chi^2_5 = 16.53(0.005)$; Box – Pierce Q(36) $\chi^2_{36} = 34.22(0.55)$, ARCH1 F(1,165) = 2.07(0.15); Normality $\chi^2_2 = 5.80(0.06)$ Five impulse dummy variables are included that have non-zero value 1 for the following months: 1997:6, 2000:7, 2004:1, 2006:9 and 2009:7.		

Graph 5.1 Two regimes of inflation persistence in Hungary



Czech Republic

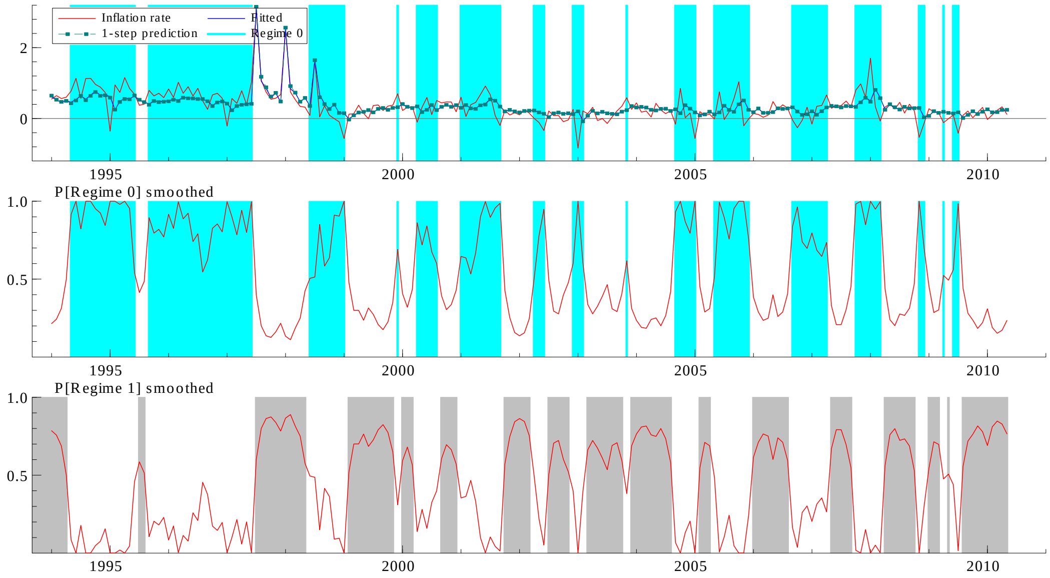
Two-state Markov switching model is also used to estimate monthly inflation rate in Czech Republic for period January, 1994-May, 2010 (Table 5.2). Regime 0 has higher inflation persistence estimated to be 0.57. During this regime inflation rate exhibits higher mean, 0.135/(1-0.570)=0.31, and variability, 0.39. Regime 1 is described as regime of lower inflation persistence with estimate 0.46. During regime 1 inflation rate has lower mean value, 0.138/(1-0.460)=0.26, and variability, 0.15. Although two estimates of persistence do not differ significantly, we proceed with two regimes analysis given significant difference in inflation mean and variability.

The probability *q* of remaining in the regime of higher persistence, while being in that regime is 0.80. The probability of staying in regime of lower persistence, as already being in that regime is 0.74, so that probability *f* of switching from the regime of lower to regime of higher persistence is 0.26. Visual inspection of regimes from Graph 5.2 indicates that two regimes are equally split with 49.75% months in regime 0 and 50.25% months in regime 1. When we take into consideration last two years of data (25 month) it appears that this economy was most of the time (20 months) in the regime of lower inflation persistence, mean and variability.

