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EURO AREA AND NEW EU MEMBER STATES

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In 2007 all ECB publications feature a motif taken from the €20 banknote.

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2 Corresponding author: Czech National Bank and CERGE-EI, P.O. Box 882, Politickych veznu 7, 111 21 Praha 1, Czech Republic; e-mail: michal.franta@cerge-ei.cz

3 Czech National Bank and CERGE-EI, P.O. Box 882, Politickych veznu 7, 111 21 Praha 1, Czech Republic; e-mail: branislav.saxa@cerge-ei.cz. Branislav Saxa worked on the paper during his internship at the European Central Bank.

4 Czech National Bank and Charles University, Postal address: Czech National Bank, Na Příkopě 28, 115 03 Prague 1, Czech Republic; e-mail: katerina.smidkova@cnb.cz

Address
Kaiserstrasse 29
60311 Frankfurt am Main, Germany

Postal address
Postfach 16 03 19
60066 Frankfurt am Main, Germany

Telephone
+49 69 1344 0

Website
http://www.ecb.europa.eu

Fax
+49 69 1344 6000

Telex
411 144 ecb d

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CONTENTS

Abstract

Is inflation persistence in the new EU Member States (NMS) comparable to that in the euro area countries? We argue that persistence may not be as different between the two country groups as one might expect. We confirm that one should work carefully with the usual estimation methods when analyzing the NMS, given the scope of the convergence process they went through. We show that due to frequent breaks in inflation time series in the NMS, parametric statistical measures assuming a constant mean deliver substantially higher persistence estimates for the NMS than for the euro area countries. Employing time-varying mean leads to the reversal of this result and suggests similar or lower inflation persistence for the NMS compared to euro area countries. Structural measures show that backward-looking behavior may be more important component in explaining inflation dynamics in the NMS than in the euro area countries.

JEL Classification: E31, C22, C11, C32;

Keywords: Inflation persistence, new Member States, time-varying mean, New Hybrid Phillips curve

Non-technical summary

According to our findings, inflation persistence in the group of new EU member states (NMS) is not considerably different from inflation persistence in the current euro area countries. If it was significantly different, fulfillment of the Maastricht criteria could be complicated, for example due to the asymmetric impact of common shocks, such as large changes in oil prices, on the two groups of countries. Moreover, differences in inflation persistence between the two groups could also suggest that convergence of the NMS towards the current euro area countries might slow down after they adopt the euro.

Inflation persistence was put forward as one of the causes of divergence when persistent inflation differentials among the euro area countries were debated several years ago. During this debate, it was pointed out that once the NMS adopt the euro, they might also face the problem of inflation differentials. Subsequently, several studies presented estimates of inflation persistence in the NMS in order to see if high inflation persistence could be a matter of concern for the NMS. These studies work mostly with micro data. A few studies that take macroeconomic approach focus solely on one country. As a result, the empirical studies available so far on NMS do not provide sufficient information for a cross-country comparison.

Our empirical cross-country analysis takes the macroeconomic approach. It is built around the same definition of inflation persistence as that employed by the Eurosystem Inflation Persistence Network. According to this definition, the higher the inflation persistence is, the longer it takes for inflation to return to its equilibrium after a shock hits the economy. We argue that it must be checked carefully if the equilibrium value of inflation can be assumed constant for the NMS, which have gone through significant structural changes during the transformation process, including changes in monetary policy regimes, and which continue to converge toward the current euro area countries. It follows that out of the three sources of inflation persistence (extrinsic, intrinsic, and expectation-driven), expectation-driven persistence – which is closely connected to inflation targets perceived by public – should be distinguished from the other two sources. Therefore, we pay considerable attention to the choice of persistence measure.

We analyze the following statistical measures of inflation persistence: measures based on autoregressive models with constant means, autoregressive models with time varying means, and autoregressive fractionally integrated moving average models. These models differ in their approach to the equilibrium value of inflation and the underlying value of the perceived inflation target. In addition to the statistical measures, we consider structural measures of inflation persistence. On the one hand, the results based on structural measures are more powerful because their estimates of inflation persistence in the NMS are relevant for both the pre-euro and euro period, and in addition they enable us to distinguish fully all three sources of inflation persistence. On the other hand, it is a real challenge to obtain these results, given relatively poor data availability and small size of the available data samples.

We start our analysis by estimating autoregressive models with constant means for GDP deflators of our two groups of countries: the NMS (the Czech Republic, Hungary, Poland, and Slovakia) and the current euro area countries (Belgium, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal, and Spain). As suspected, the estimates of inflation persistence in the four NMS are higher than in the current euro area countries. This result is robust to the choice of estimation method used (the only exception being Greece if the Hansen method is used). However, we argue that these high estimates are biased upward due to the above-mentioned problem with the constant mean assumption. When the models with time-varying means are estimated, the inflation persistence in the four NMS is comparable to that in the current euro area countries. The remarkable estimated downward shift in perceived inflation targets in the NMS is the reason why the models with time-varying means are superior to the ones with constant means. These models are also superior to the autoregressive fractionally integrated moving average models, according to our findings.

The superiority of the time-varying mean models suggests that expectations have been an important source of inflation persistence in the NMS since 1993, the starting year of our data sample. Since that time, the perceived inflation targets in the NMS have converged from 3–7% to the levels common in the current euro area countries (1–3%), according to our estimates. However, the available estimates of the New Hybrid Phillips Curve, including ours for the Czech Republic, indicate that inflation behavior are still more backward-looking in the NMS than in the current euro area countries. To conclude, while inflation persistence due to nominal and real rigidities in the NMS is comparable to that in the current euro area countries, expectation-based persistence is higher. Anchoring of inflation expectations will therefore be an important component of the euro adoption process for the NMS.

1. Introduction

In this paper, we provide input into the discussion concerning the readiness of the new EU Member States (NMS) to adopt the euro. With regards to euro adoption, the NMS face two closely related challenges. First, they need to fulfill the Maastricht criteria, including the one inflation. Second, they need to adapt their economies to life with the euro. Inflation persistence differences between the euro area countries and the NMS can represent an obstacle for successful dealing with both challenges.

The issue of differences in inflation persistence was raised by various studies[1] in 2002 when inflation divergence among the current euro area members was observed. These studies show that the inflation convergence reached prior to adopting the euro has not been sustained among the current euro area members since 1998, and they point out that inflation persistence is one of the most prominent reasons. Euro adoption candidates, therefore, have to deal with the extent of the national inflation persistence in order to prevent inflation from diverging after the euro adoption.

Furthermore, inflation persistence can influence the fulfillment of the Maastricht criteria, which is an issue for the NMS before and even after euro adoption. High inflation persistence corresponds to the slow return of inflation to its long-run value after a shock (e.g. an oil shock) occurs. Countries with a high speed of inflation adjustment determine, according to the Maastricht criterion on inflation stability, the upper inflation bound for countries that can exhibit a slower inflation rate. Countries with high inflation persistence could struggle to meet the criterion should a common shock hit European countries.

To our knowledge, there are only a few studies assessing inflation persistence in the NMS. Available results, mainly based on micro data, indicate that inflation persistence in the NMS could be higher than in the current euro area members, although in some countries it decreases slowly over time. Since disaggregate evidence makes international comparison problematic, we carry out our analysis using inflation aggregates. On the other hand, inflation aggregates can suffer from an aggregation bias, i.e. inflation aggregates exhibit higher persistence than the particular components included.

In this paper, we use several approaches to define and estimate inflation persistence in order to discuss thoroughly the appropriateness of various measures for the measurement of inflation persistence in the NMS. Furthermore, we attempt to choose the measure that enables international comparison of the euro area countries and the NMS. The list of the inflation persistence measures employed in this study is depicted by the following scheme:

Scheme of considered inflation persistence measures

Statistical measures – Parametric	i) Autoregressive model with constant mean (naive estimates)
	ii) Autoregressive model with time-varying mean
	iii) Autoregressive fractionally integrated moving average model (ARFIMA)
Structural measures	iv) Estimates of the New Hybrid Phillips Curve (NHPC)

[1] Section 2 provides a literature overview of papers related to inflation persistence in this context.

We adopt first of all a purely statistical approach and estimate several parametric measures based on the sum of autoregressive coefficients and impulse response functions, before employing a structural approach that provides the estimate of inflation inertia based on structural parameters. These approaches have already been applied to the analysis of inflation persistence in the current euro area members. Hence, we can compare our results for this control group with those of previously published work.

The first group of parametric statistical measures of inflation persistence assume a constant mean. Four NMS[2] in our sample score highly among the EU members as far as inflation persistence is concerned. The estimated inflation persistence for the these NMS is close to one, once the confidence intervals have been taken into account. This finding is in line with the available micro studies on inflation persistence in the NMS and with previously published research on inflation persistence in the current euro area Member States.

Our second more sophisticated statistical measure of inflation persistence gives, however, another picture. It allows for a time-varying mean. We separate the impact of persistence in nominal contracts and persistence in the real economy factors influencing inflation (intrinsic and extrinsic persistence) from the impact of inflation expectations and monetary policy regime changes (the two being inseparable in our model). We find that according to this measure the estimates of inflation persistence in the NMS are comparable to those in the current euro area members.

In our third statistical measure, we focus on the measure that is built on the autoregressive fractionally integrated moving average (ARFIMA) representation of the inflation process. A stationary process with parameter instability and a fractionally integrated process can look very similar (mean reversion). Nevertheless, the implications of the two representations of the inflation process for the measurement of inflation persistence differ. Values of the impulse response function based on ARFIMA suggest that persistence in some NMS is higher than in the rest of the sample. Additional statistical tests suggest that assuming a stationary process with breaks is a preferable assumption to fractionally integrated models for almost all considered countries.

Various statistical measures of inflation persistence introduced so far provide a complex picture of the actual extent of inflation inertia in the NMS in comparison with the euro area. It is worth noting that these measures can mainly serve as input for the debate about the fulfillment of the Maastricht criteria. If values are comparable for two groups of countries (euro area countries and NMS), it could be less difficult for the NMS to fulfill the Maastricht criterion on inflation, for example. However, these measures cannot serve as a basis for inferences about the country-specific effects of common monetary policy in the euro area extended to the NMS. The argument draws on the Lucas critique that views structural (deep) parameters as the only appropriate measure underlying the discussion on the consequences of unequal inflation persistence after the monetary policy regime switch.

[2] In our analysis, the NMS are represented by four countries (the Czech Republic, Hungary, Poland and Slovakia) since for these four NMS the complete data needed for the inflation persistence analysis are available.

Therefore, as a last approach to measuring inflation persistence, we introduce a model based on deep parameters that allows an international comparison of the extent of inflation inertia. We estimate the hybrid version of the structural NHPC for the Czech Republic, Poland and Slovakia, and we compare estimation results with existing studies for Hungary and the euro area. The structural measure suggests that the influence of expected future inflation for current inflation does not predominate over the influence of past inflation in the Czech Republic and Hungary.

The structure of the paper is as follows. Section 2 reviews the available literature on the topic placing special emphasis on the relevance of inflation persistence in the NMS. Section 3 describes the adopted approaches to measuring and estimating inflation persistence. Section 4 reports on and discusses the results of these alternative estimates. Section 5 concludes.

