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Changes in inflation dynamics under inflation targeting? Evidence from Central European countries☆

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ABSTRACT

Many countries have implemented inflation targeting in recent decades. At the same time, the international conditions have been favorable, so it is hard to assess to what extent the success in stabilizing inflation should be attributed to good luck and to what extent to the specific policy framework. In this paper, we provide a novel look at the dynamics of inflation under inflation targeting, focusing on three Central European (CE) countries that adopted the IT regime at similar times and in similar environments. We use the framework of the open economy New Keynesian Phillips curve (NKPC) with time-varying parameters and stochastic volatility to recover changes in price-setting and expectation formation behavior and volatility of shocks. We employ Bayesian model averaging to tackle the uncertainty in the selection of instrumental variables and to account for the possible country-specific nature of inflation dynamics. The results suggest that inflation targeting does not itself automatically trigger changes in the inflation process, and the way the framework is implemented might matter. In particular, we find rather heterogeneous evolution of intrinsic inflation persistence and volatility of inflation shocks across these countries despite the fact that all three formally introduced inflation targeting a decade ago.

1. Introduction

Understanding the nature of short-term inflation dynamics poses a major challenge for monetary policy. Sound knowledge of inflation properties is especially pressing for countries whose economies have undergone dramatic structural changes and where the institutional settings of monetary policy have been considerably changed in order to engineer a sharp disinflation process. Taming inflation has traditionally been considered costly in terms of output loss, but a better notion of the role of expectations has given policy makers hope that credible monetary policy can achieve disinflation without having a detrimental effect on real economic activity. This concept has become a hallmark of the New Keynesian Phillips curve (NKPC).

The NKPC was proposed as a structural model of inflation dynamics which is based on an optimization process at the micro-level and thus should be invariant to policy changes. However, this claim is not fully supported by recent research. There are numerous reasons why the parameters of the NKPC model can evolve over time. Importantly, a more aggressive monetary policy stance (Davig and Doh, 2008) and the implementation of credible monetary policy regimes such as inflation targeting (Benati, 2008) have been considered key drivers in reducing inflation persistence and volatility by anchoring inflation expectations. On the contrary, cross-country panel studies such as Ball and Sheridan (2004), Mishkin and Schmidt-Hebbel (2007), and Brito and Bystedt (2010) find rather mixed evidence on the relative performance of inflation targeters vs. non-targeters in both developed and emerging countries.

The countries of Central Europe (CE) represent a unique sample for analyzing changes in inflation dynamics related to the adoption of inflation targeting and overall changes in economic conditions: the three CE countries – the Czech Republic, Hungary, and Poland – are relatively similar small open economies with a strong regional and historical affinity. They jointly underwent a transition to a market economy, which could have induced similar changes in both price-setting and expectation formation behavior. Finally, they all introduced inflation targeting as a disinflation strategy (the Czech Republic in 1998, Poland in 1999, and Hungary in 2001). On the other hand, their actual monetary policy conduct has shown notable differences, in particular in the role given to

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the exchange rate. Whereas the Czech Republic and Poland have left their currencies to float freely most of the time since launching IT and have used a few time-limited exchange rate interventions (the Czech Republic in 2002 and Poland in 2011), in Hungary the IT framework was accompanied by a target zone for the exchange rate of the forint vis-à-vis the euro. This was lifted only after several years (in 2008) and a currency crisis immediately ensued.

Under these conditions, we can run a natural experiment to assess the impact of inflation targeting and the role of country-specific modifications to the IT framework. In particular, it might be of crucial importance to evaluate the effectiveness of the specific IT implementation in each country vis-à-vis changes in the inflation process such as inflation persistence, the role of inflation expectations, and the volatility of inflation shocks. The lesson learned from this analysis may shed some light on the nature of the differences in the relative performance of inflation targeters, since inflation targeting is still the preferred monetary framework being adopted by emerging countries around the globe. If inflation dynamics were homogeneous across countries, the role of domestic policy and specific issues related to the implementation of IT would be of only minor importance. On the contrary, if one observed persistent differences in inflation dynamics despite a (formally) common monetary policy regime and common foreign shocks, this would be indirect proof that good policy still matters notwithstanding the prominent role of global factors in today's world. It should be also stressed that in contrast to many other emerging and transition economies, the CE countries' membership of the OECD and EU makes the data reliable and internationally comparable.

In this paper, we provide evidence on the evolution of inflation dynamics in the CE countries based on estimates of the New Keynesian Phillips curve. This is augmented by a number of features to suit our purposes. First, we extend the open-economy version of the NKPC proposed in Galí and Monacelli (2005) to a hybrid form and a time-varying context. Second, in relation to time-varying estimation of the NKPC we provide several methodological contributions. In our two-step procedure closely related to Kim (2006), we propose to use Bayesian model averaging (BMA) to tackle the thorny issue of instrument selection in the first step. Thanks to BMA, the instruments are allowed being country specific, reflecting, for example, differences in the expected role of foreign factors in inflation expectations. Indeed, the sensitivity of the results to the choice of the conditioning instrument set has been shown to be very relevant in forward-looking models, with the NKPC being a prominent example (see e.g. (Mavroeidis, 2005)). Third, we add a stochastic volatility model for error terms into the time-varying regression because changes in inflation volatility might induce spurious variation in the estimated coefficients, as previously documented (see e.g. Koop and Korobilis, 2009). Moreover, modeling the magnitude of inflation shocks in a time-varying manner has additional analytical merits. Since a decline in inflation volatility (along the level of inflation) is often seen as the main purpose of the IT framework and can be directly linked to the effectiveness of the policy regime, obtaining relevant information on its evolution might be highly important to policy makers.

The results can be summarized as follows. First and foremost, inflation dynamics are heterogeneous across the three CE countries despite the fact that all three national central banks pursue inflation targeting. Intrinsic inflation persistence has dropped substantially only in the Czech Republic, to currently insignificant levels. More importantly, the volatility of inflation shocks decreased quickly a few years after the adoption of inflation targeting in both the Czech Republic and Poland. By contrast, the nature of the inflation process in Hungary does not seem to have changed much over the last 15 years. Second, the inflation–output trade-off seems to be blurred by potentially important supply shocks during the transition process, and those shocks cannot be fully captured by the NKPC model analyzed. However, the results tend to reveal a non-linear relationship between domestic economic activity and inflation. Indeed, the coefficient on the output gap increases consistently only in periods when output significantly deviates from its potential. Third, similar conclusions can be drawn for foreign inflation factors, tracked by the terms of trade. They turn significant in specific periods such as major exchange rate devaluations. Yet there is also some indication that foreign factors might already be well reflected in inflation expectations themselves. Finally, the overall changes in the inflation process reflect changes in the price-setting behavior of firms, which we detect in both the cross-country and temporal dimensions.

Our empirical findings have some noteworthy policy implications for (emerging) countries that adopt inflation targeting. Although all three CE countries officially adopted inflation targeting more than a decade ago, in two of them (Hungary and Poland) intrinsic inflation persistence has not decreased considerably and remains at a higher level than that reported for developed countries. The volatility of inflation shocks decreased quickly after the introduction of IT in the Czech Republic and Poland, but was again practically unaltered in Hungary. Therefore, it seems that the adoption of inflation targeting does not itself automatically produce any changes in the inflation process, and the particular way the framework is implemented might matter. If inflation targeting is not sufficiently credible, economic agents might mainly take into account observed inflation levels rather than the inflation target. This may arguably be related to the role of the exchange rate in monetary policy. Although inflation targeting is aimed at the domestic price level, Hungarian monetary policy has paid special attention to the exchange rate, and the exchange rate channel is considered the most efficient channel of monetary policy transmission. Our results suggest that this policy choice has itself had its costs, as the goal of stabilizing inflation expectations has not been fulfilled.

The paper is organized as follows. In Section 2, we review relevant literature. Section 3 presents our version of the open economy NKPC that is subject to empirical investigation. Section 4 describes our econometric framework and data. All the results and their interpretation appear in Section 5. The final section concludes.

2. Related literature

From the empirical perspective, the NKPC owes its growing popularity to the seminal papers of Galí and Gertler (1999) (GG hereafter) and Galí, Gertler, and López-Salido (2001, GGL). Despite the theoretical appeal of the (hybrid) NKPC, subsequent studies have produced rather conflicting empirical evidence, with the results varying across economies, data sets, and – most notably – estimation methods (e.g. Rudd and Whelan, 2005; Mavroeidis, 2005; Lindé, 2005). To account for the characteristics of small open economies Galí and Monacelli (2005) derive a small open economy version of the NKPC for CPI inflation, which in addition to marginal cost includes the terms of trade. Mihailov et al. (2011b) provide some favorable empirical evidence on this model based on data for selected OECD countries. With respect to theoretical underpinnings, it is probably the closest empirical study to our own. Alternative approaches include Batini et al. (2005), who propose an open-economy NKPC where the marginal cost is affected by import prices and external competition and conclude that this model fits the UK data well. Rumler (2007) extends the marginal cost measure to include the cost of intermediate inputs (both domestic and imported) and finds some plausible evidence for the euro area countries.

A few recent studies consider the effects of structural changes in the economic system and monetary policy regime and explore how these are propagated into the parameters of the NKPC. Most of the evidence is available for the US. The intrinsic inflation persistence was found to be an empirical artifact driven by specification bias inherent to fixed-coefficient models (Hall et al., 2009) or to variation in the long-run inflation trend (Cogley and Sbordone, 2008). Several authors also document that the nature of the inflation process changes with the macroeconomic environment (Zhang et al., 2008) and the monetary policy regime (Cogley et al., 2010; Kang et al., 2009). However, there are also studies (Stock and Watson, 2007) claiming that inflation persistence has not changed for decades in the US.

The evidence on changes in inflation dynamics in other economies is less abundant.[1] Benati (2008) examines data for several developed inflation-targeting countries (Canada, New Zealand, Sweden, Switzerland, and the UK) and the euro area, concluding that inflation persistence decreased almost to zero once credible monetary regimes had been implemented and, therefore, that inflation persistence is not structural. Hondroyiannis et al. (2009) apply a specific time-varying framework to data for France, Germany, Italy, and the UK, concluding consistently with previous evidence for the US (Hall et al., 2009) that the backward-looking parameter of the time-varying NKPC is almost negligible. Tillmann (2009) explores how the explanatory power of the forward-looking NKPC in the euro area evolves over time, finding that the explanatory power of the model varies substantially across the underlying monetary regimes and is influenced by events such as the ERM crisis, the Maastricht treaty, and the launch of EMU. Koop and Onorante (2011) use dynamic model averaging (Raftery et al., 2010) to study the relationship between inflation and inflation expectations in the euro area. They find strong support for forward-looking behavior, interestingly mainly since the start of the recent financial crisis.

Not much attention has been paid to changes in inflation dynamics in transition countries so far. The existing time-invariant estimates of the NKPC for the countries in question mostly imply that inflation is more persistent than in advanced economies and that external factors seem to matter (Franta et al., 2007; Mihailov et al., 2011a; Vašíček, 2011). Based on this evidence, however, it is very difficult to draw any conclusion about the effects of monetary policy on the temporal and cross-country variation in the inflation process. Only Hondroyiannis et al. (2008) provide some evidence based on a time-varying model for a panel of seven new EU member states. Panel estimation does not seem to be an appropriate technique for the (in our view) highly heterogeneous group of seven new members given that the economic structures and monetary policy frameworks of these countries are very different. The authors find that inflation persistence in these countries is practically nonexistent, which contradicts practically all the country-specific time-invariant evidence.