Table 5.2 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	0.39	9.30
(h_0+h_1)	0.15	4.26
Q	0.80	6.28
F	0.26	2.09
δ_1	-0.259	-4.33
δ_2	-0.123	-2.48
δ_3	-0.110	-2.54
Linearity test: $\chi^2_5 = 17.21(0.003)$; Box – Pierce Q(36) $\chi^2_{36} = 47.54(0.10)$, ARCH1 $F(1,181) = 0.212(0.64)$; Normality $\chi^2_2 = 1.15(0.56)$ Three impulse dummy variables are included that have non-zero value 1 for the following months: 1997:7, 1998:1 and 1998:7.		

Graph 5.2 Two regimes of inflation persistence in Czech Republic



Poland

Based on sample January, 1995 - May, 2010 monthly inflation rate is modeled within two-state Markov switching specification. Results are summarized in Table 5.3. Regime 0 is detected to have higher inflation persistence with estimate 0.89. Regime 1 is found to be of lower inflation persistence (estimate is 0.61). Parameters differ significantly (χ_1^2 =3.70(0.04)).

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

Representation of results is provided by Graph 5.3. We notice that almost 80% of the time economy was in regime 0 of higher persistence. The rest of 20% is described by regime 1 that is detected for blocks of one up to four months, mostly characterized by non-standard values. Due to these transitory shocks inflation uncertainty is estimated to be higher in this regime of lower persistence.

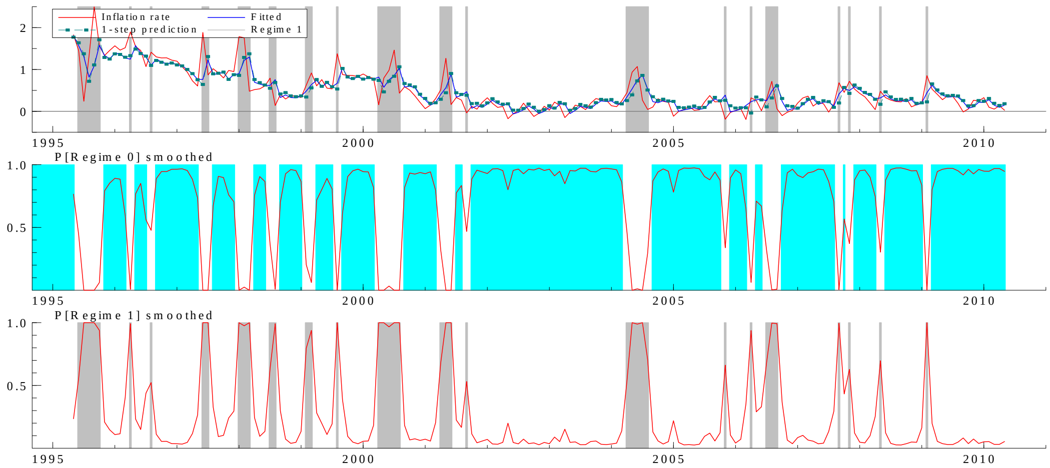
Relatively high degree of persistence is associated with last two years of estimation which, as in case of Hungary, emphasizes the potential problem with ongoing economic instability on the level of inflation persistence.

Table 5.3 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	28.9
$(\rho+\rho_1)$	0.611	4.41
h_0	0.13	9.95
(h_0+h_1)	0.51	8.68
Q	0.85	17.6
F	0.40	4.12

δ_1	-0.225	-3.60
δ_2	-0.132	-2.79
δ_3	-0.092	-2.19
Linearity test : $\chi^2_5 = 83.91\ (0.00)$; Box – Pierce $Q(36)\ \chi^2_{36} = 40.57\ (0.28)$, ARCH 1 $F(1,168) = 3.021\ (0.09)$; Normality $\chi^2_2 = 1.94\ (0.38)$		

Graph 5.3 Two regimes of inflation persistence in Poland



Romania

Two-state Markov switching model performs well for Romanian monthly inflation using period January, 2002-June, 2010 (Table 5.4). Regime 0 has higher inflation persistence estimated to be 0.73. Regime 1 is described to have lower inflation persistence (estimate is 0.52). These measures of persistence differ significantly. Average duration of regime 0 is 1.79 months and it takes 43.43% of the sample. The rest of 56.57% belongs to regime 0 that lasts on average 2.33 months (Graph 5.4). Economy moves from one to another regime frequently. In this case higher inflation persistence episodes are associated with lower mean and less variable inflation rate.