2. Related literature

Many studies focusing on the current euro area members have clearly demonstrated why inflation persistence is worth analyzing. Angeloni and Ehrmann (2004) identify the relationship between inflation persistence and persistence of inflation differentials across the euro area countries. In addition, several reports from the beginning of this decade point out that euro area monetary policy can and should target price stability in the whole euro area only, and that persistent inflation differences can cause problems for the economies of euro area members. Specifically, EC (2002, 2004), ECB (2003), OECD (2002), and IMF (2002) state that adequate national structural reforms should be adopted in countries with high inflation persistence, given the potential risks of divergence. Inflation differences and the speed of convergence have also been discussed in relation to EU enlargement. For example, some of the results in Egert et al. (2004), Bjorksten (2002) and Ca'Zorzi and De Santis (2003) suggest that inflation differences may prevail longer inside the euro area once the NMS introduce the euro.

Most of the available research on inflation persistence in NMS is based on micro data. Micro analysis is available for the Czech Republic, Hungary, Poland and Slovakia in Babetskii, Coricelli and Horváth (2006), Ratfai (2006), Konieczny and Skrzypacz (2005), and Coricelli and Horváth (2006), respectively. Two studies that draw on the macroeconomic aggregates are Darvas and Varga (2007) and Lendvai (2005). These studies focus on Hungary. Some of the results signal that high inflation persistence can indeed be a problem for some NMS.

The most often employed measures to analyze inflation persistence include the sum of autoregressive coefficients, the largest autoregressive root, half-life and spectral density at frequency zero (see Marques 2004 for a summary). In univariate analysis, the mean of the inflation process is often assumed to be constant. Yet a problem arises if a structural change, such as a change in the inflation target or an administrative price regulation, occurs and induces a change in the medium-run mean of inflation. If inflation is assumed to follow an autoregressive process, then not accounting for such a change causes upward bias in the estimated autoregressive coefficient, as shown by Perron (1989). We argue that this is a serious problem for the NMS and thus should be taken into account when estimating inflation persistence in the NMS. The choice of assumption about the mean of inflation therefore becomes one of the main issues in this paper.

More sophisticated approaches to measure inflation persistence also analyze the sources of inflation persistence. Angeloni et al. (2006) distinguish three types of inflation persistence. Intrinsic inflation persistence relates to nominal rigidities and to the way wages and prices are set. Extrinsic inflation persistence stems from persistence in the inflation-driving real variables (e.g. output gap). Finally, expectations-based inflation persistence is driven by differences between public perceptions about the inflation target and the central bank's true (explicit or implicit) inflation target. Bilke (2005) notes that the conduct of monetary policy can lead to changes in the mean of inflation time series and thus influence inflation persistence estimates. Dossche and Everaert (2005) set up a model that allows these three mentioned sources of inflation persistence to be distinguished. Moreover, their model controls for shifts in the inflation mean caused by monetary policy changes. This approach is relevant for the NMS, since it estimates inflation persistence net of expectations-based persistence and persistence related to the effects of monetary policy.

As explained earlier in this section, potential structural breaks present a serious problem for the estimation of persistence. Some recent empirical studies have approached this problem by allowing for structural breaks in inflation series. Levin and Piger (2004) estimate an autoregressive model first with the assumption of a constant mean, and subsequently allow for one structural break in the mean of inflation. They find evidence for structural breaks in several industrial countries during the period 1984-2003, and they show that allowing for a single break decreases the estimates of persistence significantly. Cecchetti and Debelle (2006) go further and estimate inflation persistence allowing for no break or one, two or three breaks. The authors suggest that allowing for one break in general reduces persistence estimates in most countries while allowing for more breaks reduces estimates further, although by a smaller amount.

Marques (2004) stresses that it is more natural to assume a time-varying mean of inflation rather than to assume a constant mean or to search for breaks in the mean of inflation. In his analysis of US and euro area inflation, Marques considers several treatments for the mean of inflation, including the application of an HP filter and a moving average. In general, his results confirm that more flexibility assumed for the mean of inflation delivers lower estimates of persistence. Similar results for the US and the euro area are provided by Dossche and Everaert (2005), who model the time-varying mean as an AR process. Benati (2006), in the framework of AR(p) representation of inflation series for 21 countries, allows for random-walk time-varying parameters. Finally, Darvas and Varga (2007) use time-varying coefficient autoregressive models to investigate Hungarian inflation persistence.

Regarding the structural breaks in parameters of the inflation process, the literature points out that stationary processes with structural breaks and fractionally integrated processes can exhibit similar time behavior along with different properties regarding persistence. The application of the fractionally integrated approach in the context of inflation persistence is introduced in Gadea and Mayoral (2006). In addition to formal tests of inflation time series, authors show how fractionally integrated behavior can emerge in heterogeneous agent sticky price models.

In contrast to purely statistical measures, our structural model of short-run inflation dynamics is derived from micro principles. The hybrid version of the model is introduced in Galí and Gertler (1999). The authors assume that a fraction of monopolistically competitive firms use a backward-looking rule of thumb to set prices. Furthermore, they argue that in the derived NHPC, real marginal cost should be used instead of the usual output gap as a measure of real economic activity

Galí and Gertler (1999) estimate the NHPC on US quarterly data for the period 1960:1-1997:4. They find that the forward looking behavior predominates in comparison with the backward looking behavior. Moreover, they obtain significant and positive coefficient in the term capturing real economic activity. Galí, Gertler and López-Salido (2001) extend the framework laid down in Galí and Gertler (1999) for the euro area. They consider the period 1970:1-1998:2, and their estimation results suggest backward-looking price setting behavior is even less important in the euro area than in the US.

Both studies use the generalized methods of moments (GMM) approach to estimate of NHPC. The use of GMM, however, has been much criticized for several reasons. The issue of weak instruments is addressed, for example, in Ma (2002). Zhang et al. (2006) also argue that the presence of serial correlation in errors influences the validity of lagged values of inflation and real variable as instruments. Zhang et al. (2006) estimate the NHPC for the US quarterly data for the period 1960:1-2005:1, and question the robustness of results in Galí and Gertler (1999) regarding the instrument set employed.

Lendvai (2005) is the only attempt to estimate a structural Phillips curve for a NMS representative. She considers quarterly Hungarian data covering the period 1995:1 – 2004:1. Results suggest that inflation exhibits higher inflation inertia in Hungary than in the euro area.

The various approaches to the measurement of inflation persistence introduced in the economic literature allow us to discuss thoroughly the extent of inflation inertia in the NMS in comparison with the euro area countries. Moreover, we discuss specific features of inflation that need to be taken into account in the case of the countries that have experienced structural breaks and transition to the new steady state.

3. Stylized facts and models for estimating inflation persistence

In this section we introduce various approaches to measuring inflation persistence. We start with naïve estimates that assume a constant mean of inflation, then move on to models that relax the constant mean assumption. We also discuss ARFIMA models. Finally, we focus on the estimations of the structural NHPC.

The literature provides several definitions of inflation persistence.[3] We stick to the usual approach that relates inflation persistence to the speed at which inflation converges to its equilibrium value after a shock. Intuitively, inflation persistence is

[3] See, for example, Batini (2002).

high if the inflation series does not frequently oscillate around its mean.[4] So, simple visual inspection of inflation plots for various countries (see Appendix 3) yields the first idea about the persistence of inflation in the euro area countries and the NMS. In addition, Table 1 reports the number of times that inflation series switched from above to below their means and vice versa.

Table 1: Number of crosses of inflation means.

Period	Czech Rep.	Hungary	Poland	Slovakia	EU12
1993:2-2006:1	11	25	11	16	19
2001:1-2006:1	6	16	11	11	11
	Belgium	Finland	France	Germany	Greece
1993:2-2006:1	28	18	16	17	19
2001:1-2006:1	16	9	9	13	12
	Ireland	Italy	Netherlands	Portugal	Spain
1993:2-2006:1	27	31	27	29	15
2001:1-2006:1	13	11	11	11	12

Source: Own calculations based on OECD OEO database.

Note: Inflation rates for Hungary available since 1995:1.

Table 1 illustrates an issue that often arises when we employ various approaches to measuring inflation persistence in the NMS. For the whole sample (1993:2-2006:1), inflation series for the NMS cross their means less frequently than inflation series for the euro area countries. According to the mentioned definition, fewer switches indicate higher inflation persistence for the NMS in comparison to the current euro area members. However, this is not necessarily so since the NMS have gone through a transformation period, during which the high initial values of inflation led to high means of inflation. In addition, price levels in the NMS have been converging to those of the euro area members. Both factors – transformation as well as convergence – may weaken the link between the persistence and the frequency of mean crosses. We indeed observe that once we restrict the sample to the period 2001:1-2006:1, the number of crosses for the NMS and euro area members is comparable (see Table 1).

Going back to the definition of inflation persistence, the focus is on the concept of the equilibrium value of inflation. Some measures of persistence introduced in the following paragraphs view the equilibrium value from a long-run perspective, while others focus rather on the medium-run[5]. Table 1 implies that the appropriateness of various measures of persistence for the NMS arises from their ability to take into account specific attributes of inflation processes in the NMS.

[4] Marques (2004) shows the inverse relationship between inflation persistence and mean reversion in the case we model the inflation process as an autoregressive process of order *k*.

[5] We find it useful to distinguish these two time horizons when discussing inflation persistence in the NMS, since long-run and medium-run equilibria may differ in period of convergence. For discussion on the importance of time horizons when dealing with the concept of equilibrium, see Driver and Westaway (2005).

3.1 Statistical measures – parametric (autoregressive models)

(i) Constant mean (naive estimate)

The most widely used measure of persistence across the literature, the sum of autoregressive coefficients, is based on the assumption that inflation follows a stationary autoregressive process of order K:

$$\pi_t = \mu + \sum_{i=1}^{K} \alpha_i \pi_{t-i} + \varepsilon_t \tag{1}$$

The sum of autoregressive coefficients is then defined as:

$$\rho_K = \sum_{i=1}^{K} \alpha_i \tag{2}$$

where π_t denotes observed inflation rate at time t. We proceed as follows. First, we obtain OLS estimates of ρ_K for specifications with lag lengths $K = 1,..,5$. The preferred number of lags is then chosen according to AIC and BIC criteria. Second, we apply Hansen's (1999) grid bootstrap procedure[6] on the same data to estimate median unbiased ρ_K and its 90% confidence intervals, again for lag lengths $K = 1,..,5$.

(ii) Time-varying mean

We draw on the model introduced in Dossche and Everaert (2005), who estimate univariate and multivariate time series models. The univariate time series model should put the naïve statistical measures from the previous subsection into a broader perspective, since the model enables us to identify the part of inflation persistence that stems from monetary policy actions.

The model Dossche and Everaert (2005) start with has the following form:

$$\pi_{t+1}^T = \pi_t^T + \eta_{1t} \tag{3}$$

$$\pi_{t+1}^P = (1-\delta)\pi_t^P + \delta\pi_{t+1}^T + \eta_{2t}, 0 < \delta < 1, \tag{4}$$

$$\pi_t = \left(1 - \sum_{i=1}^{4} \varphi_i\right)\pi_t^P + \sum_{i=1}^{4} \varphi_i L^i \pi_t + \beta_1 z_{t-1} + \varepsilon_{1t}, \sum_{i=1}^{4} \varphi_i < 1, \tag{5}$$

where π_t^T is the central bank's inflation target, π_t^P is the inflation target as perceived by the public, z_t stands for the output gap, and disturbances are mutually independent zero-mean white noise processes.

The central bank's inflation target is modeled as a random walk process in equation (3). The model assumes this equation even if the central bank does not target inflation

[6] Hansen's (1999) grid bootstrap procedure is used in several recent studies on inflation persistence, e.g. Benati (2006), Levin and Piger (2004), and Gadzinski and Orlandi (2004).

explicitly. Some countries have adopted inflation targeting during the period of interest (e.g. the Czech Republic in 1997/98). However, we do not impose known targets on the model.