Our study is conceptually related to papers tracking structural changes in the inflation process by means of time-varying estimation of the NKPC (e.g. Cogley et al., 2010; Hall et al., 2009; Hondroyiannis et al., 2008; Hondroyiannis et al., 2009). Nevertheless, those studies are aimed almost exclusively at major developed countries. We share a focus on small open economies with (Mihailov et al., 2011a,2011b), but they do not consider the possibility of structural changes, which might be relevant for emerging countries in general and for transition countries in particular. Our study also bears some resemblance to studies using alternative empirical frameworks aimed specifically at changes in inflation persistence along changes in monetary policy regimes (Benati, 2008); (Kang et al., 2009). Finally, our aim to shed some light on the viability of the inflation-targeting regime brings us spiritually close to cross-country (panel) studies (e.g. Mishkin and Schmidt-Hebbel, 2007; Brito and Bystedt, 2010). However, the approach of these studies is of an aggregate and unstructural nature. By contrast, our approach enables us to reveal that the effects of the inflation-targeting regime can differ even across relatively similar countries and so structural country-specific analysis might be more appropriate than aggregate estimation with heterogeneous country panels.

3. Open-economy hybrid NKPC

We start our exposition with the seminal hybrid NKPC model laid out in GG:

$$\pi_t = \gamma_f E_t \pi_{t+1} + \gamma_b \pi_{t-1} + \lambda s_t + \varepsilon_t, \tag{1}$$

where π_t denotes inflation, $E_t\pi_{t+1}$ represents inflation expectations conditional on the information up to time t, s_t is a proxy for the marginal cost (as a deviation from the steady-state), and ε_t is an exogenous inflation shock, such that $E_{t-1}\varepsilon_t = 0$. Unlike GG, we will later assume that parameters γ_f, γ_b, and λ are potentially time-varying, i.e., they may evolve over time because of the dynamic economic conditions in the converging economies under study. We provide a more detailed motivation and justification for the time-varying model in the next subsection.

Inflation persistence enters Eq. (1) not only through the backward-looking term γ_b, but potentially also through parameter λ as long as the markets that determine the evolution of the forcing variable (the output gap, in our case) are rigid. Moreover, the inflation process can exhibit rather persistent properties (in terms of autocorrelation) even under quite stable economic conditions where lagged inflation provides a good guess about the future inflation path. From the monetary policy perspective, however, we are chiefly interested in the intrinsic (or structural) price rigidity tracked by the backward-looking term, insomuch as it captures the persistence inherent to the inflation process itself. It contributes to a lower ability of the monetary authorities to make disinflation policy costless and, to a certain extent, implies that they have limited credibility. High intrinsic persistence can also signal poorly anchored inflation expectations. The reduced-form parameters are non-linear functions of three structural parameters (a subjective discount factor, β, the probability that prices remain fixed, θ, and a fraction of backward-looking price setters, ω):

$$\begin{aligned}
&\lambda \equiv (1-\omega)(1-\theta)(1-\beta\theta)\phi^{-1}\\
&\gamma_f \equiv \beta\theta\phi^{-1}\\
&\gamma_b \equiv \omega\phi^{-1}\\
&\phi \equiv \theta + \omega(1-\theta(1-\beta)).
\end{aligned}$$

The structural parameters can provide a closer view of the nature of the structural changes that have been affecting the economies in question. Specifically, one might be interested in finding out whether the fraction of backward-looking setters has decreased, for example, as a result of the inflation-targeting regime, or how the average time for which prices remain fixed ($1/(1-\theta)$) drifts over time. Given that the CE countries under study are all small open economies, we derive a new version of the hybrid NKPC model in the spirit of Galí and Monacelli (2005), accounting for the potential impact of external factors on inflation. Recently, Mihailov et al. (2011a) used the purely forward-looking small-economy NKPC model of Galí and Monacelli (2005) and evaluated the relative importance of domestic and external drivers in the new member states. Our version of the model can be viewed as an extension of their approach to a hybrid form and a time-varying framework.

In line with Galí and Monacelli (2005), we now assume that CPI inflation can be expressed as:

$$\pi_t = \pi_{H,t} + \alpha\Delta TT_t, \tag{2}$$

where $\pi_{H,t}$ is the *domestic inflation*, ΔTT_t denotes the current-to-past period change in the terms of trade[2] and parameter α measures the openness of the economy. Analogous to Eq. (1), the dynamics of domestic inflation are given by[3]:

$$\pi_{H,t} = \gamma_f E_t \pi_{H,t+1} + \gamma_b \pi_{H,t-1} + \lambda s_t. \tag{3}$$

Plugging Eqs. (3) into (2) and making use of the fact that $\pi_{H,t} = \pi_t - \alpha\Delta TT_t$, we get:

$$\pi_t = \gamma_f E_t(\pi_{t+1} - \alpha\Delta TT_{t+1}) + \gamma_b(\pi_{t-1} - \alpha\Delta TT_{t-1}) + \lambda s_t + \alpha\Delta TT_t,$$

[1] There are numerous studies initiated by the ESCB Inflation Persistence Network, but they are mainly based on microdata.

[2] Galí and Monacelli (2005) use an inverse definition of the terms of trade, i.e., they define them as the import price index over the export price index.

[3] We leave out the error term for expositional ease.

and after some rearrangement we obtain a hybrid open-economy NKPC model of the form:

$$\pi_t = \gamma_f E_t \pi_{t+1} + \gamma_b \pi_{t-1} + \lambda s_t + \alpha \left\{ \Delta TT_t - \gamma_f E_t \Delta TT_{t+1} - \gamma_b \Delta TT_{t-1} \right\}. \quad (4)$$

To motivate the economic interpretation of the last term in Eq. (4), it is useful first to consider the two extreme cases where either γ_f or γ_b is equal to one.[4] If $\gamma_f = 1$ and $\gamma_b = 0$, then the term becomes $(\Delta TT_t - E_t \Delta TT_{t+1})$ and model (4) collapses into the purely forward-looking open-economy model introduced by Mihailov et al. (2011a). Intuitively, as pointed out by Mihailov et al. (2011a), current demand for domestic goods in the pure NKPC will increase when $\Delta TT_t > E_t \Delta TT_{t+1}$ because the relative price of domestic goods is lower than that anticipated in the future, and this increased demand causes upward pressure on current inflation. Conversely, when $\Delta TT_t < E_t \Delta TT_{t+1}$, current-period demand for domestic goods will fall, as agents expect their relative price to decline in the future, and this exerts downward pressure on current inflation.

In a fully backward-looking setting, implied by $\gamma_b = 1$, the term shrinks to $(\Delta TT_t - \Delta TT_{t-1})$. Again, the effect on inflation can be inferred by comparing the two terms in brackets, i.e., by investigating whether $\Delta TT_t > \Delta TT_{t-1}$ or $\Delta TT_t < \Delta TT_{t-1}$ holds true. The crucial difference, however, is that backward-looking agents now anticipate the future path of the terms of trade with respect to their past value, since the lagged value is used as a simple way to make a forecast. Note that this implies, other things being equal, higher inflation inertia than in the closed-economy model, because the terms of trade now serve as another channel contributing to persistence.

When the universe is formed by both forward and backward-looking agents, one simply compares ΔTT_t with the linear combination of $E_t \Delta TT_{t+1}$ and ΔTT_{t-1}, where coefficients γ_f and γ_b serve as multiplicative constants or weights. Hence, the linear combination[5] can be viewed as a weighted average of the next-to-current difference in the terms of trade anticipated by forward-looking and backward-looking agents. Since a difference in the terms of trade is nothing else but a change in the relative prices of imports (in terms of exports), it can, in a certain respect, be interpreted as a measure of import inflation. Thus, the hybrid open-economy NKPC assumes the same hybrid formation for inflation expectations no matter whether they are defined as a rise in the general price level of goods and services or as the relative price of imports in terms of exports.

4. Econometric framework

The hybrid open-economy NKPC as represented by Eq. (4) cannot be estimated directly because $E_t \pi_{t+1}$ is, in essence, a latent quantity which must be proxied by some observable variable. Since for CE countries (like most emerging countries) inflation expectations taken from surveys cover only a very short time span and are of questionable quality, we proceed by making the common assumption that economic agents form their expectations rationally and replace $E_t \pi_{t+1}$ by π_{t+1}. Note, however, that this leads to endogeneity bias, as future inflation is by construction correlated with the error term. To see this, let $\vartheta_{t+1} \equiv \pi_{t+1} - E_t \pi_{t+1}$ be the unpredictable forecast error and rewrite Eq. (4) into the following form[6]:

$$\pi_t = \gamma_f \pi_{t+1} + \gamma_b \pi_{t-1} + \lambda s_t + \alpha TT_t + e_t, \qquad \left(e_t \equiv \varepsilon_t - \gamma_f \vartheta_{t+1} \right). \quad (5)$$

To tackle the problem of endogeneity bias in a fixed parameter setting one routinely resorts to GMM techniques, which rely on the use of instrumental variables. This has also been common practice in previous research focusing on CE countries (Franta et al., 2007; Mihailov et al., 2011a; Vašíček, 2011). A constant parameter model, however, does not seem to be fully appropriate for CE countries, as they all went through a period of transition and implemented new monetary policy regimes along the way. The changes are likely to have translated into structural instability of the key parameters in Eq. (5) and also into changes in the volatility of inflation shocks. Moreover, the traditional GMM estimator may suffer from high sensitivity to the choice of instruments. This GMM property is a particularly well-known issue in the context of NKPC estimation (for the most recent and comprehensive overview see Mavroeidis et al., 2013).

To demonstrate the tendency to parameter instability and the sensitivity of the GMM estimator to the choice of instruments we calculated GMM estimates of model (5) for different time spans as well as different instrument sets. The outputs from this exercise are depicted in Table 1. The columns of Table 1 represent GMM estimates based on several instrument sets typically exploited in the literature, with the last column showing the results for instruments selected via Bayesian model averaging, which is the approach further pursued in our analysis (see Section 4.2). One can clearly observe notable variation in the estimated parameters across different instrument sets. This variation mirrors the considerable uncertainty related to the choice of instruments when the economic theory itself cannot provide any guidance. For example, for Poland different instrument sets deliver notably different messages in terms of the relative evolution of the backward and forward-looking terms over time. Whereas with some instrument sets there seem to be no major changes in the properties of the NKPC over time, some other sets suggest dramatic and rather unexpected developments, such as an increase in the backward-looking term and, inter alia, inflation persistence.

A similar picture can be drawn if we look at Table 1 from the perspective of different time spans. The rows of the table present GMM estimates for the full sample and two distinct subsamples. The sample was split into two even subsamples so that the first represents the period prior to the introduction of inflation targeting and the early years of this framework.[7] The results again suggest relative changes in the forward versus backward-looking term in these two periods, although a clear pattern (a relative decrease in the backward-looking term) is observable for the Czech Republic only. The instability of the coefficients over time was confirmed by formal GMM breakpoint tests, which allow for changes not only in the coefficients, but also in the variance–covariance matrix of the error terms (Andrews–Fair Wald and LR-type tests were used).[8] The results in Table 2 show that the null hypothesis of parameter stability is rejected in almost all cases.