The probability of staying in regime of lower persistence, as already being in that regime is 1-f=0.66, while the probability of switching into another regime is 0.34.

Table 5.4 Estimated inflation model for Romania (January, 2002-June, 2010)

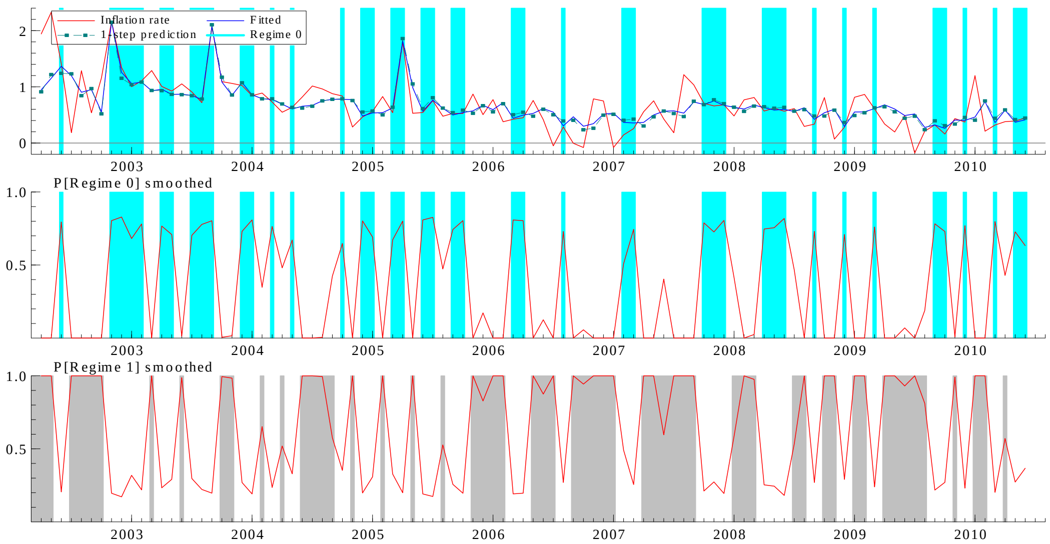
Parameter	**Estimate**	**t-ratio**
α_0	0.092	2.49
$(\alpha_0 + \alpha_1)$	0.308	3.22
ρ	0.733	18.9
$(\rho + \rho_1)$	0.522	4.46
h_0	0.06	4.44
$(h_0 + h_1)$	0.40	10.2
Q	0.41	2.80
F	0.34	3.69
δ_1	-0.253	-5.66
δ_2	-0.204	-7.43

Linearity test : $\chi^2_5 = 31.93(0.000)$; Box – Pierce Q(36) $\chi^2_{36} = 26$

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

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

Graph 5.4 Two regimes of inflation persistence in Romania



Serbia

Inflation dynamics in Serbia is examined within Markov switching model based on the sample that covers shorter time interval than used for the rest of the countries. Serbia entered transition process at the end of 2000. To avoid shocks due to price liberalization and adjustment of several price indices undertaken at the beginning of this process, our sample starts in January 2003. It ends in September 2010.

Models with three regimes fits Serbian inflation rate satisfactory well. Estimated model is presented in Table 5.5 while corresponding transition probabilities are reported separately in Table 5.5a. Regime 0 has high persistence estimated to be 0.81. Average duration of regime 0 is 1 month taking 22.47% of the sample. Regime 1 is also associated with relatively high inflation persistence (estimate is 0.72). Similarly to regime 0, this one also lasts 22.47% of the sample with average duration of 1.25 months. Smallest persistence is found in regime 2, 0.48, that lasted on average 1.58 months. Among three regimes detected highest inflation variability is estimated for regime 2. Graph 5.5 provides visual inspection of regimes.