Equation (4) captures the relationship between the central bank's inflation target and the target as perceived by the public.[7] The parameter δ measures the expectation-based persistence – a value close to zero indicates that the public forms its inflation expectations in a backward-looking manner. The effect of a shock to inflation is then prolonged via inflation expectations. On the other hand a parameter value close to one shows that a central bank is highly credible in communicating its inflation target.[8]

Equation (5) takes a form close to the traditional Phillips curve. Private inflation expectations are represented by the perceived inflation target. The sum of autoregressive coefficients captures intrinsic inflation persistence.

We make two identifying assumptions. First, we assume in accordance with Dossche and Everaert (2005) that $\beta_1 = 0$.[9] Second, to keep the estimation simple we also adopt the following assumption: $\eta_{2t} = 0$ for all *t*.

If we incorporate these assumptions, the basic version of the model has the following form:

$$\pi_t = \left(1 - \sum_{i=1}^{q} \varphi_i\right)\pi_t^P + \sum_{i=1}^{q} \varphi_i L^i \pi_t + \varepsilon_{1t} \qquad \varepsilon_{1t} \approx N(0, \sigma_\varepsilon^2)$$

$$\pi_{t+1}^P = (2-\delta)\pi_t^P + (\delta - 1)\pi_{t-1}^P + \delta\eta_{1t} \qquad \eta_{1t} \approx N(0, \sigma_\eta^2)$$

Since the model includes unobservable components (π_t^P) we transform the system into the state space form and use state space analysis methods.

$$\begin{bmatrix} \pi_{t+1}^P \\ \pi_t^P \end{bmatrix} = \begin{bmatrix} 2-\delta, \delta - 1 \\ 1,0 \end{bmatrix} \begin{bmatrix} \pi_t^P \\ \pi_{t-1}^P \end{bmatrix} + \begin{bmatrix} \delta \\ 0 \end{bmatrix} \eta_{1t}$$

[7] The model equalizes the inflation target as perceived by the public, and public inflation expectations.
[8] There is also another possible interpretation of the formula. If the public forecasts inflation ($\pi_{t+1|t}^{forecast}$) in the same way as the central bank (irrespective of what the announced inflation target is) and the central bank behaves such that the inflation forecast equals the inflation target ($\pi_{t+1|t}^{forecast} = \pi_{t+1}^T$), then the parameter δ captures the fraction of forward-looking members of the public.
[9] This assumption implies that the resulting form of the Phillips curve is equivalent to the assumption that the data-generating process for inflation has the following form:

$$\pi_t = \mu_t + \sum_{i=1}^{4} \alpha_i (\pi_{t-i} - \mu_t) + \varepsilon_t,$$

where the time-varying mean equals the perceived inflation target. This formula is the starting point for parametric measures based on AR(p) representations of the data-generating process (assuming a constant intercept $\mu_t = \mu$).

$$\pi_t = \left[\left(1-\sum_{i=1}^{4}\varphi_i\right),0\right]\begin{bmatrix}\pi_t^P \\ \pi_{t-1}^P\end{bmatrix} + \sum_{i=1}^{4}\varphi_i \pi_{t-i} + \varepsilon_{1t}$$

To estimate the unobservable series of perceived inflation π_t^P we use the exact initial Kalman filter (the case of unknown initial conditions) as described for example in Koopman and Durbin (2003). The Kalman filtering assumes known coefficients; therefore we have to estimate them.

We follow Dossche and Everaert (2005) and use a Bayesian approach combined with the method of importance sampling.

(iii) ARFIMA model

While a shock has a permanent effect in I(1) models and disappears at an exponential rate in I(0) models, the fractionally integrated approach allows for richer representation by introducing the so-called fractional differencing parameter *d,* which can be any real number $d \neq 0$. The time series y_t follows so called ARFIMA(p,d,q) model if

$$\phi(L)(1-L)^d (y_t - \mu) = \theta(L)\varepsilon_t, \qquad (6)$$

where the roots of $\phi(L)$ and $\theta(L)$ lie outside the unit circle and ε_t is white noise.

As advocated by Baum et al. (1999) or Gadea and Mayoral (2006), the ARFIMA model could be appropriate representation of the stochastic behavior of inflation time series. ARFIMA allows a high degree of persistence without assuming a unit root (i.e. I(1)) character of the process. We follow Gadea and Mayoral (2006) and estimate parameter *d* from (6) as well as the impulse response function of the appropriate ARFIMA model.

Furthermore, we employ the test suggested by Mayoral (2004) which tests the hypothesis of time series following fractionally differentiated process of order *d* versus a stationary process with breaks. Unlike Gadea and Mayoral (2006), we allow for a break not only in the level but also in the trend, to reflect the convergence process observed in the parts of inflation series of some countries.

The test statistics have the following form:

$$R(d) = T^{1-2d} \frac{\inf_{\omega\in\Omega}\left(\sum (y_t - \hat{\alpha}_1 - \hat{\delta}_1 DC_t - \hat{\beta}_1 t - \hat{\delta}_2 DT_t)^2\right)}{\sum (\Delta^d (y_t - \hat{\alpha}_0 - \hat{\beta}_0 t))^2}$$

where *d* is the order of differentiation, T is the number of periods, $\Omega = [0.15, 0.85]$ are trimming thresholds, *y* is a considered time series, $DC_t = 1$ if $t>\omega T$ and 0 otherwise, $DT_t = (t-T_B)$ if $t>\omega T$ and 0 otherwise. α_0, α_1, β_0, β_1, δ_1 and δ_2 are coefficients from appropriate regressions. Δ^d is the operator of differencing of order *d*. Critical values are computed according to Mayoral (2004).

A null hypothesis assumes a fractionally integrated process; an alternative hypothesis assumes a stationary process with breaks.

3.2 Structural measures

Both the theory and practical estimation of the structural Phillips curve is a subject of heightened debate in recent years, and no consensus concerning related issues has been achieved so far. We try to stick to the approaches used in the studies mentioned in the literature review to make the international comparison meaningful. However, we stress the possible weaknesses of the approach that are raised in the literature and that could affect the resulting estimates.

The aim of the structural Phillips curve estimation is to find out a formula that captures short-run inflation dynamics, and consequently enables us to infer the degree of inflation inertia based on the estimation of the formula.

The parameters of the model introduced in Galí and Gertler (1999) are functions of three model primitives: the probability that a firm has to keep its price unchanged (θ) (degree of price rigidity), the fraction of backward-looking firms that set their price according to the price in the previous period adjusted for inflation (ω), and the discount factor (β).

The closed-economy version of the New Hybrid Phillips curve (NHPC) takes the following form:

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E_t \pi_{t+1} + \lambda mc_t$$

$$\gamma_b = \frac{\omega}{\phi}$$

$$\gamma_f = \frac{\beta\theta}{\phi}$$

$$\lambda = \frac{(1-\omega)(1-\theta)(1-\beta\theta)}{\phi}$$

$$\text{with } \phi = \theta + \omega[1-\theta(1-\beta)].$$

Here the variable mc_t represents the percentage deviation of the average real marginal cost from its steady-state value.

The ongoing debate on the theoretical and econometric issues regarding short-run inflation dynamics is even more pronounced for the NMS. Together with the issues mentioned above, one has to deal with incomplete time series, short time spans of data, and a convergence process in the NMS. Therefore, estimating the NHPC for post-transition countries involves some additional issues.

As post-transition countries have been experiencing a transition towards a new steady state, we use an HP filter to filter out non-business cycle frequencies and thus abstract from a convergence path. This approach can result in various biases (for a detailed discussion, see Lendvai 2005). In addition, Baum et al. (2003) point out that the

GMM estimator can exhibit poor properties in the case of small samples and we therefore follow Lendvai (2005) in employing a 2SLS estimator.

4. Results

In this section we provide results of inflation persistence measures introduced in the previous section. Like some previously quoted studies, we employ a seasonally adjusted annualized quarter-on-quarter rate of change of the GDP deflator to represent inflation in all estimates and computations. All remaining data are thoroughly described in Appendix 1. The time span considered covers the period 1993:2-2006:1, if not stated explicitly otherwise. In case of Hungary, data are available since 1995:2. Country abbreviations are also explained in Appendix 1.

We provide inflation persistence estimates for individual countries (not only for the whole EU12) since a direct comparison of persistence in individual NMS and euro area as a group could be misleading. As shown in Cecchetti and Debelle (2006) and discussed in Altissimo, Ehrmann and Smets (2006) and Batini (2002), aggregation of inflation indices leads to higher persistence estimates. This holds for aggregation from sectoral to country level as well as aggregation from country indices to euro area indices.

4.1 Statistical measures – parametric (autoregressive models)

(i) Constant mean

We start with the estimation of the sum of autoregressive coefficients.[10] The results of OLS estimates of ρ_K are reported in Table 2. Estimated persistence reaches 0.68 for Poland and 0.75-0.76 for the Czech Republic, Hungary and Slovakia. In contrast, persistence is estimated below 0.68 for all the other countries. Four NMS thus have higher estimates of inflation persistence than any other countries in the sample. A similar pattern (of the six countries with the highest persistence estimates in the sample, four are NMS) is confirmed by estimating the largest autoregressive roots (not reported here).

In Table 3, we report estimates of ρ_K obtained using Hansen's (1999) grid bootstrap procedure, including 90% confidence intervals. Figure 1 shows estimates and confidence intervals for the case of $k = 5$ lags. Although confidence intervals are wide and estimates embody considerable uncertainty, one pattern is robust across the number of considered lags: estimates of persistence in NMS are high and in most cases higher than persistence in euro area countries. In all five specifications with different lag lengths, four NMS rank among the six countries with the highest persistence estimates in the sample.

[10] Note that stationarity tests of inflation time series are included in the analysis. Estimates of the coefficients for the lag length equal to one (see the last column in the Table 3) show that we can reject the null of unit root for all countries at the 90% significance level.

Table 2: OLS estimates of ρ_K (inflation based on GDP deflator)				
	Preferred model according to AIC		Preferred model according to BIC	
	Number of AR lags	Sum of AR coefficients	Number of AR lags	Sum of AR coefficients
CZE	5	0.75	4	0.76
HUN	5	0.75	4	0.75
POL	4	0.68	4	0.68
SVK	2	0.75	2	0.75
EU12	3	0.66	3	0.66
BEL	2	0.13	2	0.13
ESP	4	0.59	1	0.26
FIN	1	0.33	1	0.33
FRA	1	0.43	1	0.43
GER	4	0.50	3	0.61
GRC	4	0.67	4	0.67
IRL	2	0.11	2	0.11
ITA	2	0.14	2	0.14
NLD	3	0.62	3	0.62
PRT	5	-0.16	5	-0.16

Estimates of persistence in the NMS based on the assumption of constant mean could, however, to some extent suffer from upward bias due to the impact of administrative price changes. Gadzinski and Orlandi (2004) as well as Levin and Piger (2004) show that administrative price changes (e.g. changes in VAT) increase the persistence estimates if they are not accounted for. Due to the transition process, NMS countries experienced numerous administrative price changes during the 1990s. Besides changes in VAT and excise taxes, gradual price deregulations influenced the prices of energy and housing. Since the frequency of these changes and a relatively short sample do not allow us to control for breaks in a way some other studies do, we adopt a different approach.[11]

