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Inflation Convergence and the New Keynesian Phillips Curve in the Czech Republic

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Abstract The New Keynesian Phillips Curve has become an important part of modern monetary policy models. It describes the relationship between inflation and real marginal cost, which is derived from micro-founded models with rational expectations, sticky prices, and forward and backward looking behaviour. This answers the previous critique of the Phillips Curve. We estimate several specifications of the New Keynesian Phillips Curve for the Czech Republic between 1996 and 2009. We show that the GMM suffers under the problem of weak instruments leading to biased estimates. In turn, the FIML is robust and yields significant estimates of structural parameters implying a strong forward looking behaviour.

Keywords Inflation, New Keynesian Phillips Curve, marginal costs, output gap, real unit labour costs
JEL classification E31, E52, C32

1. Introduction

Inflation and inflation dynamics are important indicators of economic development. In particular, the euro area membership depends crucially on a sustainable stabilization of inflation. Therefore, inflation stabilization was often addressed in the literature, but papers have concentrated on the Balassa-Samuelson effect (e.g. Backé et al. 2003, MacDonald and Wójcik 2008).

Inflation dynamics and the nature of short-run inflation have been very debated issues over the years. Phillips (1958) initiated a discussion that has not been completed so far. Recent theoretical advances have produced alternative views of the inflation process with fundamentally different implications for an optimal monetary policy. The New Keynesian literature is built on the work of Fischer (1977), Taylor (1980), and Calvo (1983). Their approaches emphasize the forward looking behaviour of economic agents and sticky prices. One of the key New Keynesian concepts is generally referred to as the New Keynesian Phillips Curve (NKPC). This term was used initially by Roberts (1995). It was subsequently used widely by Sbordone (1998, 2001), Galí

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and Gertler (1999),[1] and Galí et al. (2001).[2] The latter authors also pioneered the estimation of the hybrid New Keynesian Phillips Curve to capture inflation persistence. Findings of GG (1999) encourage the use of these dynamic general equilibrium models in monetary policy analysis as they suggest that the observed dynamic of inflation can be understood with models derived from microeconomic foundations (Neiss and Nelson 2002).

The NKPC has two distinct features that characterize the relationship between inflation and economic activity. First, inflation is forward looking as a consequence of price formation. In particular, firms set prices on the basis of their expectations about the future evolution of demand and cost factors. The hybrid case of the New Keynesian Phillips Curve assumes that some firms use the backward looking rule to set prices. Thus, lagged inflation term is included according to this approach. The coefficients of this model depend on the probability of price adjustment, the share of forward looking firms, and a subjective discount factor. Second, there is a link between inflation and real activity, which comes through the potential correlation of real activity and real marginal cost.

There is a large number of papers on this issue. Most of them are looking at developed countries, including especially the USA and the euro area countries. By contrast, there are only a few early studies on Eastern European countries. We try to contribute to this literature with an estimation of the hybrid New Keynesian Phillips Curve for the Czech Republic during the progressed reform period from 1996 to 2009. We selected the Czech Republic, because it completed the disinflation process before the remaining Eastern European countries (see Fidrmuc 2009). We compare the Generalized Method of Moments (GMM) and the Full Information Maximum Likelihood (FIML) for the estimation of the NKPC. In theory, the GMM estimator should be strongly consistent and asymptotically normal. However, Monte Carlo simulations (see Fuhrer et al. 1995) show that GMM estimates are often biased in small samples. This leads to coefficients which are statistically insignificant. These results recommend the FIML with superior properties over the GMM also in misspecified models. Our results support these findings in the case of the Czech Republic. We compare several specifications of the NKPC using either the output gap or real unit labour costs as a proxy for real marginal cost. The reduced form estimates yield typically high coefficients of forward looking behaviour, while real activity is often negative or insignificant in most cases. Since FIML is invariant to normalizations, we present also significant structural estimates. For the Czech Republic, the interpretation of these results as a frequency of price changes and a share of forward looking firms is largely similar to values reported for developed economies. Similarly to Borys et al. (2009), our results confirm that the Czech Republic has already successfully converged to developed market economies.

The paper is organized as follows: Section 2 provides a basic description of the NKPC. Section 3 offers a brief review of the theoretical and empirical literature on the NKPC. Section 4 describes our data set and presents the results. We conclude in Section 5.

[1] Galí and Gertler hereinafter referred to as “GG”.

[2] Galí, Gertler and Lopez-Salido, hereinafter referred to as “GGLS”.