Adequacy of estimated model is confirmed by several specification tests. While simple AR model suggests that inflation persistence is just the value found for regime 2, 0.48, within this approach we are able to make distinction among episodes of different inflation behaviour. Similarly to results found for Romania, Serbian monthly inflation rate exhibits frequent changes of regimes that occur almost at monthly level. This could suggest extremely high level of sensitivity to unexpected random shocks. This is confirmed by inflation uncertainty being higher when the persistence is lower. In addition, given regime 2 of lower persistence there is almost equal probability (about 0.40) that economy will stay in that regime and switch to regime 0 of higher persistence .

Table 5.5 Estimated inflation model for Serbia (January, 2003-September, 2010)

Parameter	Estimate	t-ratio
α_0	-0.448	-1.67
$(\alpha_0+\alpha_1)$	0.081	7.96
$(\alpha_0+\alpha_1+\alpha_2)$	0.533	5.27
ρ	0.815	5.46
$(\rho+\rho_1)$	0.719	63.5
$(\rho+\rho_1+\rho_2)$	0.479	6.15
h_0	0.221	1.93
(h_0+h_1)	0.014	5.10
$(h_0+h_1+h_2)$	0.350	9.20
δ_1	-0.179	-14.1
δ_2	-0.191	-20.6
δ_3	-0.032	-4.48

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

ARCH1 F(1,165) = 0.02(0.90); Normality $\chi^2_2=4.12(0.13)$

Five impulse dummy variables are 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 is also introduced.

Graph 5.5. Three regimes of inflation persistence in Serbia

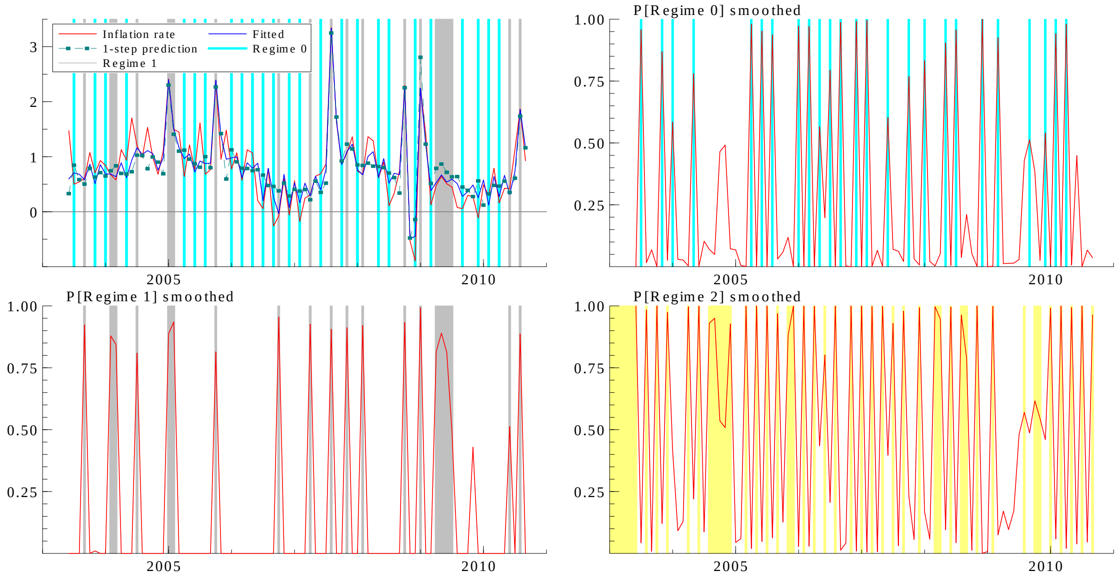


Table 5.5a Estimated transition probabilities from table 5.5

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

Slovakia

At the end of our univariate modelling Slovakian monthly inflation is considered for period January, 1995-May 2010. Estimated model (Table 5.6) implies one regime of inflation persistence, but two regimes of inflation mean and variability. Estimated inflation persistence is 0.28. Regime of lower inflation mean covers 96% of the sample and has average duration of 21.75 months. The rest of 4% is described by higher mean inflation rate that lasted on average 1 month. As seen from Graph 5.6, Slovakian economy has been permanently in the regime of lower mean inflation since November 2005.