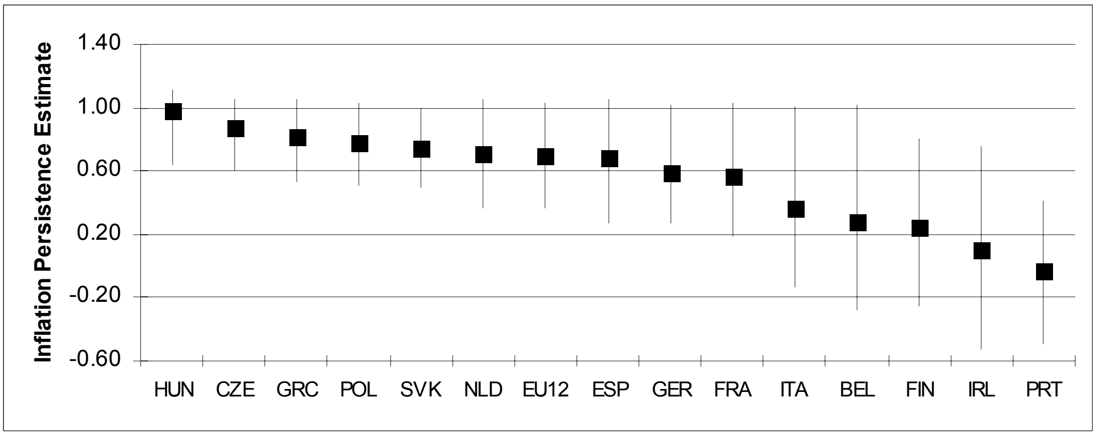
While we abandon the constant mean assumption in the next section, in Appendix 2 we present results of the same methodology as previously, this time applied to inflation based on non-food, non-energy CPI. The reason is that non-food, non-energy CPI inflation is supposed to be less influenced by price deregulations[12] and therefore allows for a better comparison of inflation persistence between the NMS and the rest of the sample. Nevertheless, even in the case of core inflation, the estimates of persistence in NMS are (with the exception of Slovakia) still higher than in most of the ten other countries.[13] Using Hansen's (1999) grid bootstrap estimation on the core inflation data, we can observe that inflation persistence in Slovakia is relatively low, whereas the Czech Republic, Hungary and Poland rank in the half of the sample with higher persistence, regardless of the number of lags (see Appendix 2 for tables and figures reporting resulting for core inflation).

[11] Fidrmuc and Tichit (2004) discuss the role of structural breaks in transition data. They attempt to detect structural breaks in a growth regression for data frequency and time period similar to ours.
[12] Prices of energy were among most heavily regulated prices in NMS over the transition period.
[13] Another reason for including non-food, non-energy CPI inflation is to examine the robustness of our results with respect to the choice of inflation time series.

Table 3: ρ_K its 90% confidence intervals estimated using Hansen's (1999) grid bootstrap procedure
(inflation based on GDP deflator)

	Lag length = 5			Lag length = 4			Lag length = 3			Lag length = 2			Lag length = 1		
	lower bound	mean	upper bound	lower bound	mean	upper bound	lower bound	mean	upper bound	lower bound	mean	upper bound	lower bound	mean	upper bound
CZE	0.60	**0.88**	1.05	0.62	**0.87**	1.04	0.49	**0.71**	0.96	0.52	**0.76**	1.01	0.43	**0.63**	0.84
HUN	0.64	**0.98**	1.11	0.60	**0.98**	1.10	0.57	**1.00**	1.08	0.44	**0.76**	1.04	0.11	**0.41**	0.68
POL	0.51	**0.78**	1.03	0.55	**0.83**	1.03	0.52	**0.79**	1.04	0.47	**0.69**	0.99	0.27	**0.50**	0.72
SVK	0.49	**0.74**	1.00	0.52	**0.74**	1.00	0.62	**0.85**	1.03	0.64	**0.84**	1.03	0.52	**0.70**	0.89
EU12	0.36	**0.70**	1.03	0.45	**0.78**	1.05	0.48	**0.80**	1.05	0.21	**0.50**	0.84	0.10	**0.33**	0.58
BEL	-0.28	**0.28**	1.02	-0.16	**0.38**	0.94	-0.33	**0.10**	0.54	-0.15	**0.20**	0.56	-0.39	**-0.15**	0.08
ESP	0.27	**0.69**	1.06	0.35	**0.79**	1.07	0.02	**0.37**	0.79	-0.03	**0.24**	0.54	0.05	**0.31**	0.54
FIN	-0.25	**0.24**	0.81	-0.24	**0.12**	0.56	-0.06	**0.29**	0.70	0.12	**0.41**	0.71	0.13	**0.35**	0.62
FRA	0.19	**0.57**	1.03	0.06	**0.36**	0.68	0.27	**0.56**	0.90	0.35	**0.62**	0.96	0.24	**0.47**	0.69
GER	0.27	**0.59**	1.02	0.28	**0.60**	0.98	0.43	**0.72**	1.03	0.16	**0.41**	0.71	0.13	**0.35**	0.58
GRC	0.53	**0.82**	1.06	0.52	**0.54**	1.07	0.07	**0.41**	0.80	0.06	**0.35**	0.63	0.02	**0.24**	0.46
IRL	-0.53	**0.10**	0.76	-0.35	**0.16**	0.73	-0.25	**0.21**	0.74	-0.19	**0.19**	0.57	-0.53	**-0.31**	-0.09
ITA	-0.13	**0.37**	1.01	0.10	**0.56**	1.05	0.01	**0.46**	1.02	-0.15	**0.20**	0.58	-0.32	**-0.06**	0.20
NLD	0.36	**0.71**	1.05	0.49	**0.93**	1.08	0.42	**0.77**	1.05	0.20	**0.52**	0.88	-0.04	**0.20**	0.44
PRT	-0.49	**-0.03**	0.41	-0.20	**0.20**	0.65	-0.23	**0.14**	0.53	-0.10	**0.22**	0.59	-0.18	**0.04**	0.30

Figure 1: Inflation based on GDP deflator, ρ estimate and its 90% confidence intervals
(lag length = 5, Hansen's (1999) grid bootstrap procedure)



(ii) Time-varying mean

In this section, we present the results of the autoregressive model of inflation, allowing for time-varying mean. The model measures inflation persistence net of the effects of a monetary policy authority.

Tables 4a and 4b report parameter estimates and 90% confidence intervals that are obtained by the method of importance sampling.[14] Intrinsic inflation persistence (sum of AR coefficients) and expectation-based inflation persistence (δ) are statistically significant.

TABLE 4a: Estimation results of the model with time varying mean – NMS

	Czech Republic			Poland			Slovakia		
	lower bound	*parameter estimate*	*upper bound*	*Lower bound*	*parameter estimate*	*upper bound*	*lower bound*	*parameter estimate*	*upper bound*
φ_1	0.19	0.31	0.54	-0.01	-0.01	-0.01	0.06	0.21	0.32
φ_2	-0.13	0.01	0.09	0.05	0.05	0.05	0.06	0.18	0.29
φ_3	-0.06	0.05	0.18	0.22	0.22	0.22	-0.13	-0.01	0.14
φ_4	-0.21	-0.12	-0.07	-0.14	-0.14	-0.14	-0.21	-0.10	0.00
$\sum_{i=1}^{4}\varphi_i$	-0.12	0.26	0.49	0.12	0.12	0.12	0.10	0.28	0.49
δ	0.16	0.26	0.33	0.07	0.07	0.07	0.15	0.27	0.39
σ_ε^2	2.11	2.37	2.74	2.80	2.80	2.80	1.80	2.04	2.33
σ_η^2	0.05	0.13	0.22	0.12	0.12	0.12	0.02	0.08	0.17

[14] During the estimation of coefficients for the filtering algorithm we encountered two main numerical problems. First, for Hungary and Ireland the algorithm for finding the minimum of constrained nonlinear multivariable function does not converge in a reasonable number of iterations. We therefore do not report estimation results for the two countries. Note that minimization is a first step in a method of importance sampling to obtain importance density. Second, for Greece and Poland we take only a subsample since the full sample Hessian matrix obtained during the minimization is too large to be useful for importance density. Even for the restricted sample, the Hessian matrix for Poland is quite large and thus lower and upper bounds of 90% confidence intervals differ little.

TABLE 4b: Estimation results of the model with time varying mean – EU12

	Belgium			Finland			France		
	lower bound	*parameter estimate*	*upper bound*	*lower bound*	*parameter estimate*	*upper bound*	*lower bound*	*parameter estimate*	*upper bound*
φ_1	-0.06	0.08	0.23	0.16	0.19	0.25	0.12	0.26	0.40
φ_2	0.05	0.20	0.35	-0.01	0.06	0.13	0.01	0.16	0.30
φ_3	-0.15	0.00	0.14	-0.12	-0.04	0.05	-0.12	0.02	0.16
φ_4	0.04	0.18	0.32	-0.25	-0.15	-0.07	-0.22	-0.07	0.08
$\sum_{i=1}^{4}\varphi_i$	0.11	0.45	0.86	-0.09	0.07	0.21	0.07	0.37	0.67
δ	0.06	0.21	0.37	0.15	0.23	0.39	0.09	0.22	0.37
σ_ε^2	0.87	1.15	1.46	2.10	2.28	2.36	0.40	0.58	0.82
σ_η^2	0.00	0.03	0.09	0.00	0.02	0.09	0.00	0.04	0.10

	Germany			Greece			Italy		
	lower bound	*parameter estimate*	*upper bound*	*lower bound*	*parameter estimate*	*upper bound*	*lower bound*	*parameter estimate*	*upper bound*
φ_1	0.06	0.20	0.35	-0.18	-0.06	0.00	-0.11	0.03	0.14
φ_2	-0.07	0.07	0.22	-0.19	-0.09	0.00	-0.01	0.11	0.24
φ_3	0.08	0.22	0.36	-0.21	-0.12	-0.03	0.00	0.13	0.28
φ_4	-0.14	0.00	0.14	0.30	0.40	0.57	-0.09	0.06	0.16
$\sum_{i=1}^{4}\varphi_i$	0.19	0.50	0.85	-0.09	0.13	0.43	0.13	0.33	0.57
δ	0.09	0.22	0.37	0.12	0.20	0.29	0.08	0.20	0.37
σ_ε^2	0.70	0.94	1.25	1.87	2.19	2.39	1.80	2.05	2.30
σ_η^2	0.00	0.03	0.09	0.00	0.04	0.09	0.00	0.04	0.11

	Netherlands			Portugal			Spain		
	lower bound	*parameter estimate*	*upper bound*	*lower bound*	*parameter estimate*	*upper bound*	*lower bound*	*parameter estimate*	*upper bound*
φ_1	0.01	0.14	0.26	-0.12	-0.01	0.13	0.08	0.22	0.36
φ_2	0.07	0.20	0.34	-0.09	0.04	0.16	-0.10	0.04	0.18
φ_3	0.06	0.19	0.33	-0.14	-0.03	0.10	-0.08	0.07	0.21
φ_4	-0.02	0.11	0.24	0.09	0.19	0.31	0.04	0.18	0.32
$\sum_{i=1}^{4}\varphi_i$	0.36	0.64	0.97	-0.13	0.19	0.51	0.17	0.50	0.90
δ	0.10	0.23	0.39	0.10	0.26	0.40	0.11	0.23	0.37
σ_ε^2	1.44	1.71	1.99	1.62	1.90	2.17	0.88	1.16	1.48
σ_η^2	0.00	0.05	0.12	0.00	0.03	0.10	0.00	0.05	0.12

Note: Data: seasonally adjusted q-o-q change of the GDP deflator.
Time span: 1993:1 - 2006:1, Greece since 1994:4, and Poland since 1995:3.
Results reported were obtained by importance sampling.
90% confidence interval bounds are reported.