Although the results presented above provide a clear indication that parameter instability deserves deeper exploration, they are not very informative about the time-varying nature of the coefficients. In light of the specific economic factors discussed above, we tend to believe that gradual smooth changes in parameters are a more plausible framework a priori than single (or multiple) structural breaks. This view is in line with some previous literature. Namely, Primiceri (2005) argues that smooth evolution of coefficients seems to be the most flexible framework given that discrete breaks models may well describe rapid shifts in policy but seem to be less suitable for capturing changes in private sector behavior, where aggregation among agents usually plays the role of smoothing most of the changes.

[4] Although we do not impose the restriction $\gamma_f + \gamma_b = 1$ a priori, the results usually show close-to-convexity properties.

[5] If we restrict the coefficients to sum to 1, we obtain a convex combination with the straight interpretation of a weighted average.

[6] To spare space, we refer to the term $\{\Delta TT_t - \gamma_f E_t \Delta TT_{t+1} - \gamma_b \Delta TT_{t-1}\}$ merely as TT_t in the subsequent sections.

[7] Given the sample size, it seems reasonable to split the sample around the middle so that the coefficients can be reliably estimated in each subsample. Consequently, we assumed a breakpoint in 2003 Q1. This breakpoint is also reasonable given the evidence that inflation targeting might affect agents' behavior around 2 or 3 years after its implementation.

[8] We again assumed a breakpoint in 2003 Q1, but alternative breaks at earlier or later dates provide very similar results.

Table 1
Time-invariant GMM estimates across different time spans and instrument sets.

	GG99		GGL05		All lags 1–2		All lags 1–4		BMA	
	Coef	*pval*	Coef	*pval*	Coef	*pval*	Coef	*pval*	Coef	*pval*
CZ										
γ_f	0.54	*0.00*	0.43	*0.32*	0.41	*0.22*	0.41	*0.11*	0.59	*0.00*
γ_b	0.45	*0.01*	0.56	*0.13*	0.55	*0.05*	0.61	*0.01*	0.41	*0.00*
λ	0.05	*0.62*	−0.02	*0.81*	0.00	*0.99*	−0.07	*0.61*	−0.01	*0.93*
α	0.13	*0.85*	1.60	*0.00*	2.08	*0.13*	1.45	*0.11*	1.73	*0.00*
Subsample 96–02										
γ_f	0.44	*0.21*	0.58	*0.17*	0.10	*0.90*	0.28	*0.47*	0.51	*0.00*
γ_b	0.56	*0.07*	0.47	*0.20*	0.84	*0.13*	0.77	*0.05*	0.48	*0.00*
λ	0.00	*1.00*	−0.14	*0.46*	0.16	*0.83*	−0.09	*0.71*	−0.04	*0.80*
α	0.95	*0.05*	1.55	*0.00*	2.79	*0.07*	1.50	*0.02*	1.96	*0.00*
Subsample 03–10										
γ_f	0.61	*0.03*	0.71	*0.00*	0.67	*0.03*	0.64	*0.01*	0.68	*0.00*
γ_b	0.41	*0.09*	0.31	*0.10*	0.37	*0.14*	0.39	*0.09*	0.29	*0.01*
λ	0.05	*0.55*	0.04	*0.72*	0.04	*0.69*	0.02	*0.73*	0.05	*0.47*
α	−0.43	*0.65*	1.09	*0.22*	−0.58	*0.57*	0.04	*0.97*	1.35	*0.06*
POL										
γ_f	0.58	*0.00*	0.59	*0.01*	0.81	*0.00*	0.71	*0.00*	0.63	*0.00*
γ_b	0.43	*0.00*	0.42	*0.01*	0.29	*0.09*	0.35	*0.03*	0.40	*0.05*
λ	0.01	*0.96*	−0.03	*0.81*	0.04	*0.77*	−0.02	*0.91*	−0.01	*0.92*
α	0.02	*0.82*	−0.02	*0.79*	0.07	*0.41*	−0.01	*0.88*	−0.04	*0.57*
Subsample 96–02										
γ_f	0.51	*0.03*	0.60	*0.27*	0.73	*0.00*	0.64	*0.00*	0.38	*0.06*
γ_b	0.49	*0.00*	0.43	*0.15*	0.35	*0.03*	0.39	*0.00*	0.32	*0.04*
λ	−0.16	*0.52*	−0.17	*0.50*	−0.03	*0.77*	−0.01	*0.94*	0.40	*0.15*
α	−0.03	*0.77*	−0.04	*0.70*	0.01	*0.95*	0.05	*0.34*	0.09	*0.08*
Subsample 03–10										
γ_f	0.54	*0.01*	0.52	*0.13*	0.65	*0.04*	0.38	*0.27*	0.45	*0.03*
γ_b	0.54	*0.00*	0.49	*0.01*	0.64	*0.07*	0.76	*0.00*	0.67	*0.00*
λ	0.04	*0.82*	0.03	*0.91*	−0.14	*0.69*	−0.09	*0.66*	0.01	*0.95*
α	0.21	*0.00*	0.35	*0.02*	0.60	*0.03*	0.02	*0.94*	−0.21	*0.74*
HUN										
γ_f	0.27	*0.68*	0.88	*0.11*	0.45	*0.12*	0.49	*0.15*	0.64	*0.00*
γ_b	0.56	*0.00*	0.42	*0.00*	0.57	*0.03*	0.58	*0.11*	0.51	*0.00*
λ	−0.03	*0.85*	0.06	*0.72*	−0.02	*0.88*	−0.01	*0.96*	−0.04	*0.75*
α	−0.75	*0.31*	−0.15	*0.81*	−1.18	*0.66*	−1.09	*0.66*	0.23	*0.82*
Subsample 96–02										
γ_f	0.78	*0.55*	0.37	*0.39*	0.95	*0.17*	0.77	*0.18*	0.36	*0.26*
γ_b	0.31	*0.45*	0.30	*0.04*	0.25	*0.78*	0.41	*0.36*	0.40	*0.25*
λ	0.06	*0.84*	0.01	*0.98*	0.11	*0.76*	0.07	*0.79*	−0.51	*0.32*
α	0.95	*0.39*	0.72	*0.06*	1.36	*0.40*	0.86	*0.59*	0.76	*0.42*
Subsample 03–10										
γ_f	0.76	*0.11*	0.55	*0.04*	0.46	*0.09*	0.67	*0.03*	0.41	*0.01*
γ_b	0.42	*0.24*	0.41	*0.00*	0.50	*0.03*	0.40	*0.08*	0.67	*0.00*
λ	0.05	*0.84*	0.10	*0.65*	0.01	*0.98*	0.07	*0.75*	−0.10	*0.58*
α	−1.49	*0.17*	−0.71	*0.37*	−1.02	*0.59*	−1.08	*0.65*	−1.01	*0.23*

Note: The specifications follow various models in the literature, extended for inflation in the EU to account for the openness of the respective economies. GG99 (Galí-Gertler, 1999): four lags of inflation, the output gap, unit labor costs, the interest rate spread, the price of crude oil, and inflation in the EU. GGLS05 (Galí-Gertler-Lopez-Salido, 2005): four lags of inflation, two lags of the output gap, unit labor costs, the interest rate spread, and inflation in the EU. All lags 1–2: two lags of inflation, the output gap, unit labor costs, the interest rate spread, the price of crude oil, unemployment, the short-term interest rate, the NEER, and inflation in the EU. All lags 1–4: four lags of inflation, the output gap, unit labor costs, the interest rate spread, the price of crude oil, unemployment, the short-term interest rate, the NEER, and inflation in the EU. BMA: instruments selected by BMA with a probability larger than 0.5, see Fig. 2.

One way to check for parameter instability that is driven by smooth changes (rather than abrupt breaks) is the method of flexible least squares (FLS, Kalaba and Tesfatsion, 1989). Although FLS is not a formal testing procedure it provides handy descriptive tools for analyzing variation in parameters. The objective function of the FLS estimator consists of two subcriteria, namely: the goodness-of-fit, determined by the squared residuals (the measurement error, R_M^2) and the smoothness of the coefficients, given by the sum of their squared first differences (the dynamic error, R_D^2). The relative weight assigned to each criterion is regulated through smoothing constant μ. When μ equals one, no

Table 2
Breakpoint tests with breakpoint at 2003 Q1.

Model	Czech Republic		Poland		Hungary	
	A–F Wald	A–F Lik. Rat.	A–F Wald	A–F Lik. Rat.	A–F Wald	A–F Lik. Rat.
GG99	57.932	136.735	15.436	45.476	33.054	4148.214
	0.000	**0.000**	**0.004**	**0.000**	**0.000**	**0.000**
GGL05	12.809	73.019	19.798	36.941	7.231	42.208
	0.0123	**0.000**	**0.000**	**0.000**	**0.124**	**0.000**
All lags 1–2	181.38	919.697	44.99	476.562	26.653	88.952
	0.000	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**
All lags 1–4	49.135	244.063	1.574	13.403	8.074	18.010
	0.000	**0.000**	0.813	**0.010**	**0.089**	**0.001**
BMA	3.876	5.774	12.562	24.344	2.656	13.520
	0.423	0.217	**0.014**	**0.001**	0.617	**0.009**

Note: The table shows the results for the Andrews and Fair GMM breakpoint test (based on the Wald and Likelihood ratio statistics). The first row of each instrument set shows the value of the test statistics, while the second row shows the corresponding p-value. Structural instability is indicated by p-values printed in boldface. The same instrument sets were used as in Table 1.

variation in the parameters is allowed, the dynamic error equals zero, and the measurement error attains its maximum. By contrast, when μ equals zero, one allows for maximum variation in the coefficients, which results in a perfect fit. One of the useful outputs of FLS is the residual efficiency frontier, which consists of all pairs of measurement and dynamic errors which are implied by μ and are compatible with the minimum value of the objective function. The residual efficiency frontiers for all the CE countries are shown in Fig. 1. Its shape suggests that there is systematic parameter instability in all three countries. This instability is most pronounced in the Czech Republic and rather less marked in Poland and Hungary. This can be seen from the large decrease in the measurement error once we allow for even a very small variation in the coefficients. At the same time, an inspection of the coefficients' paths for alternative values of the smoothing parameter reveals that gradual changes are a more plausible framework than the presence of a single break.

Against this backdrop, in our empirical analysis we opt for an approach which deals with the endogeneity problem similarly to GMM but can additionally tackle the issue of smooth parameter variation. Since our approach also relies on the use of instruments and as such may be prone to the same sensitivity to the choice of those instruments as the GMM estimator, we additionally use appropriate methods to reduce the uncertainty in instrument selection.