Table 5.6 Estimated inflation model for Slovakia (January, 1995-May, 2010)

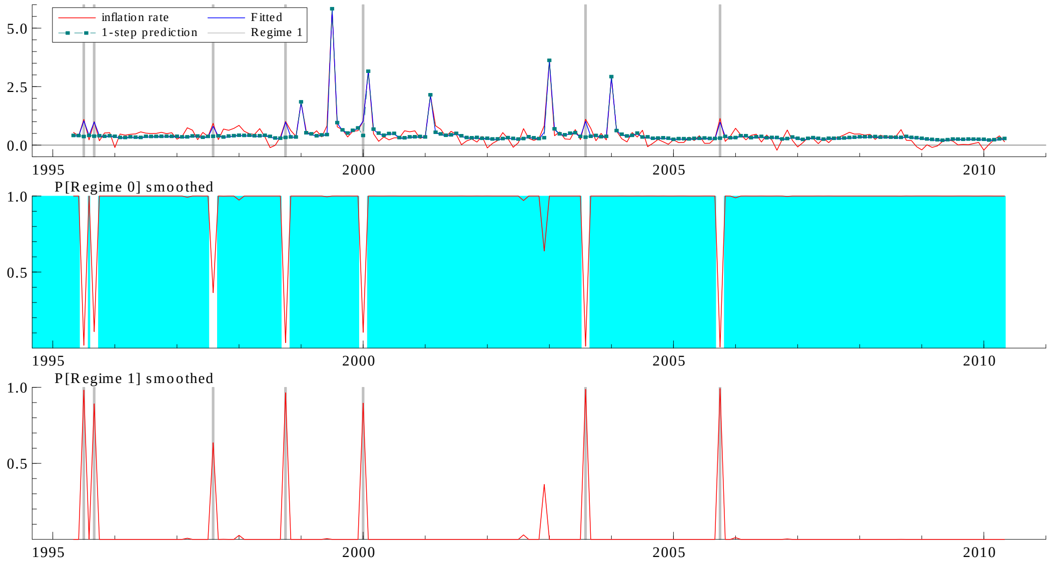
Parameter	**Estimate**	**t-ratio**
α_0	0.206	7.11
$(\alpha_0 + \alpha_1)$	0.914	14.9
ρ	0.283	5.81
h_0	0.201	16.5
h_1	0.100	2.40
δ_1	-0.174	-3.87
δ_2	-0.130	-3.30
δ_3	-0.109	-3.21
δ_4	-0.064	-2.41

Linearity test: $\chi^2_3 = 16.70\ (0.0)$; Box − Pierce $Q(36)\ \chi^2_{35} = 42.590$

ARCH 1 $F(1,163) = 0.37\ (0.54)$; Normality $\chi^2_2 = 0.85\ (0.65)$

Four impulse dummy variables are included to have non-zero value 1 for the following months: 1999:1, 1999:7, 2000:2 and 2003:1. Also. two composite dummies are introduced such that the first one takes non-zero value 1 for 2001:1 and 2002:1, while the second one has non-zero value 1 for 2004:1 and 2003:1.

Graph 5.6 Two regimes of inflation mean and variance in Slovakia



Findings reported are now summarized in Table 5.7. Inflation in Slovakia exhibits by far the lowest level of persistence. Relatively modest magnitude of persistence was determined for Czech Republic. Only in these two economies there was no significant difference between inflation persistence across regimes.