Tables 4a and 4b provide a parameter estimate comparison of the extent of inflation persistence in selected NMS and euro area countries. Because of possible aggregation bias, we compare inflation persistence at the level of individual countries. The table suggests that inflation persistence adjusted for the effects of monetary policy in the Czech Republic, Poland and Slovakia belong to the group of euro area countries with lower inflation persistence (Finland, Portugal). For example, the intrinsic and extrinsic inflation persistence in Slovakia is 0.28, while in Belgium the persistence reaches 0.45. On the other hand, 90% confidence intervals often reject statistical difference of the inflation persistence between countries.

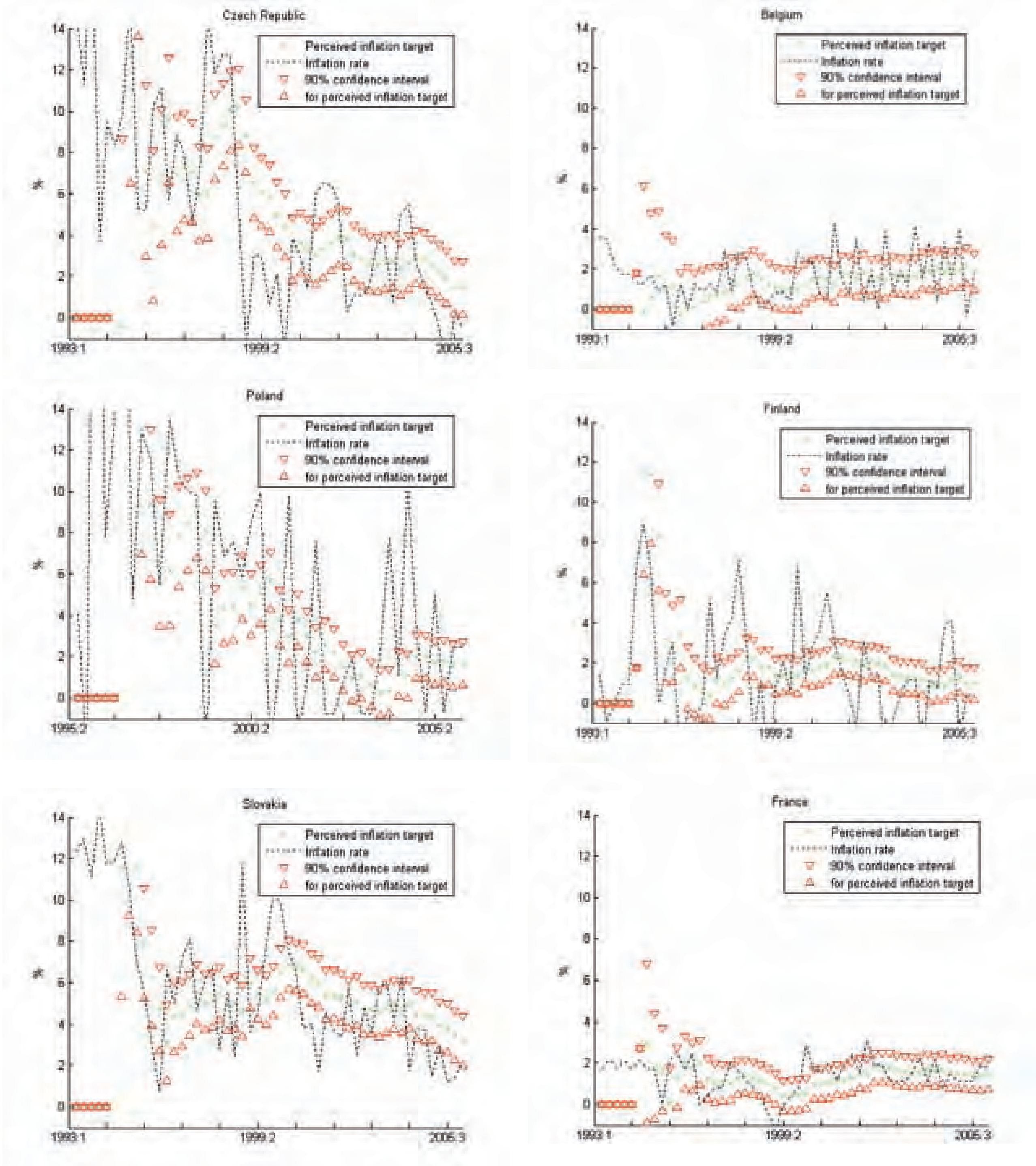
The time-varying mean model enables a discussion of the credibility of monetary authorities and the extent of expectation-based persistence. The values of parameter δ are lower for selected NMS than for selected euro area countries, suggesting that the public in the NMS sets its expectations about inflation rates less in accordance with the modeled targets announced by central banks than in euro area countries (or alternatively, that the fraction of forward-looking members of the public is lower in selected NMS). The conclusion often holds even in terms of 90% confidence intervals.

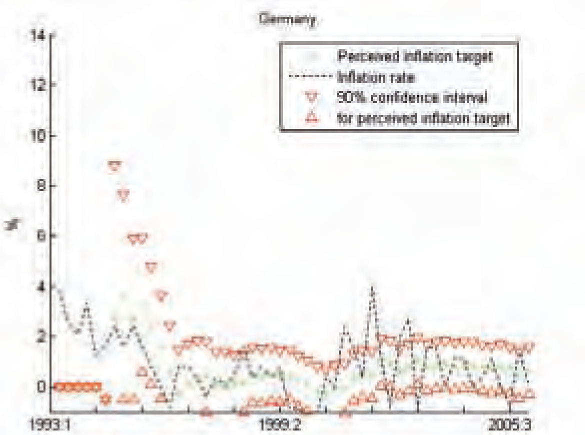
Finally, the Czech Republic, Poland and Slovakia experience higher variance of shocks to the modeled inflation target and also shocks in the inflation equation than the euro area countries. This is a consequence of the transition phase in the 1990s which included cost-push shocks, significant changes in monetary strategies, etc.

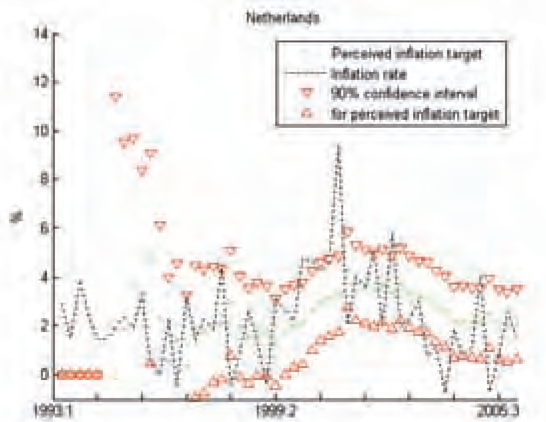
With estimated parameters, it is possible to use the method of the exact initial Kalman filter to estimate the unobservable components of the system. One has to bear in mind that we do not know the exact parameter values, and must work with estimates. However, for our purposes our knowledge of parameter estimates is sufficient.

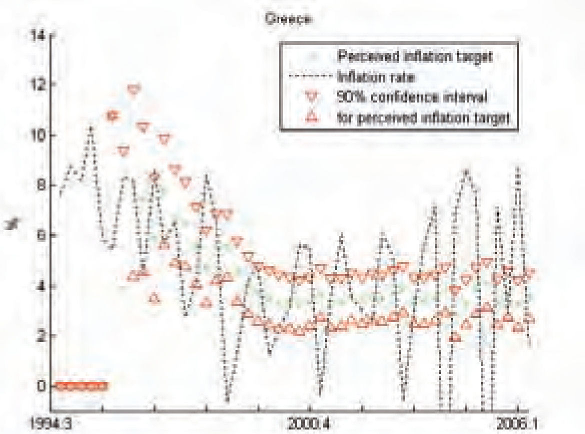
The results of the Kalman filtering are depicted in Figure 2. Note that the inflation target pursued by the central bank is modeled as a random walk and the perceived inflation target that serves as a time-varying mean follows an AR(2) process.

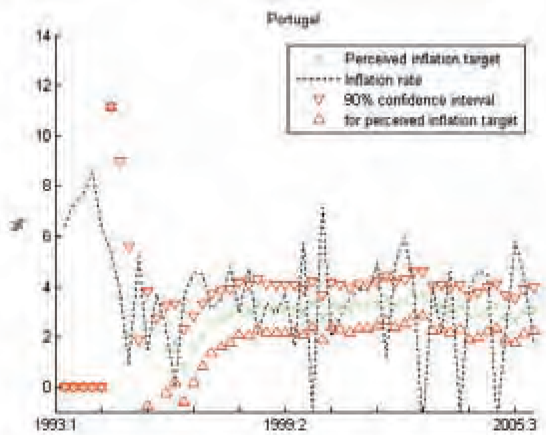
Figure 2: Perceived targets in the time-varying mean models.

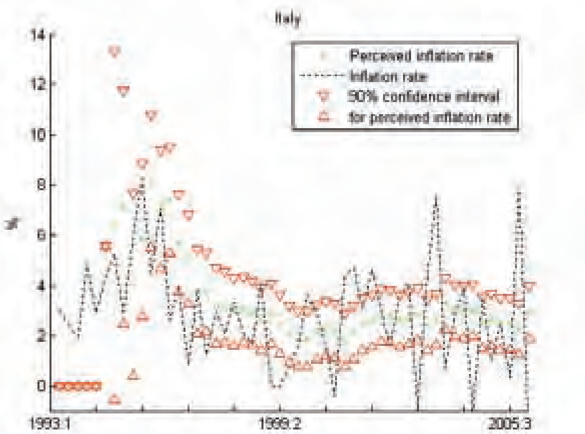


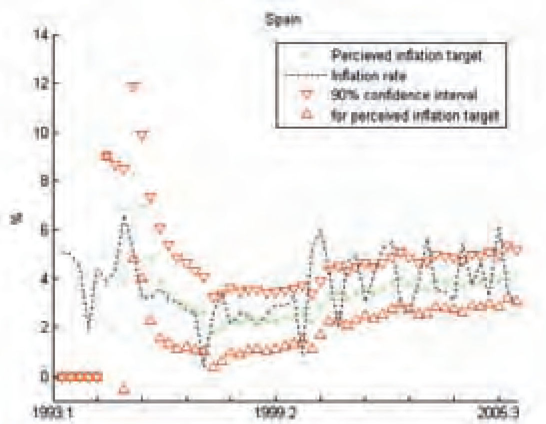

Germany
Perceived inflation target
Inflation rate
90% confidence interval
for perceived inflation target
%
1993:1
1999:2
2005:3


Netherlands
Perceived inflation target
Inflation rate
90% confidence interval
for perceived inflation target
%
1993:1
1999:2
2005:3


Greece
Perceived inflation target
Inflation rate
90% confidence interval
for perceived inflation target
%
1994:3
2000:4
2006:1


Portugal
Perceived inflation target
Inflation rate
90% confidence interval
for perceived inflation target
%
1993:1
1999:2
2005:3


Italy
Perceived inflation rate
Inflation rate
90% confidence interval
for perceived inflation rate
%
1993:1
1999:2
2005:3


Spain
Percieved inflation target
Inflation rate
90% confidence interval
for perceived inflation target
%
1993:1
1999:2
2005:3

First note that 90% confidence intervals for the perceived inflation target time series are zero for the few first quarters, and then larger in comparison to the rest of the time span considered. This is a consequence of the method of exact initial Kalman filter used, which assumes infinite variances for initial values of unobserved components (π_0^P) of the system. Thus we do not report the first few confidence intervals, so as to keep the figures in a reasonable range.