4.1. Dealing with time variation and endogeneity

To overcome the problem of endogeneity in a time-varying framework we broadly stick to the strategy proposed by Kim (2006), who suggests estimating a time-varying regression model with endogenous regressors by employing a two-step procedure. In the first step, one runs an OLS regression[9] of the endogenous variables on a set of instruments that are uncorrelated with the error term in Eq. (5). To finish the first step we get residuals $\hat{v}_t = y_t - \hat{y}_t$, estimate Σ_v by $\hat{\Sigma}_v = \sum_{t=1}^{T} \frac{1}{T} \hat{v}_t \hat{v}_t'$, and obtain the standardized residuals $\hat{v}_t^* = \hat{\Sigma}_v^{-1/2} \hat{v}_t$. These residuals are used as the additional regressors in Eq. (5), where they serve as the endogeneity correction terms. After insertion of the correction terms into the NKPC model the whole system with time-varying coefficients can be cast into the state-space form (see Subsection 4.3) and estimated using slightly modified Kalman filter formulas which correct

[9] For reasons that we will explain later we consider a time-invariant relation between the endogenous regressor and the instruments.

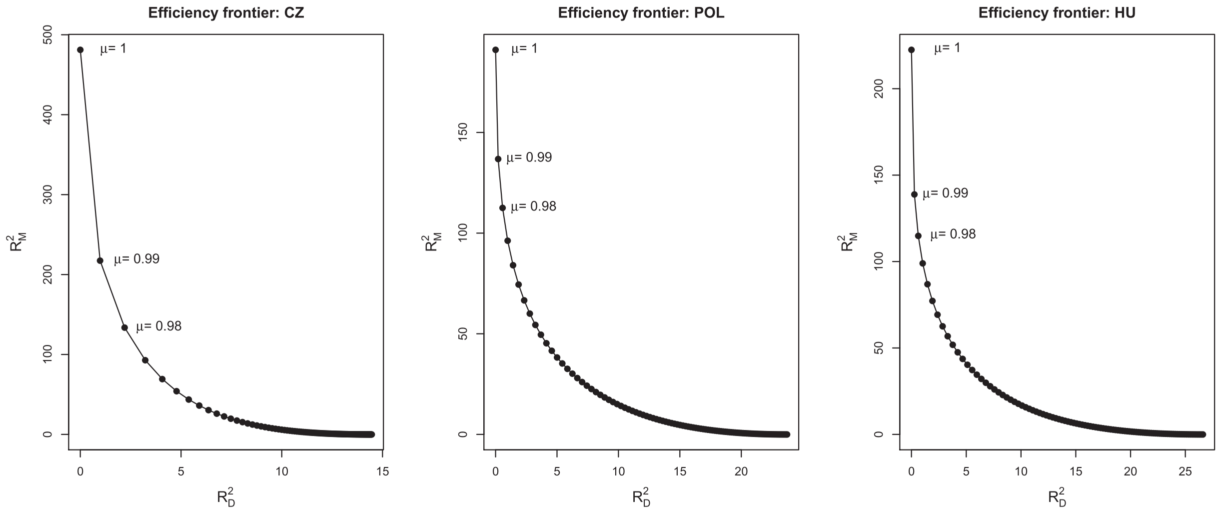

Fig. 1. Flexible least squares: residual efficiency frontier.

for bias in the variance of the estimated states. Full details are given in Kim (2006) and Kim (2008).[10]

Despite the practical appeal of the two-step procedure of Kim (2006), we are still left with considerable uncertainty about which instruments should be used in the first-step OLS regression. Another problem is that the standard estimation of linear state-space models assumes constant Gaussian shocks to the target variable, which is unlikely to hold for the inflation process (as argued, for example, in Koop and Korobilis, 2009). Applying methods that ignore possible variation in the volatility of the error term in the NKPC model may lead to serious bias of the estimated time-varying coefficients. Indeed, we find that varying volatility of shocks does matter and its omission leads to highly erratic and unstable results.[11]

Based on these facts, we improve upon the original procedure of Kim (2006) and introduce some necessary modifications. Namely, we employ Bayesian model averaging to select a proper set of instruments and use a stochastic volatility model for the error term in Eq. (5) to account for potential changes in the volatility of inflation shocks.[12] In the following two subsections we provide a deeper justification for each modification and give some technical details of our estimation approach.

4.2. Instrument selection and Bayesian model averaging

As suggested by the GMM results in Table 1 models with rational expectations such as the NKPC are subject to considerable uncertainty in instrument selection. Under these conditions, the Bayesian model averaging (BMA) provides a coherent framework to account for model uncertainty and instrument sensitivity. It is a relatively new method (see Hoeting et al., 1999) that was introduced to a wider audience in the mid-1990s. To our knowledge, BMA is new in the NKPC literature, although similar ideas have already been tossed around in the context of rational expectations models (see Wright, 2003). Unlike the 'traditional' approach to estimation of the NKPC, where a researcher typically selects instruments (and thus conditions her model) in quite a subjective manner, BMA effectively weights all the possible models based on the posterior model probability. The relative importance of each instrument can be then inferred from its posterior inclusion probability, which is equal to the sum of the posterior probabilities over the models in which it is included.

The application of BMA in our setting can be summarized as follows. Let Z be a $T \times k$ matrix of instruments containing the information available to economic agents. Under standard assumptions, the unrestricted model can be represented as:

$$y_t = a + Z_t\delta + \epsilon_t \qquad \epsilon \sim N\left(0, \sigma^2\right), \tag{6}$$

where y_t denotes the outcome variable π_{t+1}, a is an intercept, and δ is a vector of parameters. Since economic theory leaves us rather agnostic about the 'true' model, the researcher may have some uncertainty over which instruments to include or exclude. All possible combinations of instruments form the model universe $M = [M_1, M_2, \ldots, M_K]$, where $K = 2^k$. The BMA solution to the problem is to weight the outcomes of all the models by their posterior probability. The fitted value $\hat{y}_t^{BMA}$ can then be expressed as:

$$\hat{y}_t^{BMA} = \sum_{k=1}^{K} \hat{y}_{t,k} p(M_k|y, Z), \tag{7}$$

where $\hat{y}_{t,k}$ denotes a fitted value conditional on the model k, and weights $p(M_k|y, Z)$ are the posterior model probabilities that arise from Bayes' theorem:

$$p(M_k|y, Z) = \frac{p(y|M_k, Z)p(M_k)}{p(y|Z)} = \frac{p(y|M_k, Z)p(M_k)}{\sum_{s=1}^{K} p(y|M_s, Z)p(M_s)}, \tag{8}$$

where $p(y|M_k, Z)$ denotes the marginal likelihood of the model, $p(M_k)$ is the prior probability that M_k is the 'true' model, and the denominator represents the integrated likelihood, which is constant over the model universe. The expressions for the marginal likelihood $p(y|M_k, Z)$ depend on the problem at hand and vary across different kinds of models. In a linear regression setting, the marginal likelihood has a closed-form solution or can be obtained by approximation (depending on the nature of

[10] Note that we did not modify the variance of the states in our empirical exercise and our estimates of the time-varying parameters may thus be subject to higher uncertainty than that presented in Section 5. The variance of the time-varying parameters depends on the (square) value of the estimated coefficient on the standardized residuals. Since in our case it was usually quite small and statistically insignificant, the differences in the confidence intervals should mostly be negligible.

[11] Due to their apparent failure, we do not present these results here, but they are available upon request.

[12] Although the (Kim, 2006) procedure in principle allows heteroskedasticity to be dealt with by means of GARCH, the prevailing view in the literature is that stochastic volatility models are more flexible and usually outperform ARCH-type models (Kim et al., 1998).

the priors on the coefficients).[13] Before running BMA, the researcher needs to specify the model universe (set of instruments), the model priors $P(M_k)$, and the parameter priors $P(\varpi|M_k)$, with $\varpi \equiv (a, \delta', \sigma^2)'$.

In our setting, y_t represents the endogenous variables in Eq. (5)[14] and the instrument set includes four lags of inflation, the output gap, the unit labor cost, long-term interest rates, the interest rate spread, unemployment, the nominal effective exchange rate, and the crude oil price. We aimed to include the most comprehensive set of instruments consistently with previous papers, subject to data availability. We use the hyper-g prior on the coefficients proposed by Liang et al. (2008) and run the Bayesian adaptive sampling algorithm (Clyde et al., 2011) to obtain the posterior probabilities over the models.

In light of our considerations above, it may also seem reasonable to account for the time-varying relation between the target (endogenous) variable and the instruments. Recently, Raftery et al. (2010) proposed a new method called dynamic model averaging (DMA) that accounts for both model uncertainty and parameter variation. Since forecasting exercises have shown that BMA and DMA perform comparably at short horizons (see Koop and Korobilis, 2009), and given that DMA is still computationally unfeasible for large instrument sets, we regard BMA as a reasonable option for the first-step regression.[15]

4.3. The complete model

The hybrid NKPC in Eq. (5) with added correction terms, time-varying coefficients, and stochastic volatility can be cast into the following state-space representation (see Nakajima, 2011, for general representations of time-varying regression and VAR models with stochastic volatility):

$$\pi_t = c_t'\kappa + x_t'f_t + \psi_t, \quad \psi_t \sim N\left(0, \sigma_t^2\right) \tag{9}$$

$$f_{t+1} = f_t + u_t, \quad u_t \sim N(0, \Sigma) \tag{10}$$

$$\sigma_t^2 = \gamma exp(h_t) \tag{11}$$

$$h_{t+1} = \rho h_t + \eta_t, \quad \eta_t \sim N\left(0, \sigma_\eta^2\right), \tag{12}$$

where $c_t \equiv (v_{t,\pi}^*, v_{t,gap}^*)'$ is a vector of the endogeneity correction terms, $x_t \equiv (\pi_{t+1}, \pi_{t-1}, s_t, TT_t)'$ is a vector containing key model covariates, κ is a vector of constant parameters, and $f_t \equiv (\gamma_{f,t}, \gamma_{b,t}, \lambda_t, \alpha_t)'$ represents a vector of time-varying coefficients.

The time-varying coefficients are constrained to follow a random walk, which allows for both permanent and transient shifts. Such a specification is designed to capture gradual smooth changes and/or structural breaks in the coefficients. Disturbances in Eq. (9), denoted ψ_t, are normally distributed with time-varying variance σ_t^2. The log-volatility, $h_t = log(\sigma_t^2/\gamma)$, is modeled as an AR(1) process.

The system of Eqs. (9)–(12) forms a non-linear state-space model with state variables f_t and h_t. The presence of stochastic volatility (the source of the non-linearity) makes traditional estimation difficult because the likelihood function is intractable. However, Bayesian inference is still possible and we can estimate the model efficiently using Markov chain Monte Carlo (MCMC) methods.[16] Now the only remaining issue is how to estimate parameter α_t on the linear combination of the terms of trade. Recall that the terms of trade enter in Eq. (5) as $\{\Delta TT_t - \gamma_f E_t \Delta TT_{t+1} - \gamma_b \Delta TT_{t-1}\}$, which means that they are dependent on the value of coefficients γ_f and γ_b, which are not known beforehand. To solve this issue we first estimate the closed-economy version of the NKPC and obtain the initial values of γ_f and γ_b. These are used to calculate the compound expression for the terms of trade. Then we estimate the open-economy version in Eq. (5) and again obtain new values for γ_f and γ_b, which may be used to recalculate the terms of trade. We repeat these steps until all the parameter values converge.