Inflation persistence remains at relatively high level in Hungary and Poland with dominant regimes of higher inflation persistence. The same was concluded for Romania and Serbia. Additional finding for later two economies is that they switch from one to another regime frequently such that relatively lower inflation persistence is characterized by higher inflation uncertainty. This could also be a sign of high sensitivity of inflation rate to unexpected random shocks. Reaction may be described by lower persistence in one of detected regimes, but at the cost of more variability in inflation rate.

Table 5.7 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	

6. Empirical findings from structural approach

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

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

Upon trying different combination of instruments and proxies for the real sector forcing variables, we present the most satisfactory regression estimates for each country. Results are reported below in Tables 6.1-6.6. Overall conclusion from econometric results is that NKPC captures well inflation dynamics in CSEECs.

We performed several diagnostic tests to evaluate NKPC specification. To check for the potential weaknesses of the instruments, we present the F-statistic from the first stage regression of inflation (π_t) against the instruments set. The rule of thumb advocated by Staiger and Stock (1997) is that a value bigger than 10 indicates "no weak instruments" in the case of one endogenous regressor. Another evidence of GMM validity is the Hansen J-test of overidentifying restrictions, which examines weather the orthogonality conditions for instruments are met. The p-value of the Hansen J-statistic is denoted as p-over in tables below. It should be noted that J-test has deficient small-sample properties and could be deceptive when the estimated residuals manifest serial correlation (cf. Zhang and Clovis, 2010, Bardsen et al., 2004).

Previous empirical studies have not found empirical support for NKPC using Polish and Slovakian data. Moreover, this is the first analysis of the inflation dynamics from the New Keynesian perspective in case of Romania and Serbia.

Hungary

NKPC model for Hungary was estimated with real gross wage inflation as real marginal cost variable (Table 6.1). This variable has positive and significant impact on inflation dynamics. Weights on past and future inflation are highly significant. Estimates reached are robust to the change of estimation method. Parameter α_b is estimated in 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 then backward-lookingness. Both parameters sum to 1, as in case of Poland reported next.

Our estimates are not completely comparable to that of Menyhert (2008) where annualized quarterly core inflation rate is applied for the sample that ends in 2006. It was found that portions of backward and forward looking behaviour are about the same and close to 0.5. Taking into account this finding, we may argue that during recent period backward-looking element of inflation dynamics reduced, while forward-looking element remains at about the same level.

Table 6.1 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	0.479	0.100	17.667	0.199
	(0.065)	(0.081)	(0.039)		
LIML	0.413	0.508	0.074	5.431	
	(0.082)	(0.092)	(0.100)		

Note 1: The full instruments set 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 are obtained by the disaggregation of quarterly measured GDP (seasonally adjusted). For a definition of these instruments, see Appendix 1.
Note to Tables 6.1-6.6: GMM results are obtained by the TSLS estimator; HAC-robust standard errors are reported in parenthesis; p-over denotes p-value of Hansen *J*-test; the F-statistics is a measure of the instruments strength (the joint significance test of the instruments in the first stage of regression).

Poland

Real broad money growth is employed in NKPC specification in Poland (Table 6.2). It was estimated to have positive and significant impact on inflation dynamics. Parameters on past and future inflations are also highly significant. Parameter α_b is estimated to be 0.41 by GMM and 0.35 by LIML, while estimate on α_f is around 0.48 in both equations. Forward-looking behaviour predominates in comparison with the backward-looking activity.