2. The New Keynesian Phillips Curve

The NKPC is one of the key elements of New Keynesian economics. It is based on the Calvo sticky-pricing model (Calvo 1983). Even though there are more realistic formulations (Taylor 1980; Fischer 1977), Calvo pricing is more comfortable and it gives similar results to those derived by more complex models. The approach assumes a continuous environment of monopolistically competitive firms. These firms are basically identical with the exception of differentiated products and pricing history. Each firm faces the same constant elasticity demand function. A fraction of firms $(1-\theta)$ is able to adjust prices[3] in period t, and future developments are discounted by a factor β. Generally, the pricing decision is based on a monopolistic competitor's profit maximization problem subject to the constraint of price adjustment in different periods. Then NKPC is derived as

$$\pi_t = \beta E_t \pi_{t+1} + \lambda mc_t^r + \varepsilon_t,$$

where $\lambda = (1-\theta)(1-\theta\beta)/\theta$. Thus, inflation depends positively on the expected future inflation and real marginal costs. We have to keep in mind that coefficient λ depends negatively on θ and β. Therefore, inflation is less sensitive to the value of real marginal cost if θ is high. In turn, full price rigidity, $\theta = 1$, implies $\lambda = 0$ and $\pi_t = \beta E_t \pi_{t+1}$. In this specific case, contemporaneous inflation is determined only by inflation expectations and the subjective discount factor.

However, Fuhrer (1997) suggests that the pure forward looking specification of prices is empirically unimportant in explaining inflation behaviour. Rudd and Whelan (2005) criticise the forward looking NKPC due to the lack of inflation inertia which enables a costless trade-off between economic activity and inflation. Moreover, price changes are caused not only by the rational expectations but also by the persistence of firms' behaviour. Firms often use past information in their expectation formation. For this reason GG (1999) consider two types of firms with different price strategies. Firms behave in a forward looking way with probability $(1-\omega)$. They use backward looking price setting with probability ω. Thus, the hybrid NKPC introduces lagged inflation as an additional variable

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

where the coefficients are functions of the underlying structural parameters:

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

The hybrid NKPC converges to the NKPC if all firms are forward looking ($\omega = 0$). An interesting feature is $\gamma_f + \gamma_b = 1$ if subjective discount factor $\beta = 1$. As the inflation

[3] The parameter θ can be also interpreted as a probability that a single firm cannot adjust its price in a given period or as a measure of price-stickiness. We have fully flexible prices if $\theta = 0$.

process is highly persistent in general, we focused only on the estimates of the hybrid NKPC below.

However, the NKPC cannot be directly estimated because the real marginal costs are unobservable. Therefore, we consider two proxy variables recommended in the literature. First, real marginal costs are generally assumed to be a cyclical variable similar to inflation. GG (1999) argue that the relation between real marginal cost and output gap is proportional under sticky prices. Boom periods are characterized by high competition for the available production factors. Consequently the real marginal cost increases with the output gap, which is defined as the difference between the log of real output and the log of the potential level of output. The disadvantage of this approach lies in the possible systematic bias, which may rise during the computation of the potential level of output.

Alternatively, under the assumption of the Cobb-Douglas production technology, we use labour income share or equivalently real unit labour costs. This approach assumes that real unit labour costs above their potential value imply a higher growth of nominal wages compared to the development of labour productivity.

3. Literature review

The NKPC has recently become an important part of monetary policy models. Its major advantage over the traditional Phillips Curve is its structural interpretation, which can be used in policy analysis. GG (1999) create an important baseline for most future discussions and pioneered the estimation of the NKPC by GMM. The baseline model was extended by backward looking behaviour. According to their approach, real unit labour costs (RULC) are preferred to model inflation persistence while the output gap measure yields negative coefficients and/or is insignificant. In the subsequent research, GGLS (2001) present the NKPC for the euro area between 1970 and 1998. The hybrid NKPC seems to fit the euro area data possibly better than the earlier estimations for the USA. Moreover, the forward looking component was found to be higher for the euro area than for the USA. These papers caused an intense discussion.

The GG approach assumes rational expectations meaning that expected inflation term $E_t(\pi_{t+1})$ is substituted with realized future inflation and forecasting error term.[4] Thus, equation (1) can be transformed to

$$\pi_t = \gamma_f \pi_{t+1} + \gamma_b \pi_{t-1} + \lambda mc_t^r + e_t, \tag{2}$$

where $e_t = \varepsilon_t - \gamma_f v_t$. However, future inflation is endogenous because the error term also includes the forecasting error, v_t. Therefore, equation (2) has to be estimated by the IV method in order to avoid biased estimates. The instruments should include all exogenous variables available at time t, which are correlated with the endogenous explanatory variables. The disadvantage of IV methods is that their results can be

[4] The relationship between expected inflation and future inflation may be expressed as $\pi_{t+1} = E_t\pi_{t+1} + v_t$, where v_t stands for a forecasting error with zero mean, which is not predictable using information available at time t.

sensitive to specification changes, e.g. with respect to the proxy for real marginal costs and selected instrument sets.

The rational expectation assumption and endogeneity problems may be avoided if inflation predictions are used. However, we do not have inflation predictions for the Czech Republic and therefore have to assume rational expectations. Nevertheless, we would like to mention some contributions focusing on this issue. This approach is discussed by Adam and Padula (2003), who use data from the Survey of Professional Forecasters. Similarly, Paloviita (2006) uses the OECD forecasts. Henzel and Wollmershaeuser[5] (2006) use data from ifo World Economic Survey. While Adam and Padula (2003) assume a finite number of professional forecasters that form expectations for a set of firms, HW (2006) consider individual firms as individual forecasters. The latter approach allows to introduce backward looking firms into the NKPC.