The figures suggest why classical measures of inflation persistence could be inappropriate, especially for the NMS. While the time-varying mean (the perceived inflation target) exhibits breaks for the NMS, no such clear breaks can be observed for the euro area countries.

The figures also capture the effect that the adoption of inflation targeting had on the inflation perceived by the public. For example, in the Czech Republic inflation targeting was adopted in 1997/1998. A year later one may observe that there is a switch in the formation of public perception of inflation can be observed. Since then, the time-varying mean of inflation has been close to the target of 3%.

(iii) The ARFIMA model

At first, we estimate the fractional differencing parameter *d.* We opt for Geweke and Porter-Hudak's technique[15] and report results in the Table 5. Based on estimated value of parameter *d*, we estimate the impulse response function of ARFIMA(0,*d*,0)[16]. To compare the persistence of shocks in the time series, we follow Gadea and Mayoral (2006) and report the values of the impulse response function for selected time horizons (*h=4* and *h=12*) after the realization of a shock.

The results show that Hungary, Poland and Slovakia score high in the persistence suggested by ARFIMA, together with Greece and Spain. The Czech Republic ranks midway in the whole sample of 14 countries.

To assess the relevance of the ARFIMA model in inflation modeling, we test the hypothesis that inflation series follow a fractionally integrated process, against the hypothesis that series follow a stationary process with breaks. The results of the test outlined in subsection 3.1 are reported in Table 5b. In most cases, the fractionally integrated process hypothesis can be rejected at the 1% level. The only inflation process for which we cannot reject the null of a fractionally integrated process at any reasonable significance level is the inflation series for Slovakia.

[15] Implemented in STATA by Baum and Wiggins (1999).
[16] The impulse response function measures the effects of the realization of shock in y_t on subsequent values of time series. See Andrews and Chen (1994) for details. We used STATA implementation for ARFIMA written by Baum (2000).

Table 5: Estimation of fractional differencing parameter *d* and value of impulse response function for selected time horizons

Inflation based on GDP deflator

Country	*d*	SE(*d*)	IPF(4)	IPF(12)
cze	0,59	0,25	0,38	0,24
hun	0,74	0,21	0,56	0,42
pol	0,93	0,15	0,87	0,81
svk	0,90	0,42	0,81	0,73
bel	0,63	0,33	0,42	0,28
esp	0,84	0,32	0,72	0,60
eu12	0,75	0,21	0,58	0,44
fin	0,23	0,63	0,08	0,04
fra	0,19	0,28	0,07	0,03
ger	0,54	0,29	0,32	0,20
grc	1,06	0,14	1,13	1,20
irl	0,22	0,20	0,08	0,03
ita	0,40	0,33	0,20	0,10
nld	0,75	0,24	0,57	0,43
prt	0,52	0,32	0,30	0,18

Table 5b: Test of fractional integration process of order *d* versus stationary process with breaks

	d				
Country	0,5	0,6	0,7	0,8	0,9
cze	0,817	**0,353** **	0,151 *	0,064	0,027
hun	0,604 ***	0,258 ***	**0,110** ***	0,046 **	0,020 *
pol	0,507 ***	0,212 ***	0,088 ***	0,036 ***	**0,015** ***
svk	0,784 *	0,342 **	0,147 **	0,062	**0,026**
bel	0,615 ***	**0,252** ***	0,103 ***	0,042 ***	0,017 **
esp	0,658 ***	0,281 ***	0,119 ***	**0,050** *	0,021 *
eu12	0,659 ***	0,280 ***	0,118 ***	**0,049** **	0,020 *
fin	**0,774** *	0,337 **	0,145 **	0,062	0,026
fra	**0,700** ***	0,308 ***	0,133 **	0,057	0,024
ger	**0,675** ***	0,287 ***	0,121 ***	0,050 *	0,021 *
grc	0,614 ***	0,255 ***	0,105 ***	0,043 **	0,018 **
irl	**0,546** ***	0,223 ***	0,090 ***	0,037 ***	0,015 ***
ita	**0,611** ***	0,254 ***	0,105 ***	0,043 **	0,018 **
nld	0,623 ***	0,261 ***	**0,109** ***	0,045 **	0,018 **
prt	**0,592** ***	0,246 ***	0,102 ***	0,042 ***	0,017 **
1% critical values	0,715	0,335	0,132	0,043	0,016
5% critical values	0,768	0,364	0,147	0,050	0,020
10% critical values	0,797	0,381	0,156	0,054	0,022

Notes: Computation of the test statistics and critical values are based on Mayoral (2004). ***, **, and * denote the significance at 1%, 5% and 10% levels, respectively. For each country, the cell in bold determines the column closest to the value of d estimated using the Geweke and Porter-Hudak technique and reported in Table 5.

4.2 Structural measures

The estimation of the NHPC is significantly influenced by the data availability, especially for Slovakia. Some time series are only available for a part of the time span considered. Moreover, some time series are available only annually. Therefore, we compromise between the data availability and the possibility to carry out the analysis, and use yearly instead of quarterly data for some instruments. Data used are described in Appendix 1.

The hybrid version of the Phillips curve incorporates the pure forward-looking version as a special case. Hence, we can only estimate the hybrid version. We also purely deal with the closed version of the model, since the instrument set employed performs poorly for the open economy version of the NHPC. Finally, as it is usual in the related literature, we assume rational expectations. Future actual inflation rate, therefore, stands for expected inflation rate in estimation of NHPC.

Zhang et al. (2006) point out the influence of the instrument set on the estimation results, especially when autocorrelation of residuals is present. We employ the sets of instruments introduced in Galí and Gertler (1999), Galí, Gertler and López-Salido (2001), Zhang et al. (2006) and Lendvai (2005). We also add some instruments that we think of as valid for the estimation in the case of the NMS. Table 6 below reports estimation results for the Czech Republic, Poland and Slovakia for various sets of instruments. Estimates for Hungary are available, as already mentioned in Part 2. Staiger and Stock (1997) suggest a rule of thumb for the relevance of instruments: the F-statistics of the overall relevance of excluded instruments should exceed 10. F-statistics below 10 imply a bias in the estimated coefficients. We therefore do not report estimation results for sets of instruments that are not relevant according to this criterion.

TABLE 6: New Hybrid Phillips Curve: Estimation for Various Sets of Instruments - Czech Republic, Poland and Slovakia

Excluded instruments (lags) see Appendix 1	Czech Republic						Poland						Slovakia					
	GG	GGL	ZO	L	IS1	IS2	GG	GGL	ZO	L	IS1	IS2	GG	GGL	ZO	L	IS1	IS2
infl_d	2,3,4	2,3,4,5	x	2	2,3	x	2,3,4	2,3,4,5	x	2	2,3	x	2,3,4	2,3,4,5	x	2	2,3	x
lrulc_d	2,3,4	1,2	x	1,2	2,3	x	2,3,4	1,2	x	1,2	2,3	x	2,3,4	1,2	x	1,2	2,3	x
irspread	1,2,3,4	x	x	x	1,2,3,4	x	1,2,3,4	x	x	x	1,2,3,4	x	1,2,3,4	x	x	x	1,2,3,4	x
ogap	1,2,3,4	x	x	x	1,2	x	x	x	x	x	x	x						
deficit	x	x	x	1,2	1,2	x	x	x	x	1,2	1,2	x	x	x	x	1,2	1,2	x
diff_rer_d	x	x	x	1,2	0,1,2	x	x	x	x	1,2	0,1,2	x	x	x	x	1,2	0,1,2	x
rer_d	x	x	x	0	x	x	x	x	x	0	x	x	x	x	x	0	x	x
u_rate	x	x	1,2,3,4	x	x	1,2	x	x	1,2,3,4	x	x	1,2	x	x	1,2,3,4	x	x	1,2
diff_treasury	x	x	1,2,3,4	x	x	1,2,3,4	x	x	1,2,3,4	x	x	1,2,3,4						
output_d	x	1,2	x	x	x	1,2,3,4	x	1,2	x	x	x	1,2,3,4	x	1,2	x	x	x	1,2,3,4
rg_exp_d	x	x	1,2,3,4	x	x	x	x	x	1,2,3,4	x	x	x						
winfl_d	1,2,3,4	1,2	x	1,2	1,2,3,4	x	x	x	x	x	x	x						
cap_ut	x	x	1,2,3,4	x	x	1,2,3,4	x	x	x	x	x	x						
diff_1day	x	x	1,2,3,4	x	x	1,2,3,4	x	x	1,2,3,4	x	x	1,2,3,4	x	x	1,2,3,4	x	x	1,2,3,4
Results																		
$\hat{\gamma}_f$	0.47*		0.45*		0.42*	0.42*		-0.66**	-0.35			0.02					0.18	0.34
	(0.14)		(0.12)		(0.14)	(0.12)		(0.28)	(0.23)			(0.27)					(0.36)	(0.24)
$\hat{\gamma}_b$	0.35*		0.38*		0.35*	0.38*		-0.31**	-0.26***			-0.19					0.14	-0.02
	(0.08)		(0.09)		(0.07)	(0.09)		(0.15)	(0.13)			(0.15)					(0.18)	(0.15)
$\hat{\lambda}$	-0.09		-0.06		-0.06	-0.04		-0.05	-0.05			-0.05					0.15	0.13
	(0.18)		(0.16)		(0.16)	(0.15)		(0.07)	(0.06)			(0.05)					(0.12)	(0.11)
Instrument relevance																		
F statistics	10.81	1.09	19.51	1.06	72	35.26	3.12	10.46	13.84	2.59	1.84	30.9	8.75	3.64	7.64	1.8	15.41	11.76
Partial R2	0.56	0.19	0.59	0.28	0.57	0.58	0.41	0.34	0.41	0.23	0.08	0.53	0.45	0.14	0.33	0.12	0.49	0.46

* 1% significance level ** 5% ***10%

3-lag HAC-robust standard errors are reported in parentheses.

Estimation results for relevant (F statistics above 10) sets of instruments are reported.

We employ instrument sets that replicate Galí and Gertler (1999) **GG**, Galí, Gertler and López-Salido (2001) **GGL**, Zhang et al. (2006) **ZO**, and Lendvai (2005) **L**. We also add some instruments that we consider as valid for the estimation: **IS1** and **IS2**.

For a definition of these instruments, see Appendix 1. The suffix _d denotes HP filtered time series.

In Table 6, the numbers in upper panels report the lags of variables that are included in various sets of instruments. The panels in the middle of the table provide estimates of reduced form coefficients. Finally, the lower panels report F-statistics and partial R^2.

For sets of instruments resulting in F-statistics above 10, we carry out a Hansen J test for overidentifying restrictions. In all cases we cannot reject the null of satisfied overidentifying restrictions at all relevant significance levels. Furthermore, we test for homoskedasticity employing the Pagan-Hall test and for residual autocorrelation using the Breusch-Godfrey test. We detect serially correlated residuals in all cases and we reject homoskedasticity for Poland.[17] Based on the results of diagnostics test mentioned, we correct for serial correlation and heteroskedasticity using three-lag HAC-robust standard errors.

Overall, the estimation results suggest that structural NHPC is not an appropriate short-run inflation dynamics model for Poland and Slovakia. The estimated coefficients for these countries are not significant and often have a sign that does not correspond to the underlying theory. On the other hand, for the Czech Republic the estimated reduced form coefficients $\hat{\gamma}_f, \hat{\gamma}_b$ of the model are significant with the expected sign and within the range predicted by the micro theory. However, the slope parameter at the real marginal cost term $\hat{\lambda}$ is not statistically significant.