To obtain the results, we drew $M = 70{,}000$ samples from the posterior distribution and discarded the first 50,000 samples as a burn-in period. Below we report the results for the default (quite loose) coefficient priors implemented by Nakajima (2011) in his code. As a robustness check we also experimented with other parameter settings in the prior densities, but the results do not seem to be severely affected by the choice of prior. Nevertheless, the mixing properties of the Markov chain improved as the priors got tighter. To check for convergence, we computed inefficiency factors (Geweke, 1992), which measure how well the Markov chain mixes. The inefficiency factors were usually quite low (below 50). Occasionally, however, they reached values close to 100 for some coefficients. Despite this fact, it still implies that we get about $M/100 = 200$ uncorrelated samples, which is considered enough for posterior inference (see Nakajima, 2011). As a robustness check we also obtained the posterior distribution of the coefficients by sampling only every tenth draw, as this can reduce the potential autocorrelation in the chain. Results, however, remained almost identical.

As indicated above, one might also be interested in the structural parameters of the NKPC model. However, it would be extremely difficult in practice to estimate them directly from a highly non-linear state-space model. Since under quite mild conditions there is one-to-one mapping between the reduced-form coefficients and the structural parameters, we avoid direct estimation of the structural parameters and instead use a non-linear solver to obtain their value from the median of the posterior distribution of the reduced-form coefficients.[17]

4.4. Data

Our dataset combines time series taken from several data sources (ECB, Eurostat, OECD, IMF, and national statistical offices). They were mainly downloaded from the E(S)CB data warehouse, which integrates series collected by the key supranational data providers. We used seasonally adjusted (SA) data or performed our own adjustment based on X12 ARIMA when SA series were not directly available and statistical tests detected seasonality. Due to the limited data availability induced by the transition from a command to a free-market economy we are forced to use a relatively short time span, running from 1996 Q1 to 2010 Q4. One also has to take into account lower data quality — especially at the beginning of the sample, as the statistical services in CE countries still faced some difficulties in meeting newly adopted statistical standards. In this respect, the results should be interpreted with some caution.

In line with Galí and Monacelli (2005) the inflation rate is measured as the annualized quarter-on-quarter (log) difference in the

[13] BMA for linear models has been implemented in several statistical products. Here, we make use of the BAS package (Clyde et al., 2011), which is freely available in Development Core Team (2011).

[14] As we have shown above, the endogeneity problem enters the model through the replacement of inflation expectations with the observable value of future inflation. The forcing variable (the unit labor cost or the output gap) is usually considered exogenous. However, we believe that endogeneity of the output gap cannot be rejected a priori. For this reason, we formally treat the output gap as endogenous in the first-step regression and test for the presence of endogeneity in the second step by inspecting the statistical significance of the coefficient on the endogeneity correction term. The terms of trade in all specifications are considered exogenous. All other variables are predetermined since they enter the equation with some lag.

[15] Note that DMA requires full enumeration of all models, which is memory and time consuming for K greater than, say, 2^{20}.

[16] Nakajima (2011) shows how to sample from the posterior distribution of coefficients using a Gibbs sampler and provides all the necessary computational details. See Nakajima (2011) also for the reference to his Ox and Matlab codes, which were (after some modifications) used for the estimation.

[17] We fixed the subjective factor β to 0.99.

harmonized index of consumer prices. To proxy the marginal cost we stick to the output gap taken from the OECD Economic Outlook[18] rather than the commonly used unit labor cost (labor share of income, LIS). The latter measure performed rather poorly in the cross-correlation pre-analysis and in the pre-estimation exercise. The terms of trade series are calculated as the ratio of the import price index to the export price index as taken from the Eurostat database.

In addition to the lags of the variables described above, our initial instrument set includes (four lags of) the unit labor cost, unemployment, the nominal effective exchange rate, the crude oil price, the long-term interest rate, the interest rate spread, and the foreign (EU) inflation rate. The spread is defined as the difference between 3 M and overnight interbank interest rates.[19] As noted above, the number of lags – four – corresponds to that in most previous studies (see for example Galí et al., 2005).

It is important to highlight that the inflation rate (especially for Hungary and Poland), along with some other variables, shows a clear non-stationary pattern. Since it is not evident whether the non-stationarity is a result of the time-varying environment or is of an intrinsic nature, we rendered inflation stationary by shortening the estimation period to 1999 Q1–2010 Q4 and re-estimated model (4). Given that the overall results remained largely identical we report the outcomes for the longer time span only.

5. Results

5.1. Fitting inflation expectations

Unless a reliable measure of actual inflation expectations is available, most empirical studies on the NKPC assume that inflation expectations are formed rationally. In practice, inflation expectations are proxied by regressing actual future inflation on the set of instruments containing lags of different variables (see Section 4.4 for a description of the full instrument set). As explained above, Bayesian model averaging (BMA) becomes our vehicle for formally selecting the relevant instruments. By running BMA separately for each country, we assume that the determinants of inflation expectations can differ across the three countries. The R-squared of the models with the highest posterior probability was around 0.8 for all three countries.

The results of the BMA are presented in Fig. 2. These figures show the inclusion probabilities for each instrument and each country, with red bars indicating a posterior inclusion probability higher than 0.5. It can be clearly seen that these inclusion probabilities are indeed country specific, reflecting differences in the predictability of inflation in the CE countries. In particular, the results show that in Hungary, where the central bank strongly considers exchange rate fluctuations when making its monetary policy decisions (Vonnak, 2008), the foreign variables dominate over the domestic ones.

A more detailed inspection of the results reveals that these differences might be related to country-specific mechanisms for the formation of inflation expectations. In the Czech Republic, domestic inflation and real economic activity have the highest inclusion probabilities and only the fourth lag of inflation in the eurozone has an inclusion probability higher than 0.5. In Hungary, we observe the opposite pattern, with the highest inclusion probabilities for the nominal effective exchange rate and other foreign variables and economic activity, while the lags of domestic inflation have inclusion probabilities lower than 0.1. The results of the first step for Poland lie somewhere in between and the BMA selects both domestic and foreign variables. Fig. 2 also shows how the BMA estimates of future inflation track actual future inflation. The BMA estimate tracks the long-term trend in inflation and also some pronounced peaks in the inflation rate. Some of the spikes, however, were evidently unexpected given the information available to the agents. In this respect, the BMA estimates fit our intuition about inflation expectations in all three countries. It should be stressed that consistently with GG we focus on quarter-on-quarter changes in the price level, whereas most surveys provide year-on-year inflation.

5.2. Open-economy NKPC with time-varying parameters

Figs. 3–5 present the estimated time-varying open economy reduced-form coefficients for the Czech Republic, Hungary, and Poland respectively.

In general, we find evidence that the forward-looking inflation term is more important than the backward-looking one. This implies that inflation expectations formed in forward-looking fashion play an important role and are (at least partially) anchored in the CE countries. Consequently, monetary policy might be able to affect future inflation by influencing inflation expectations as such, for example by making a credible commitment to future policy actions. However, the backward-looking term, which arguably tracks intrinsic inflation persistence, remains largely significant (with a partial exception for the Czech Republic). This is a somewhat different picture to that found in studies for developed countries that use a comparable econometric framework (e.g. Cogley et al., 2010; Hall et al., 2009). These studies usually argue that a proper treatment of potential structural instabilities enables one to fully abstract from the existence of intrinsic inflation persistence.

The small overall variation in the backward and forward-looking coefficients underlines their relative stability.[20] However as usual, there are some exceptions to the rule which should not go unnoticed. Despite considerable uncertainty, as evidenced by the wide credible intervals, one can occasionally observe that the median as well as the whole posterior distribution tends to go downwards or upwards. The most conspicuous instance is the decrease of the backward-looking term γ_b to practically insignificant levels in the Czech Republic. Correspondingly, the coefficient on the forward-looking term exhibits a slight upward tendency. A similar, albeit less pronounced, pattern can be found for Poland just until the outbreak of financial crisis. In 2008, the observed resemblance seems to come to a halt. In the case of Hungary both coefficients are remarkably stable, showing only a mild decrease in the wake of the financial crisis.

Rather interestingly, we do not observe any peak or change in trend around the time when inflation targeting was adopted (i.e., in 1998 in the Czech Republic, in 1999 in Poland, and in 2001 in Hungary). This seems to imply that the shift to the new regime did not produce an immediate effect. However, we can observe some gradual changes in the coefficients for the first two countries, with the already mentioned drop in the backward-looking term 3 years after the implementation of inflation targeting in the Czech Republic.[21] These changes were accompanied by an overall slump in the inflation rate below the inflation target in both countries. In the Czech Republic, the disinflation appeared shortly after the Czech National Bank decided to move from periodic setting of targets for the end of the year to continuous targeting of headline inflation within a predefined target range.[22] The National Bank of Poland originally set inflation targets in a similar manner as in the Czech Republic, but during the first 2 years after IT implementation actual inflation ran well above the upper bounds of the announced targets. Inflation expectations were anchored to the inflation targets

[18] It seems to correspond by and large to the output gap obtained by the HP filter.

[19] We resort to this rather simplistic definition due to the limited availability of other interest rate data in the given period.

[20] However, due to the non-linear relation between reduced-form and deep coefficients, this may still imply sizable changes in the deep parameters.

[21] For Poland, Lyziak (2003) and Orlowski (2010) document that inflation expectations actually became anchored to the target path about 2 years after inflation targeting was adopted.

[22] Initially, the target was continuously decreasing, from 3–5% to 2–4% between 2002 and 2005. However, since the inflation rate already often crawled below the inflation target, the effects of the subsequent shift to point targets in 2005 and the change of the targeted inflation rate from 3% to 2% in 2009 were negligible.