Table 6.2 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	0.484	0.055	22.235	0.524
	(0.071)	(0.071)	(0.024)		
LIML	0.354	0.487	0.118	11.032	
	(0.094)	(0.086)	(0.069)		

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

Czech Republic

Inflation driving variable in Czech Republic is real broad money. All variables included in NKPC were found to be significant with positive impact (the only exception is real broad money within LIML estimation, Table 6.3). Estimate on past inflation is in interval 0.24-0.36. Forward-looking weight is estimated in range from 0.21 to 0.31 depending on the method employed. While GMM suggests slightly higher impact of forward-looking element, LIML indicates just the opposite. Our estimate on forward-looking behavior is close to one found in Franta et al. (2007), where, however, higher impact of backward-inflation is estimated (0.42-0.47). It could be the case that during recent years the reduction on backward-looking weight occurred.

Table 6.3 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	0.310	0.005	10.065	0.106
	(0.081)	(0.069)	(0.003)		
LIML	0.358	0.209	0.002	7.200	
	(0.125)	(0.121)	(0.585)		

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

Coefficients α_b and α_f do not sum to 1. As the value of F-statistics is just above 10, we carry out with the Hansen J-test for overidentifying restrictions. The null hypothesis that overidentified restrictions hold is accepted at 11% significance level.

Slovakia

NKPC model for Slovakia is obtained with output gap employed as a measure of economic driving variable (Table 6.4). This variable is not significant. On the other side, both parameters, α_b and α_f , are significant with estimates 0.27 and 0.45 respectively. According to results reached, inflation persistence is more due to expectations about future inflation than to backward-looking behaviour. The estimated coefficients α_b and α_f do not sum to near unity.

The value of F-statistics is a bit higher than 10 (11.06), while the corresponding p-value of the Hansen J-test is 0.06. It could be the case that some of the instruments play endogenous role. However, we were not able to reach better quality model with them.

Table 6.4 Estimation of NKPC for Slovakia using output gap (January, 1997-July, 2010)

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

Notes: The full instruments set 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.

Romania

Real gross wage inflation was taken as inflation forcing variable while estimating NKPC in Romania. It has significant and positive influence on inflation dynamics (Table 6.5). Estimate on backward-looking element is in interval 0.31-0.40, while the portion of forward-looking behavior is in range 0.59-0.67. Estimates slightly depend on the method applied, but they uniformly suggest more important role of expectations about future inflation in determining inflation dynamics.

Table 6.5 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
	0.314	0.668	0.031	10.118	0.665
GMM	(0.095)	(0.105)	(0.009)		
	0.399	0.586	0.039	3.375	
LIML	(0.130)	(0.138)	(0.013)		

Notes: The full instruments set 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).

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

Serbia

Estimated NKPC for Serbia includes real broad money growth as inflation driving variable (Table 6.6). It appears as significant regressor, but only if GMM is applied. However, both tests that check the appropriateness of instruments clearly accept their validity and thus the implementation of GMM. Parameters on past and future inflation are highly significant in both versions: parameter α_b is estimated to be around 0.35-0.36, while estimate of α_f is around 0.51-0.55. Forward-looking weight is higher then weight on backward behavior.

Table 6.6 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
	0.353	0.513	0.046	10.071	0.371
GMM	(0.061)	(0.082)	(0.027)		
	0.361	0.552	0.053	2.570	
LIML	(0.156)	(0.142)	(0.367)		

Notes: The full instruments set 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.

In summary, as shown in Table 6.7, inflation persistence is driven both by expectations about future inflation and lagged inflation. The contribution of backward term is of relatively smaller importance. The effect of lagged inflation is further assessed by the value of root g_1 that appears of moderate size across countries. The highest influence of lagged inflation is calculated for Romania, 0.65, while the lowest values are reached for Czech Republic (0.26) and Slovakia (0.32). For the remaining three countries this estimate takes a value from interval 0.42-0.49. Finally, in five out of six NKPC models (Slovakia is an exception) significant impact of inflation driving variable has been detected. Therefore, results reached are supportive for the validity of NKPC for CSEECs.