We focus on the comparison of reduced form coefficients γ_b, γ_f since we are purely interested in the extent of inflation inertia.[18] Detailed analysis of the structural Phillips curve estimation lies beyond the scope of this current study. The comparison suggests that the predominance of the expected future inflation over past inflation in the euro area (and the US) is not detected for the Czech Republic and Hungary.[19] If we follow the definition of (intrinsic) inflation persistence from previous sections, we can conclude that the Czech Republic and Hungary exhibit comparable or higher inflation persistence than euro area countries. Moreover, lower predominance of the forward-looking term is in accordance with the results of statistical measures based on the autoregressive model with time-varying mean from the preceding subsection.

[17] For discussion on the possible sources of residual autocorrelation, see Galí, Gertler and López-Salido (2001).
[18] Note that the reduced form coefficients are a sole function of deep parameters.
[19] See results for the US, euro area and Hungary in Galí and Gertler (1999), Galí, Gertler and López-Salido (2001), Zhang et al. (2005), and Lendvai (2005). We summarize the results of interest in the Section 6.

5. Conclusions

According to our analysis, inflation persistence in the NMS is comparable to that in the current euro area Member States. Table 7 illustrates that the NMS are characterized by lower intrinsic and extrinsic (both estimated together) inflation persistence under the time-varying mean assumption, placing them among the euro area countries. On the other hand, the four NMS score among the top five when all countries are ordered according to the scope of inflation persistence estimated under the constant mean assumption.

We argue that time-varying mean models should be a preferred option for inflation persistence measurement in the NMS. According to our tests, the models with breaks in parameters are superior to the ARFIMA models for most considered countries, while the constant mean assumption is too restrictive for estimating the inflation persistence in the NMS for two reasons. First, means are estimated from data samples covering both transformation as well as convergence process. They are much closer to the notion of a medium-run equilibrium that gradually moves towards the long-run equilibrium. However, this is only possible in models with time-varying parameters. Second, changes in expectations and monetary policy regimes are likely to contribute to changes in inflation targets that are closely related to means estimated from data.

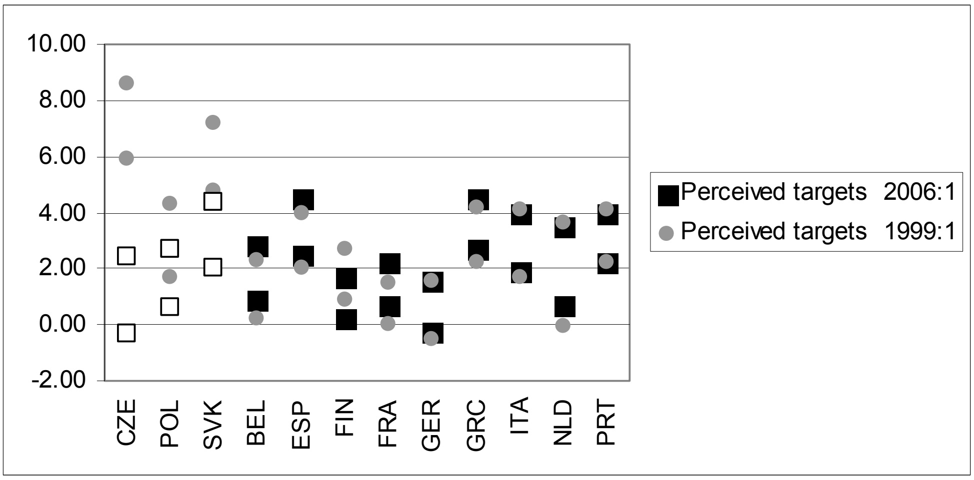
Table 7: Summary of results – statistical measures

	ρ_K (OLS)		ρ_K (Hansen) lag =5		$\Sigma\varphi_i$ (Time-varying mean)	
CZE	0.75 - 0.8	(1)	0.88	(2)	0.26	(8)
HUN	0.75	(2-3)	0.98	(1)	x	
POL	0.68	(4)	0.78	(4)	0.12	(11)
SVK	0.75	(2-3)	0.74	(5)	0.28	(7)
EU12	0.66	(x)	0.70	(x)	x	
BEL	0.13	(12)	0.28	(11)	0.45	(4)
ESP	0.26 - 0.6	(10)	0.69	(7)	0.50	(2-3)
FIN	0.33	(9)	0.24	(12)	0.07	(12)
FRA	0.43	(8)	0.57	(9)	0.37	(5)
GER	0.50 - 0.6	(7)	0.59	(8)	0.50	(2-3)
GRC	0.67	(5)	0.82	(3)	0.13	(10)
IRL	0.11	(13)	0.10	(13)	x	
ITA	0.14	(11)	0.37	(10)	0.33	(6)
NLD	0.62	(6)	0.71	(6)	0.64	(1)
PRT	-0.16	(14)	-0.03	(14)	0.19	(9)

Note: For each approach, we report parameter estimates. Intervals indicate estimates by various methods as presented in the paper. In brackets, countries are ordered according to the scope of estimated inflation persistence.

Changes in expectations and monetary policy regimes are key to analyzing inflation persistence in the NMS. Figure 3 shows how important it is to account for changes in expectations and monetary policy regimes when intrinsic and extrinsic inflation persistence is being estimated. The perception of inflation has changed profoundly in the NMS. This is in contrast with remarkably stable perceived inflation in the current euro area Member States. Recently, perceived inflation in the NMS has been similar to that of the current euro area members. This finding supports the hypothesis that the medium-run equilibrium, which is important for estimates of inflation persistence in the NMS, converges towards the long-run equilibrium that defines the theoretical concept of inflation persistence. However, the NMS should not take it for granted that values of perceived targets can only decrease over time. The examples of some current euro area countries show that upward changes are also possible.

Figure 3: Perceived inflation targets from 1999 to 2006



Note: Targets are represented with confidence intervals (vertical axis is in %).

The importance of expectations is further emphasized by the estimates of structural measures of persistence. To previously published results for Hungary, we add our estimates of the NHPC for the Czech Republic. For these two NMS, past inflation is relatively more important than in the current euro area Member States for whom the expected future inflation dominates. This result might indicate that although inflation persistence in the NMS is comparable to that in the current euro area Member States, it does not have the same roots. The reduced importance of the forward looking term for the NMS, together with slightly lower estimated credibility of inflation targets (parameter δ in the time varying mean models), indicates that anchoring expectations may be very important for the NMS during the euro adoption process.

Table 8: Estimates of New Hybrid Phillips Curve in various studies

Coefficient	Czech Republic Summary	Hungary L (2005)	euro area Summary	GG (1999) a	GG (1999) b	GGL (2001) a	GGL (2001) b	ZO (2005)
$\hat{\gamma}_b$	0.42-0.47	0.467	0.04-0.59	0.252	0.378	0.043	0.272	0.587
		(0.084)		(0.023)	(0.020)	(0.115)	(0.072)	(0.085)
$\hat{\gamma}_f$	0.35-0.38	0.553	0.43-0.77	0.682	0.591	0.773	0.689	0.429
		(0.084)		(0.020)	(0.016)	(0.064)	(0.047)	(0.089)

Note: See Galí and Gertler (1999) – Table 2, Galí, Gertler and López-Salido (2001) – Table 2, Zhang et al. (2005) – Table 2, and Lendvai (2005) – Table 3a. The two versions of Galí and Gertler (1999) and Galí, Gertler and López-Salido (2001) correspond to the two versions of orthogonality conditions. For the euro area GDP deflator is used, Lendvai (2005) uses core inflation. Standard errors are reported in parentheses.

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Appendix 1: Data description

Inflation based on the GDP deflator: seasonally adjusted annualized q-o-q rate of change of GDP deflator published in the OECD OEO Database. Time span: 1993:2-2006:1. The exception is Hungary, for which we used time span 1995:1-2006:1. Transformation: $inf_{GDP} = 400ln(GDPdefl/GDPdefl_{-1})$.

Inflation based on non-food non-energy CPI: annualized q-o-q rate of change of Consumer Price Index published in the OECD MEI Database. Timespan: 1996:2-2006:2. Transformation: $inf_{coreCPI} = 400ln(coreCPI/coreCPI_{-1})$.

Country name abbreviations: BEL (Belgium), CZE (Czech Republic), FIN (Finland), FRA (France), GER (Germany), GRC (Greece), HUN (Hungary), IRL (Ireland), ITA (Italy), NLD (Netherlands), POL (Poland), PRT (Portugal), SVK (Slovakia), ESP (Spain)

The structural Phillips curve is estimated based on the quarterly data covering the period 1993:2-2006:1 for the Czech Republic and Poland, and 1995:2-2006:1 for Slovakia.[20] We take over the inflation based on the GDP deflator (*infl*) and the real effective exchange rate (*reer*) from the preceding analysis. Real marginal costs are represented by the log of real unit labor cost deflated by the GDP deflator (*lrulc*). In addition, we employ the following series:

ogap: output gap is a percentage of total GDP
irspread: the difference between short-term (1 day) and long-term (3 months) interest rate
deficit: government surplus or deficit in terms of GDP
rer: real exchange rate
diff_rer: q-o-q change of real exchange rate
u_rate: unemployment rate
diff_treasury: first difference of long-term interest rate (10 years)
output: GDP
rg_exp: real government expenditure (deflated by the GDP deflator)
winfl: wage inflation (annualized q-o-q change)
cap_ut: capacity utilization
diff_1day: first difference of short-term interest rate.

The output gap is available for the Czech Republic and for the period 1995:1-2006:1 only; we impose zeros for the period 1993:1-1994:4. Government deficit is available annually since 1995:1 (we impose zeros for the period 1993:1-1994:4). GDP is available quarterly since 1996 for the Czech Republic and since 1995 for Poland. Only annual values for the long-term interest rate (10 years) are available for Slovakia, and therefore we do not include *diff_treasury* in the analysis for that country. Quarterly values of the long-term interest rate (10 years) are available since 1999:2 for Poland. Government expenditures are available since 1996 for Poland; we impose values as of 1996 in the period before. For Slovakia, government expenditures are not available. For the Czech Republic, wage inflation is available quarterly since 1998; annual values are available in the preceding period. Time series of wage inflation and capacity utilization are not available for Slovakia and Poland.

[20] Data were downloaded from the ECB Statistical Data Warehouse, Data sources are: OECD Economic Outlook, OECD Main Economic Indicators, ECB Euro Area Accounts and Economic Statistics – Government Statistics and ESA, and the Czech Statistical Office.