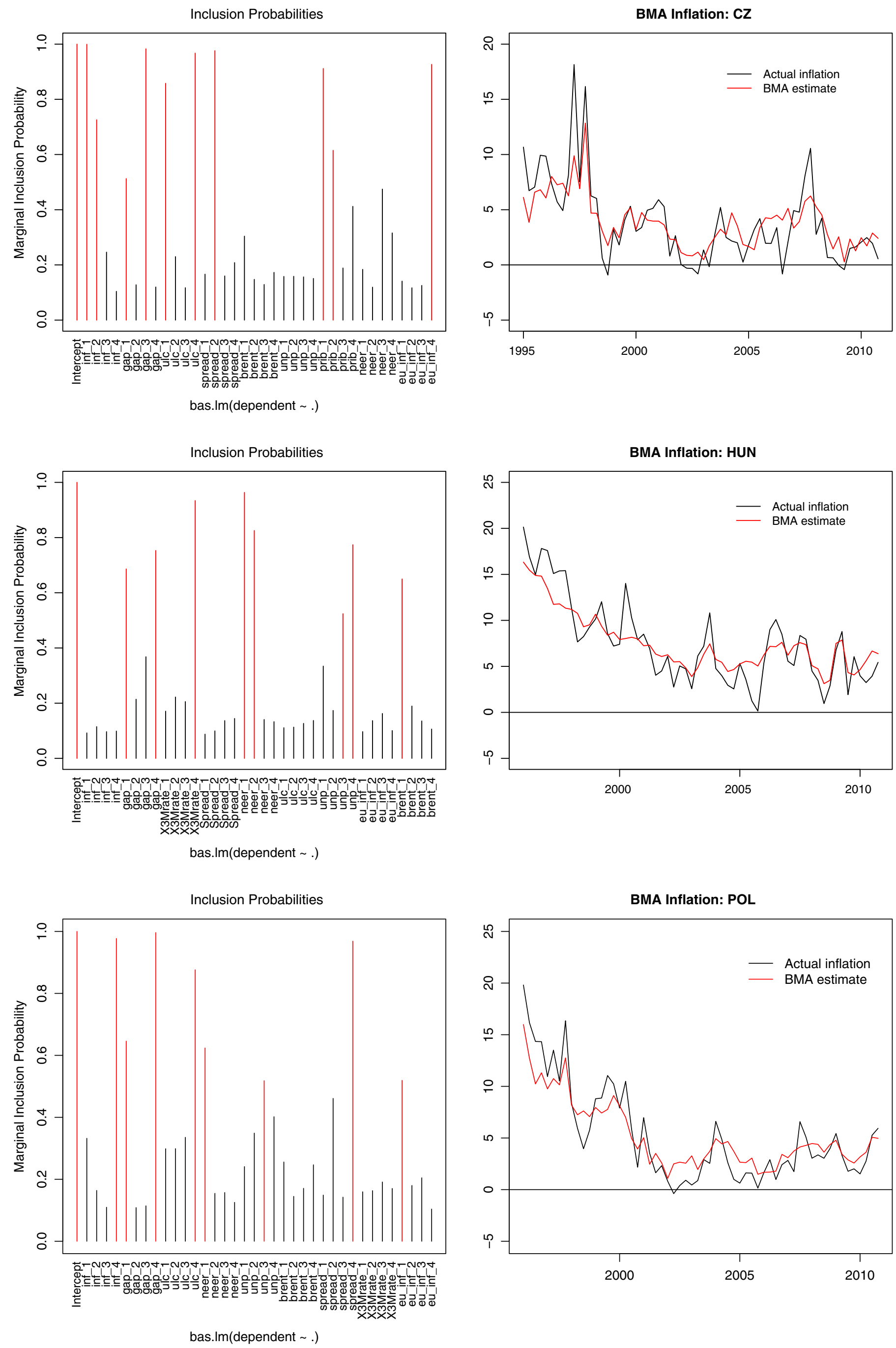

Fig. 2. BMA results: posterior inclusion probabilities and model fit.

shortly after the crawling peg was replaced with a pure float of the Polish zloty in April 2000.

On the contrary, the stability of the coefficients in Hungary prior to the Great Recession is somewhat surprising given that, at least formally, the monetary policy framework has changed significantly over the past years.[23] The main policy difference vis-à-vis the former two countries lies arguably in the role attributed to the exchange rate (Vonnak, 2008). Simultaneously with the shift to inflation targeting, the previous crawling band exchange rate regime was replaced by a 'shadow' ERM II regime of a fixed exchange rate with a fluctuation band of +/−15% around the central parity against the euro (the other two countries did not declare any specific exchange rate target and maintained a free float for most of the time). Although this de facto meant the formal adoption of an exchange rate target (alongside the official inflation target), the Hungarian central bank was not fully able to fulfill it in practice and some inflationary depreciation periods followed.[24] This is also supported by the fact that the exchange rate was identified as one of the most important factors of inflation expectations. With the evolution of the coefficients in mind, this narrative evidence seems to suggest that continuous targeting is a preferable vehicle for communicating monetary policy intentions and it seems preferable to disregard explicit

[23] Inflation targets were announced at the end of the year for the following one until 2007. A policy based on a predefined medium-term target (set at 3%) was implemented only in 2008.

[24] The most significant one in terms of its effect on inflation occurred in 2004, when inflation increased from 3% to 7%.

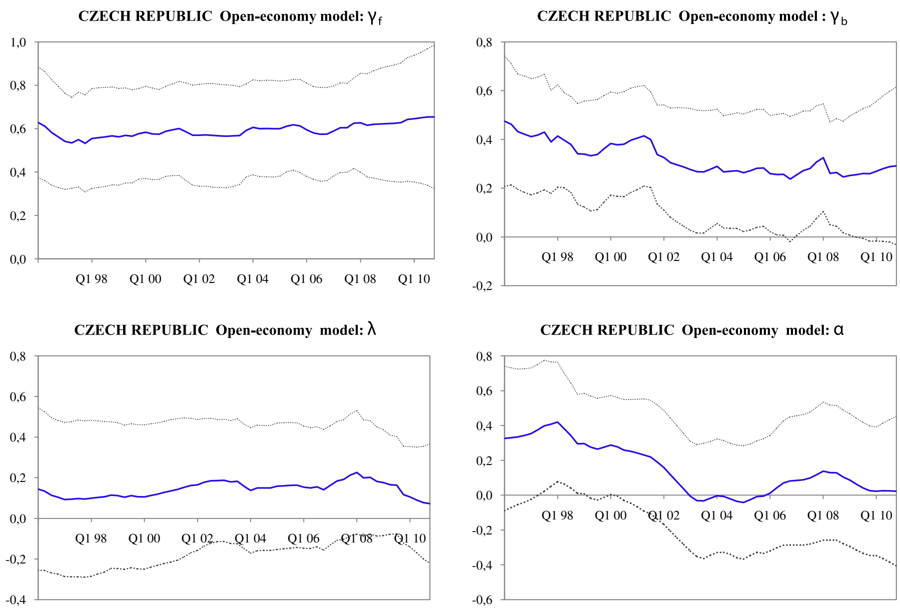

Fig. 3. Czech Republic, open-economy NKPC coefficients.

exchange rate targets, in line with the Impossible Trinity hypothesis (Obstfeld et al., 2005).

In this regard it is also useful to inspect more closely the estimated volatilities of inflation shocks, which are depicted in Fig. 6. Most notably, the volatility of inflation shocks decreased quickly a few years after the adoption of inflation targeting in the Czech Republic and Poland (standard deviation commonly around 2.5 in the Czech Republic and 2.0 in Poland) but remained quite stable and on average higher in Hungary (around 4). The most significant inflation shocks can be found in the Czech Republic (mostly induced by administrative measures such as changes in VAT), but these were short-lived and do not seem to have affected the longer-term properties of the inflation dynamics. This arguably demonstrates that one-off shocks do not affect inflation expectations if they are well anchored.

The identification of inflation forcing variables turns rather challenging. Coefficient λ, measuring the impact of real economic activity (the output gap) on inflation, remains insignificant over the entire sample in all three countries. The median estimates vary in the common range of 0–0.2, but the credible intervals are too wide to draw any particular conclusion.[25] If we limit our attention solely to the estimated median, we can observe a mild increase for all three countries starting in 2001 and common decline since the onset of global economic crisis in 2008.

As a robustness check, we also explored the performance of alternative domestic forcing variables: (*i*) the real unit labor cost, and (*ii*) the unemployment rate, but they both led to even less satisfactory outcomes than the output gap itself. Moreover, to address the uncertainty related to individual measures of real economic output we estimated a factor from all the available series potentially tracking domestic inflation pressures (real GDP, domestic demand, industrial production, index of domestic wages and salaries, real unit labor cost, unemployment rate, consumer confidence indicator, current level of capacity utilization). The resulting factor tracked the common variation of all the series (besides real unit labor cost, whose development is entirely idiosyncratic) relatively well, and was in fact quite similar to the original output gap. Consequently, the estimated coefficient λ was also almost identical.[26]

These results may have a number of explanations. First, the hypothesis often put forward for developed countries is that the Phillips curve has flattened in the last few decades. Indeed, if central banks are successful in keeping inflation rates close to the target and the volatility of inflation is limited, it is rather difficult to find a stable relationship vis-à-vis the output gap, which is substantially more volatile. Second, the trade-off between inflation and economic activity may be nonlinear, which means that the slope of the Phillips curve might depend on the actual size of the output gap: in normal times without recessions and with only mild output gaps, the relationship implied by the Phillips curve is negligible, but when larger expansions or recessions occur, the curve steepens (Stock and Watson, 2010). This argument may partially apply here, as there is some volatility apparent in the GDP growth rates in all three countries. The most blatant example relates to the economic boom in the Czech Republic and Poland in 2005–2007, when growth rates exceeded 6% annually. The output gap was then at its highest positive values, accompanied by an apparent (though still insignificant) increase in coefficient λ. Third, and perhaps most importantly in the context of transition or emerging economies, the low λ may be associated with factors specific to transition, such as supply shocks caused by the changing production structure of the economy or the fading impact of changes in regulated prices. These factors cause shifts of the Phillips curve rather than movement along it.[27] Fourth and finally, there is some intuition that open trade and capital flows weaken the effect of domestic real activity on inflation (Razin and Loungani, 2005; Razin and Yuen, 2002). This is consistent with the alternative hypothesis of a flattening of the Phillips curve, which points to the effects of globalization rather than to monetary policy (Borio and Filardo, 2007). All three countries under study are small open economies highly integrated with international markets, especially the euro area. As a result, a large proportion of domestic production is destined for foreign markets and a significant proportion of both intermediate and final products are imported. Therefore, domestic consumer inflation should rather be determined (at least partially) by external factors.

As advocated by Galí and Monacelli (2005) and subsequent authors, the terms of trade, which track relative changes in import and export prices, can thus be considered a second forcing variable for the inflation

[25] Note that in the conventional closed-economy set-up this coefficient is usually higher and significant for some periods. These results are available upon request.

[26] These results are not reported to save space, but are available upon request.

[27] However, the first-step BMA results show that real factors are often relevant for inflation forecasts (inflation expectations). It may follow that the relationship between inflation and real drivers is actually more complicated than the standard NKPC suggests.

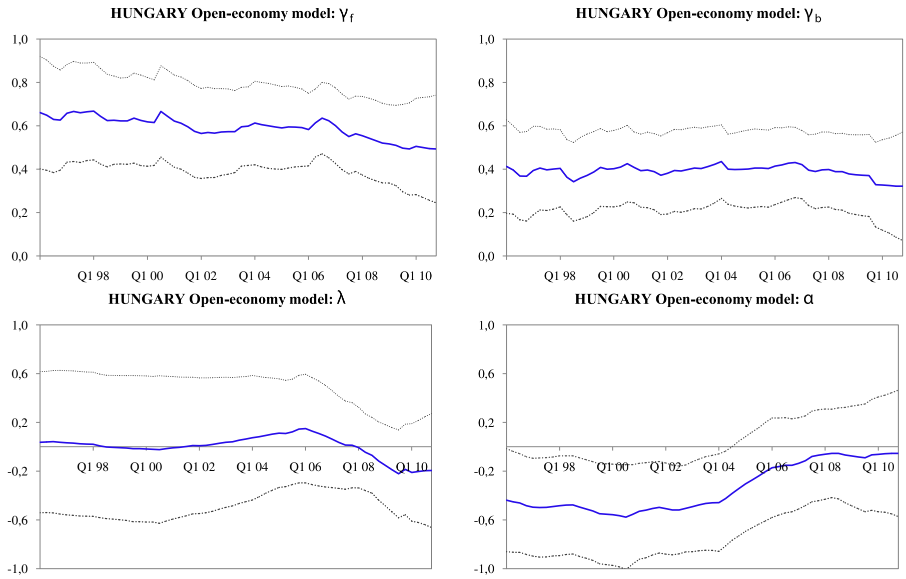

Fig. 4. Hungary, open-economy NKPC coefficients.

dynamics. However, the corresponding estimate of coefficient α_t offers mixed evidence. A kind of trade-off between domestic and foreign inflation factors can be found in the Czech Republic, where foreign factors dominate in the first half of the sample and domestic ones in the latter. For Poland, we find a predominance of foreign factors, with the corresponding coefficient α_t being significant in several periods following large depreciations (1997 Q4–1998 Q2, 2003 Q4–2004 Q3) and again at the onset of the late-2000's recession (2007 Q1–2008 Q2). For Hungary the estimated coefficient α is significantly negative on part of the sample. This goes against the underlying theory, but matches the findings of Mihailov et al. (2011a).