Table 6.7 Summary from estimated NKPC models

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

7. Conclusions

Paper offers two set of results concerning inflation persistence in selected CEECs. First, detailed analysis of correlation pattern in inflation dynamics is undertaken within the univariate approach that allows for random changes of economic regimes. It is revealed that inflation persistence is still of moderate to high magnitude in Hungary, Poland, Romania and Serbia. In Slovakia and Czech Republic inflation persistence is estimated to be of smaller order. It is documented that changes in inflation persistence often correspond to changes in variability and mean of inflation. Contrary to countries from Central Europe, two Balkan countries, Romania and Serbia, exhibit more frequent switches of regimes with different level of inflation persistence. This indicates higher sensitivity of these economies to inflation random shocks.

Second, we provide evidence that NKPC represents a valid structural approach to describe inflation dynamics in this region. In all six cases weights on backward and forward looking behavior are significant, while the impact of driving variable was insignificant only once. We found that significant influence of economic driving variable is captured by real gross wage inflation and real broad money growth. Our estimates show that in general forward-looking element has slightly higher magnitude than backward-looking term in determining inflation persistence.

Relatively high inflation persistence suggests that agents employ simple adaptive expectations in price setting which is often the case when monetary policy lacks its credibility. In this situation the appropriate type of monetary policy should be inertial such that reduction in inflation is associated with output loss. On the other side, as NKPC indicates that expectations about future inflation cannot be neglected, it seems that agents in this region are in learning process. Therefore, expectations formation is not exclusively adaptive. Monetary policy could be partly effective if it was able to change expectations about future inflation that are found to be present. It is fair to say that monetary authorities in these countries should take into consideration all instruments available for conducting credible policy as implementing only one may not be enough to reach announced inflation target level.

Our analysis is based on monthly data which provide accurate insight into the inflation dynamics without inducing false autocorrelation either in univariate or structural set-up. We opt to work with switching regime time-varying framework in univariate models to discover specific features of time-series structure in inflation. We did not proceed with similar framework when NKPC is estimated, since included driving variable has a potential to capture information about the changes. In addition, estimation of NKPC is complicated enough such that linear form is the preferred one.

Given that most researches on this topic take quarterly or yearly data, our results are not directly comparable to the existing ones. Comparison among current euro zone countries and economies covered here is relevant as enables assessment of extent to which inflation dynamics in CSEES might converge to that of euro zone. We can take the case of Slovakia as a benchmark, for which we have an evidence of the lowest persistence level, while the weight on forward-looking term is almost double the weight on lagged element. Let us emphasize again that except Czech Republic other economies in the region are characterized by higher persistence and that forward-looking element is not as dominant in comparison to the backward term as in case of Slovakia. There is no enough evidence to conclude that inflation dynamics in the countries analyzed is likely to converge to the inflation dynamics of the current euro zone members. In order to strengthen our conclusion we plan to perform similar econometric work for more euro zone countries.

8. Literature

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Table A1. Data Definitions and Sources

Variable	Definition	Sources
inflation rate, based on CPI base monthly index (P_t).	transformation: $\pi_t = 100(\log P_t - \log P_{t-1})$	**OECD data:** http://stats.oecd.org/Index.aspx **data for Romania:** WIIW 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
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.en.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.	**WIIW Monthly Database on Eastern Europe:** http://mdb.wiiw.ac.at/
real braod money, (BMR_t)	transformation: broad money (base index) /CPI index, monthly data.	**OECD data:** http://stats.oecd.org/Index.aspx **data for Serbia:** National Bank of Serbia http://www.nbs.rs/export/internet/english/80/index.html
growth of real broad money	transformation: $\Delta BMR_t = BMR_t - BMR_{t-1}$	
real gross wages, (WR_t)	transformation: gross wages (total economy) /CPI index, monthly data.	**OECD data:** http://stats.oecd.org/Index.aspx **data for Romania:** WIIW 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: $\Delta WR_t = WR_t - WR_{t-1}$	
real exchange rate to euro	monthly averages, adjusted for domestic and EU inflation.	**WIIW Monthly Database on Eastern Europe:** http://mdb.wiiw.ac.at/