Appendix 2: Sum of autoregressive coefficients – core inflation

Table: OLS estimates of ρ_K
(inflation based on non-food non-energy CPI)

	Preferred model according to AIC		Preferred model according to BIC	
	Lag length	Sum of AR coefficients	Lag length	Sum of AR coefficients
CZE	4	0.75	2	0.65
HUN	4	0.85	4	0.85
POL	4	0.84	4	0.84
SVK	1	0.21	1	0.21
BEL	5	-0.14	5	-0.14
ESP	3	-1.56	1	-0.95
FIN	5	0.65	5	0.65
FRA	4	0.75	4	0.75
GER	4	0.20	4	0.20
GRC	4	0.51	4	0.51
IRL	5	0.49	4	0.57
ITA	4	0.33	2	-0.05
NLD	5	0.67	4	0.85
PRT	4	0.72	4	0.72

Figure 2: Inflation based on non-food non-energy CPI, ρ estimate and its 90% confidence intervals
(lag length = 5, Hansen's (1999) grid bootstrap procedure)

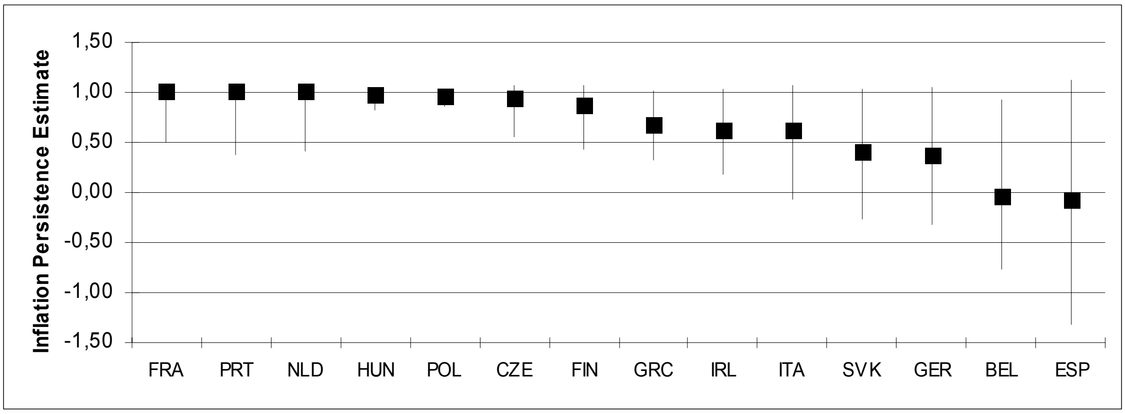
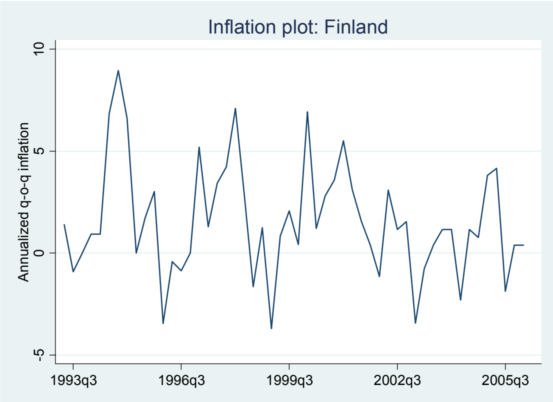
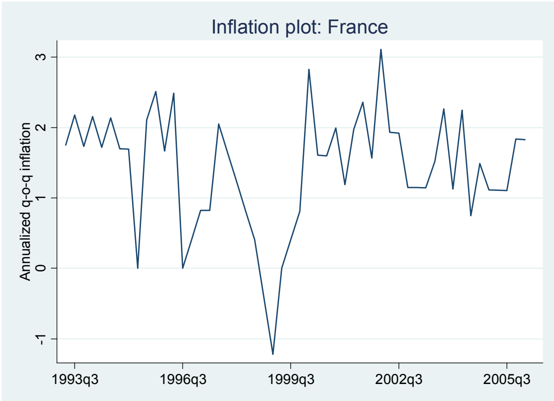


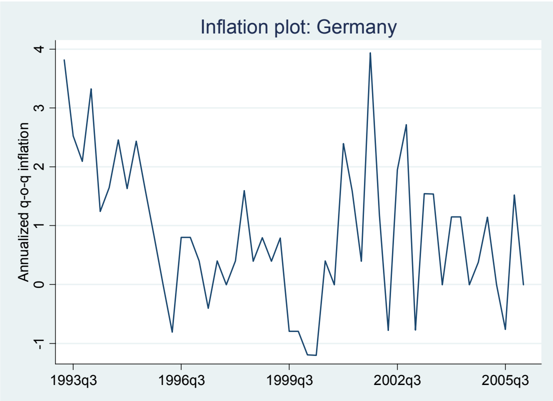
Table: ρ_K its 90% confidence intervals estimated using Hansen's (1999) grid bootstrap procedure
(inflation based on non-food non-energy CPI)

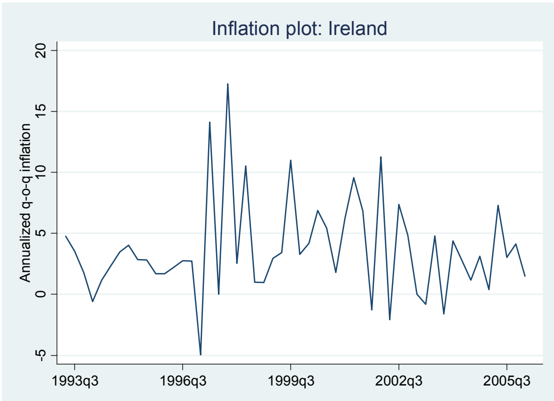
	Lag length = 5			Lag length = 4			Lag length = 3			Lag length = 2			Lag length = 1		
	lower bound	mean	upper bound	lower bound	mean	upper bound	lower bound	mean	upper bound	lower bound	mean	upper bound	lower bound	mean	upper bound
CZE	0.55	**0.95**	1.07	0.60	**1.00**	1.09	0.44	**0.80**	1.04	0.50	**0.81**	1.04	-0.33	**-0.06**	0.17
HUN	0.83	**0.98**	1.04	0.81	**0.95**	1.02	0.73	**0.99**	1.06	0.26	**0.53**	0.83	0.33	**0.55**	0.79
POL	0.85	**0.97**	1.03	0.81	**0.92**	0.99	0.79	**0.99**	1.05	0.76	**0.96**	1.04	0.71	**0.86**	1.02
SVK	-0.26	**0.41**	1.04	0.02	**0.54**	1.05	-0.19	**0.28**	0.79	-0.03	**0.33**	0.75	-0.02	**0.24**	0.53
BEL	-0.77	**-0.03**	0.93	-0.49	**0.28**	1.06	-0.82	**-0.20**	0.46	-0.40	**0.08**	0.59	-0.63	**-0.38**	-0.14
ESP	-1.32	**-0.08**	1.13	-2.11	**-1.03**	0.10	-2.23	**-1.51**	-0.78	-1.31	**-0.76**	-0.18	-1.06	**-0.97**	-0.85
FIN	0.43	**0.87**	1.07	0.39	**0.85**	1.07	0.21	**0.59**	1.03	0.36	**0.76**	1.04	-0.52	**-0.23**	0.04
FRA	0.50	**1.02**	1.09	0.59	**1.03**	1.12	0.12	**0.56**	1.03	0.13	**0.48**	0.95	-0.14	**0.13**	0.41
GER	-0.32	**0.38**	1.06	-0.20	**0.39**	1.04	-0.52	**0.07**	0.65	-0.24	**0.23**	0.71	-0.60	**-0.37**	-0.10
GRC	0.32	**0.67**	1.02	0.32	**0.63**	0.96	0.09	**0.41**	0.80	0.33	**0.72**	1.05	-1.00	**-0.86**	-0.71
IRL	0.17	**0.62**	1.04	0.33	**0.78**	1.06	0.11	**0.51**	1.02	0.23	**0.60**	1.02	-0.16	**0.11**	0.38
ITA	-0.07	**0.62**	1.08	-0.09	**0.56**	1.06	-0.28	**0.17**	0.69	-0.35	**0.00**	0.33	-0.05	**0.19**	0.47
NLD	0.41	**1.01**	1.09	0.71	**1.04**	1.27	-0.14	**0.40**	1.03	0.12	**0.56**	1.03	-0.50	**-0.22**	0.04
PRT	0.37	**1.02**	1.09	0.46	**1.02**	1.14	-0.07	**0.44**	1.03	0.09	**0.55**	1.03	-0.89	**-0.70**	-0.50

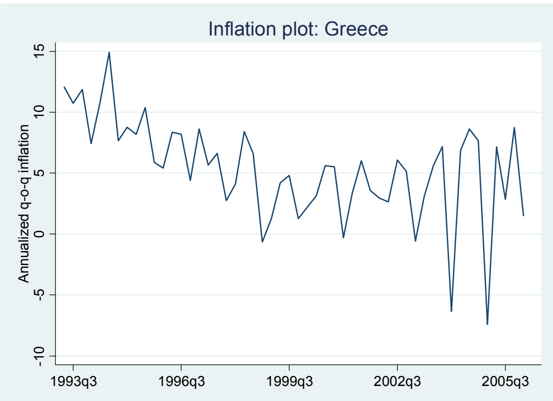
Appendix 3: Inflation plots for selected countries.

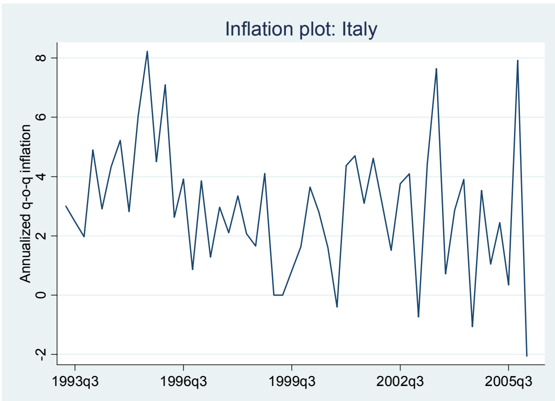

Inflation plot: Finland
Annualized q-o-q inflation
-5
0
5
10
1993q3
1996q3
1999q3
2002q3
2005q3

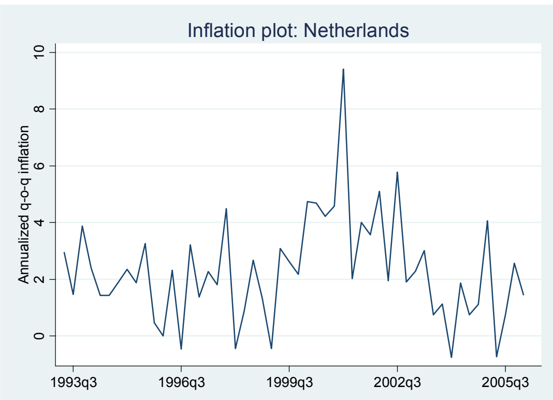

Inflation plot: France
Annualized q-o-q inflation
-1
0
1
2
3
1993q3
1996q3
1999q3
2002q3
2005q3

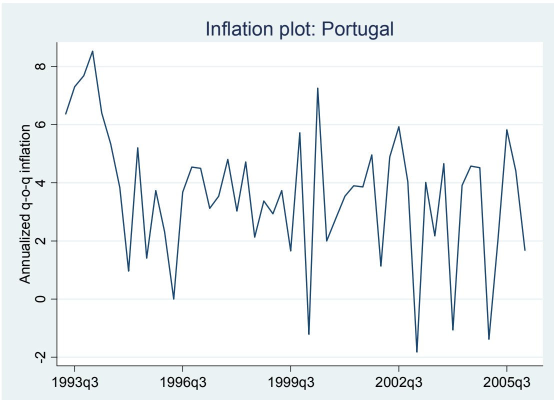

Inflation plot: Germany
Annualized q-o-q inflation
-1
0
1
2
3
4
1993q3
1996q3
1999q3
2002q3
2005q3


Inflation plot: Ireland
Annualized q-o-q inflation
-5
0
5
10
15
20
1993q3
1996q3
1999q3
2002q3
2005q3


Inflation plot: Greece
Annualized q-o-q inflation
-10
-5
0
5
10
15
1993q3
1996q3
1999q3
2002q3
2005q3


Inflation plot: Italy
Annualized q-o-q inflation
-2
0
2
4
6
8
1993q3
1996q3
1999q3
2002q3
2005q3


Inflation plot: Netherlands
Annualized q-o-q inflation
0
2
4
6
8
10
1993q3
1996q3
1999q3
2002q3
2005q3


Inflation plot: Portugal
Annualized q-o-q inflation
-2
0
2
4
6
8
1993q3
1996q3
1999q3
2002q3
2005q3

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