A possible explanation of why the effects of external factors on inflation are only temporary (and possibly non-linear) might be that they are already reflected in inflation expectations. If domestic firms engage in foreign trade, their inflation expectations are likely to be influenced by the exchange rate. For instance, in Hungary the exchange rate was on a depreciating path for the most of the time and thus arguably made agents take future currency depreciation already into account when forming their inflation expectations. Indeed in Hungary, the NEER turned out to be the most relevant variable in the first-step BMA results, but the effect of external factors in Eq. (4) proved to be very limited and disputable. By contrast, Poland experienced a few sudden and generally unexpected depreciation episodes (as noted above), which caused a huge temporary blip in its terms of trade with a significant impact on inflation.

As in the case of the domestic forcing variables we checked the robustness of our results using (*i*) the simple first difference of the terms of trade as well as their deviation from the HP-filtered trend, but the results based on the theoretical model are still preferable (as also found by Mihailov et al., 2011b), (*ii*) both the difference and the deviations from the HP-trend of the NEER, but these variables were generally insignificant.

5.3. Are the NKPC deep parameters truly structural?

While structural coefficients are routinely reported in papers based on the time-invariant framework, time-varying studies do not usually go that far. The idea that the deep structural coefficients of the NKPC vary over time is rather controversial. However, macroeconomic developments are inseparable from changes on the microeconomic level. When a change occurs in the macroeconomic setting, agents' behavior might gradually adapt to the new conditions. In particular, recent evidence suggests that firms' decisions on the frequency of price adjustment are prone to be state-dependent (in contrast to the time-dependent pricing that is assumed by the NKPC). As corroborated by most microeconomic studies on price setting, price adjustment is mainly influenced by the level and variability of inflation (see Klenow and Malin, 2010, for a survey). Similarly, Fernandez-Villaverde and Rubio-Ramirez (2008) show within the DSGE framework that movements of pricing parameters are indeed correlated with inflation. Therefore, there is no a priori reason why either the reduced-form or 'structural' parameters of the NKPC should be time invariant.

This also seems to hold from a purely modeling perspective. Estrella and Fuhrer (2003) and others offer a suite of methodological explanations of why structural models derived from agents' optimizing behavior and based on the assumption of rational expectations do not guarantee immunity to the Lucas critique. Therefore, the stability of the structural model and its ability to withstand the Lucas critique should not be an a priori assumption, but should rather be a hypothesis subject to empirical testing. Arguably, the need to verify it is stronger in transition countries than anywhere else.

To check for potential instabilities, we reconstruct a sequence of structural coefficients from the reduced-form parameters,[28] namely, *i*) the share of backward-looking price setters ω and *ii*) the average time for which prices remain fixed as a function of θ. As already mentioned above, the structural coefficients were derived under the assumption that the subjective factor, β, is fixed to 0.99. The results are reported in Fig. 7. It is also important to note that since the estimates of λ are insignificant, one has to interpret the results with great caution. Moreover, for Hungary, the reduced-form coefficient of the output gap, λ, turns negative in some periods, and this impedes obtaining the structural coefficient θ in a reasonable range (these periods are dropped and the corresponding coefficient series is discontinuous, although it still holds some information).

In the upper panel we can see some differences in the share of 'rule of thumb' firms, ω. The most economically consistent picture can be drawn for the Czech Republic. The share of backward-looking firms that adjust prices simply to the inflation observed in the previous period has been slowly trending downwards. These developments correspond to our expectations. First, during the transition, firms faced continuously increasing competition and needed to change their pricing policy with

[28] The whole distribution is obtained by calculating the structural coefficients from every posterior draw of the reduced-form parameters.

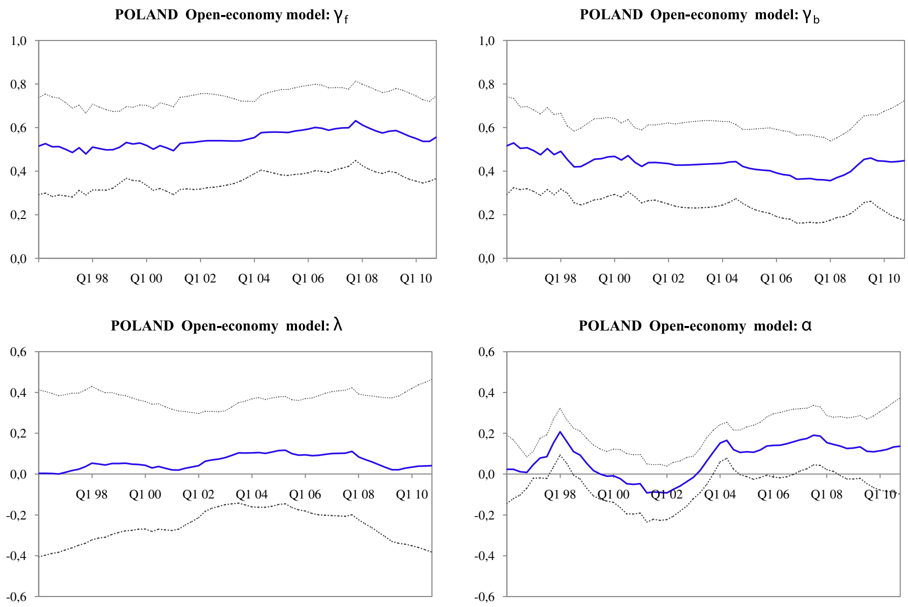

Fig. 5. Poland, open-economy NKPC coefficients.

respect to market conditions. Second, over time, the share of administered prices (which are typically set in a backward-looking manner) in overall inflation decreased. Third, forward-looking price setting is arguably subject to learning. Therefore, a decreasing share of backward-looking price setters may signal that agents are becoming more sophisticated and forming their expectations in a rational (i.e., forward-looking) rather than adaptive (i.e., backward-looking) fashion. These findings seem to suggest some kind of convergence in price setting to the euro area, where backward-looking price setting was found to be negligible (Galí et al., 2001). A similar, although a rather more humpy pattern[29] can be found for Poland, where this share decreases from substantially higher levels. The pattern for Hungary, where the estimate is subject to the largest uncertainty, is again humpy and a decreasing trend is apparent only between 2000 and 2007. Interestingly, for all three countries we can observe an increase in this share as the global crisis unfolds. The cross-country variation in the share of backward-looking price setters may be driven by multiple institutional differences in price-setting and expectation formation behavior (which are unfortunately not well explored in a cross-country context yet).

In the bottom panel we depict the structural parameter θ, capturing the overall price rigidity, and, more interestingly, the average length of price fixation $1/(1-\theta)$. Again, we find a rather stable pattern for the Czech Republic and a humpy pattern for Poland and Hungary. For the Czech Republic we find the shortest period of average price fixation, which slowly increases from two to three quarters. This result may be associated with decreased volatility of inflation, which in turn translates into longer periods over which prices remain unchanged. For Poland, we can observe a clear pattern of trendless oscillation around a mean value close to three quarters. However, there are two notable peaks, the first around the late-2000's recession and the second around the global financial crisis (late 2007 through mid-2009). A concurrent upward shift in the structural parameter θ and a deep slump in output suggest the existence of downward price rigidities, which mirror the increase in average fixation during these periods. The average price fixation values for Hungary are subject to very substantial variation and uncertainty, which can be attributed to very low and sometimes even negative estimates of λ (unreasonable price fixation values are not depicted in the chart). If anything, we can again see some variation over time.[30] The cross-country differences may be linked with numerous institutional disparities. One possible reason noted in Rumler (2007) is that more open economies may be prone to lower structural price rigidity, as firms import from volatile international markets and need to change prices more often. This could explain why the average price fixation period is shorter in the more open Czech economy than in the Polish one (discarding quite unreliable results for Hungary).

The previous results suggest that the time variation in the structural parameters may be related to macroeconomic developments. Economic intuition says that in a situation of higher and more volatile inflation it becomes more complicated for economic agents to distinguish changes in relative prices from changes in the overall price level (Lucas, 1972). Therefore, agents might be motivated to change prices more often. Besides, the level and variability of inflation can affect the size and frequency of price changes due to menu cost (Sheshinski and Weiss, 1977). That is, when inflation is high, the relative cost of keeping the price unchanged can be perceived to be smaller. Consequently, a negative relationship should exist between the average length of fixation and the inflation rate or the volatility of inflation (a discussion of this issue can be found in Taylor, 1999). The potential negative relationship between the average length of fixation and the size of the output gap may be related to the presence of downward price rigidities, i.e., prices are adjusted (upwards) more actively during expansions than (downwards) during recessions. This implies that the average price fixation is longer during recessions than during expansions.

6. What can be learned from the experience of Central European countries?

Our empirical findings on the changes in inflation dynamics have some noteworthy implications for countries considering adopting a new monetary policy regime or modifying their existing one. The differences in the implementation of inflation targeting across countries in our study shed some light on how the implementation of the IT regime affects the success of the new monetary policy regime in terms of the materialization of its potential benefits. The main benefits of IT suggested by both the theoretical and empirical literature include:

[29] This is mainly due to low estimates of λ in some periods and highly non-linear mapping between the reduced-form and structural coefficients.

[30] These findings can be confronted with stylized facts on pricing behavior available mainly for developed countries (Klenow and Malin, 2010; Taylor, 1999).

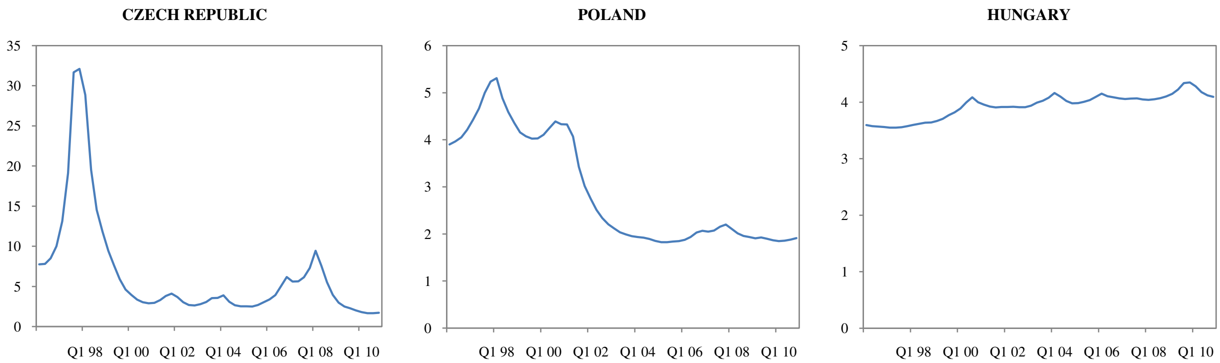

Fig. 6. Stochastic volatility.

(*i*) moderation of the level and volatility of inflation (Abo-Zaid and Tuzemen, 2012; Goncalves and Salles, 2008; Svensson, 1997), (*ii*) anchoring of inflation expectations (Blinder et al., 2008), and (*iii*) increased transparency and credibility of monetary policy and thus reduced costs of disinflation (Batini and Laxton, 2007; Mishkin, 1999).

In general, the implementation of the IT framework as a disinflation strategy proved to be successful (to some degree, at least) in all three countries, since their high inflation rates dropped to lower levels shortly after its adoption. However, as far as the volatility of inflation shocks and the credibility of the target are concerned, the three central banks seem to have met this goal with varying success. Whereas an overall reduction of the inflation level has been achieved in all three CE countries, a decrease in inflation volatility can only be observed for the Czech Republic and Poland, where the spikes in inflation volatility are rather short-lived and realized inflationary shocks dissipate quite quickly. Consequently, the link between (temporarily) higher volatility of inflation and increasing persistence seems to have been broken as well. The most significant decrease in inflation persistence has been recorded in the Czech Republic, where it has been accompanied by anchoring of inflation expectations. As a result, the Czech inflation process seems to be driven mainly by its forward-looking component. By contrast, significant inflation persistence has remained an important phenomenon in both Hungary and Poland, suggesting that past inflation bears relevant information for the formation of inflation expectations. Moreover, Hungary does not even have an observable pattern of reduced inflation persistence.

Although even ex post it is not fully clear whether the observed changes in inflation dynamics can be wholly attributed to inflation targeting, the fact that the changes occurred shortly after the adoption of the IT regime, and that the differences in IT implementation in the three CE countries match the observed differences in inflation dynamics, suggests that IT played a considerable role. Whereas the Czech Republic and Poland seem, according to several characteristics, to adhere to full-fledged inflation targeting, the monetary policy of Hungary in most categories can be characterized as inflation targeting lite (using the IT typology classification of Carare and Stone, 2006, see

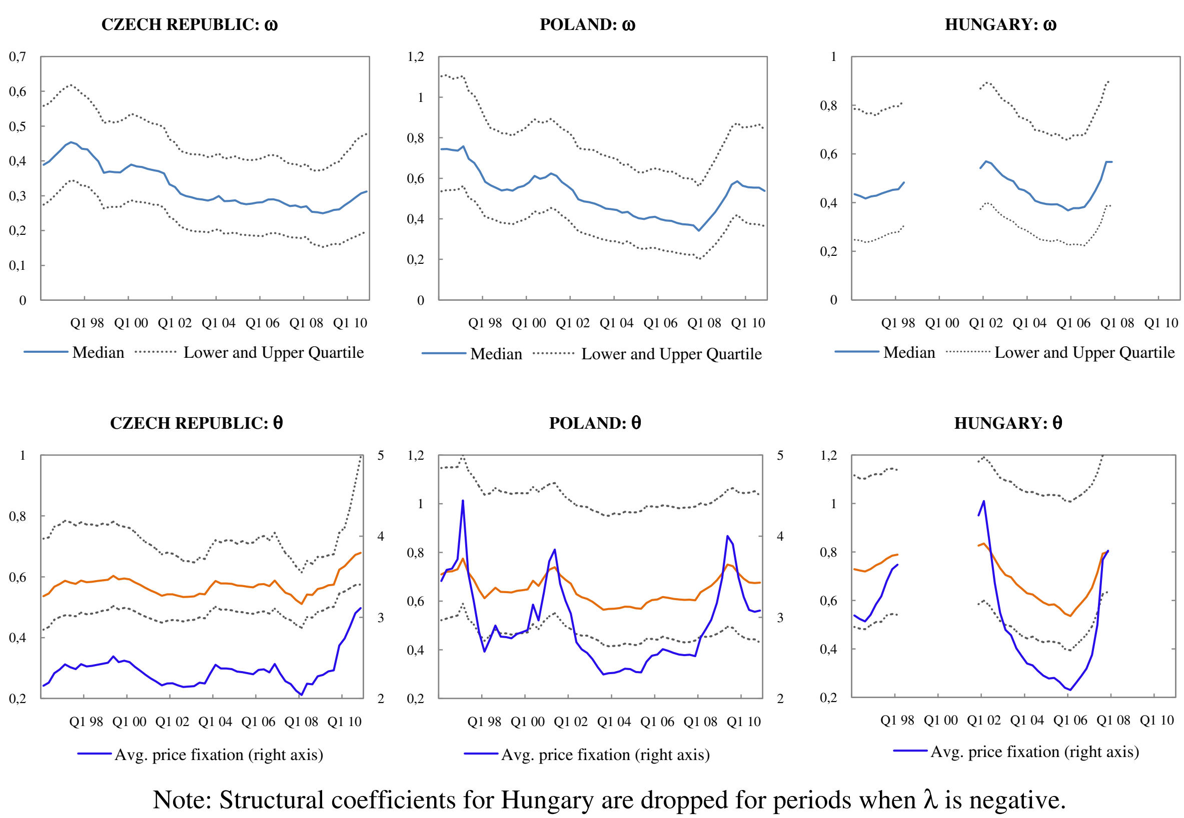

Fig. 7. Estimates of deep parameters. Note: structural coefficients for Hungary are dropped for periods when λ is negative.

Table 3
Varieties of inflation-targeting regimes and realized outcomes.

	Czech Republic	Poland	Hungary
Clarity of commitment to inflation target	Main objective is price stability	Main objective is price stability	Main objective is price stability
Numeric value of target	2% ± 1 pp	2.5% ±1 pp	3%
Definition of target and evaluation of fulfillment	Continuous, without pre-defined date of revision	Continuous, without pre-defined date of revision	Continuous, medium-term target, reassessed after 3–5 years
Exchange rate regime	Managed float	Free float	Crawling band
Credibility: past departures from target	Short-term, both positive and negative deviations	Short-term, both positive and negative deviations	Persistently above target, deviations exceed 2%
Average inflation (before/after 2003)	5.17%/2.7%	8.56%/2.72%	10.98%/5.1%
Moderation of inflation level	Yes	Yes	Yes
Inflation volatility: average of 3-year standard deviation (before/after 2001)	3.76/2.39	3.63/1.85	3.09/2.74
Moderation of inflation volatility	Yes	Yes	No
Credibility: S&P sovereign domestic-currency rating since IT implementation	A to AA	A	A to BB

Note: The table shows the dominant monetary policy characteristics over the years of IT and the mean values of variables proxying the key IT outcomes. However, both the characteristics and outcomes of the IT regime changed over time.

Table 3 for details). Most notably, the Czech and Polish numerical inflation targets are lower and are continuously assessed, and these countries' policy track records show better target fulfillment. Last but not the least, the exchange rates of the Czech koruna and Polish zloty have floated freely most of the time. On the contrary, in Hungary we do not observe major changes in inflation dynamics after the adoption of inflation targeting, and the expected positive effects of the IT regime (e.g. lower inflation volatility, reduced inflation persistence, and increased credibility) did not materialize. The narrative record of Hungarian monetary policy suggests that several aspects of the implementation of inflation targeting undermined the credibility of the target. In particular, although inflation targeting should be aimed at the domestic price level, Hungarian monetary policy has paid special attention to the exchange rate and the exchange rate channel is considered the most efficient channel of monetary policy transmission (Vonnak, 2008). This policy choice has had its costs in terms of regime credibility, as the inflation target has often been missed. This seems to reflect the Impossible Trinity of simultaneously stabilizing domestic inflation and hitting the exchange rate target in the presence of free capital flows. Consequently, the prevailing inflation rates in Hungary did not allow for a significant reduction in interest rates, and the high interest rate differential between the Hungarian forint and other currencies drove a significant expansion of foreign currency lending, which ultimately turned into a severe financial stability problem.

These conclusions support the view of (e.g. Mishkin and Schmidt-Hebbel, 2007) that inflation-targeting countries do not necessarily perform better (in terms of either output or inflation volatility) than those that do not target inflation. Based on panel data methods, they argue that what seems to matter the most is the establishment of a strong nominal anchor and the anchoring of inflation expectations. Benati (2008) also points to credibility rather than to a particular monetary policy regime as such. The time series view of three rather homogeneous CE countries with minor but important differences in policy design clearly demonstrates that these differences matter. From this perspective, the key policy objective of lowering the inflation level was met and the policy can be seen as successful. However, taking a broader view and other objectives on board, the assessment of the performance of the IT regime is more equivocal across countries. The results confirm that the adoption of the IT framework does not automatically go hand in hand with building credible and transparent policy, and these challenges become even more striking when maintaining inflation at the targeted value is not the only monetary policy objective.

7. Conclusions

This paper aims to shed some light on possible changes in inflation dynamics in the presence of structural changes in the economy and the monetary policy regime. It analyzes the dynamics of inflation through the lens of the New Keynesian Phillips curve nested within a time-varying framework using data for three CE countries. Although originally the NKPC was proposed as a structural model of inflation dynamics which is invariant to policy changes, it is likely that substantial changes on the macroeconomic level coupled with large-scale restructuring of whole economies also resulted in significant changes at the microeconomic level.

In general, we find that the considerable structural changes experienced by the economies under study, coupled with shifts in their monetary regimes, produced some notable changes in the inflation process. However, the nature of those changes is still rather heterogeneous across countries — despite their relative regional and historical affinity. Our results imply that the adoption of inflation targeting per se does not guarantee desirable outcomes such as a reduction in intrinsic inflation persistence or inflation volatility. The implementation details might matter as well, supporting the need for a structural country-level analysis.

As regards the cross-country differences, we found that intrinsic inflation persistence decreased substantially in the Czech Republic only. This implies that lower inflation can be achieved by anchoring inflation expectations and is not necessarily accompanied by output or employment losses. These findings are supported by the estimated volatility of inflation shocks, which decreased quickly a few years after the adoption of inflation targeting in the Czech Republic and also in Poland. Indeed, the predominantly forward-looking nature of the inflation process and the short-lived inflation shocks point to well-anchored inflation expectations. By contrast, almost all the characteristics describing the inflation process in Hungary are surprisingly stable. In our opinion, the comparatively low impact of the introduction of inflation targeting on inflation dynamics may be associated with the significant role of the exchange rate in Hungarian monetary policy.

The overall changes in the inflation process can be traced to changes in microeconomic behavior, in particular the price-setting behavior of firms. We found some evidence that the 'structural' coefficients describing this behavior are stable neither over time nor in the cross-sectional dimension. While the time variation can be related to the macroeconomic environment, the cross-country variation can be linked to institutional differences determining the nature of price-setting behavior and expectation formation.

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