A departure from the rational expectations assumption leads to surprising results on the output gap position in the pure forward looking NKPC formulation. While GG (1999) conclude that the output gap fails to be a relevant proxy, the analysis using survey data find that the output gap is correctly signed and significant. HW (2006) compare their results with other similar publications and show that the forward looking coefficient γ_f seems to be lower in an analysis based on the rational expectations assumption.[6] They explain this puzzle by non-rationalities in survey data. Overall, backward looking behaviour is more relevant according to their estimations. These findings are confirmed by Zhang et al. (2009), who use several measures of the output gap and inflation.

Alternatively, Fuhrer (2006) studied the importance of the lagged inflation term in NKPC under the assumption of rational expectations. He showed that inflation persistence comes from the persistence of real marginal costs. Inflation persistence was studied also by Franta et al. (2007). Their results suggest that inflation persistence in the new member states is comparable to the inflation persistence of earlier member states. Vašíček (2009b) finds that the inflation process in the new member states is highly persistent. By contrast, Roberts (1997) presents empirical evidence of flexible prices. Hondroyiannis et al. (2007) apply the time-varying coefficient (TVC) estimation proposed by Chang et al. (2000).[7] The TVC approach provides evidence that the high weight of lagged inflation in estimates of the NKPC might be due to the specification bias and spurious correlation.

Mavroeidis (2005, 2007) raises two issues related to the selection of the appropriate estimation method. First, weak instruments lead to an overestimation of the forward looking coefficient (at all sample sizes and without any tendency to converge to the true value of the coefficient). Second, the estimations are biased if endogenous regressors are correlated with the instruments. Stock et al. (2002) offers a deeper discussion of the weak identification problem and the selection of an appropriate test procedure. Menyhért (2008) examines the problem of weak instruments related to the two stage

[5] Henzel and Wollmershaeuser, hereinafter referred to as “HW”.

[6] Averages of forward looking coefficients reported by HW(2006) are different. While rational expectations average is 0.59, survey data generate an average of 0.4 for US data.

[7] The TVC allows to separate the bias-free component of each coefficient from the other components so that specification bias can be corrected.

least squares proposed by Lendvai (2005), the countinuous-updating GMM estimator and the full information maximum likelihood estimator (FIML). He concludes that the FIML has superior properties in small samples.

Rudd and Whelan (2005) present one of the most critical papers on the NKPC. They criticize several issues: First, the forward looking NKPC is not appropriate for monetary analysis because this specification lacks inflation inertia, hence it supports a free trade-off between output and inflation. Second, unit labour costs are shown to be not a valid proxy for the real marginal cost because they do not follow sufficiently the cyclical movements of real marginal costs. Finally, the GMM is not appropriate for the estimation of the hybrid NKPC because the estimation is subject to omitted variables problem, while potential omitted variables are included in the instrument set (and correlated with π_{t+1}). Consequently, the influence of omitted variables is captured by a proxy for $E_t\pi_{t+1}$ which leads to an overestimation of γ_f. Moreover, Rudd and Whelan (2005) argue that if inflation lags are included in the instrument set, the lagged inflation role may be captured by the forward looking term.

Lindé (2005) considers the GMM estimates to be severely biased in small samples and dependent on changes in monetary policy. Based on Monte Carlo simulations, he concludes that reliable estimates of NKPC cannot be obtained by single equation methods. Therefore, he favours the FIML that performs well also under model misspecification and non-normally distributed measurement errors.

GGLS (2005) defend against most of these critical points and conclude that the main conclusions in GG (1999) and GGLS (2001) remain intact also under alternative methods of estimation. They conclude that their estimates are robust to a variety of different econometric procedures, including the GMM estimation of the closed form as suggested by Rudd and Whelan (2005) and nonlinear instrumental variables in the spirit of Lindé (2005). They also review publications with similar results using alternative econometric approaches including e.g. Sbordone (2005), who presents the two-step minimum distance estimation procedure.

Jondeau and LeBihan (2006) compare GMM and ML specifications of NKPC with output gap and RULC. The GMM leads to an overestimation of the forward looking coefficient in both specifications for all selected countries except Italy. Furthermore, Monte Carlo simulations presented by Fuhrer et al. (1995) show that GMM estimates are often statistically insignificant and unstable. A moderate degree of instrument relevance can lead to biased estimates in small samples. Therefore, they support the superior properties of the FIML estimator which is robust also in misspecified models and small samples.

Vašíček (2009a, 2009b) presents NKPC estimates for twelve new EU member states. His approach is based on the open economy Phillips Curve, which covers more broader factors than a typical analysis for closed developed economies. He recommends to focus on the post-reform period with low, one-digit inflation levels. The inflation dynamics of the NMS is found to be highly persistent with a significant forward looking component.

4. Empirical part

4.1 Data description

We focus on the Czech Republic because disinflation was achieved faster in the Czech Republic than in the remaining Eastern European countries (Fidrmuc 2009). Therefore, we can use longer time series which were less influenced by monetary regime changes than in other countries (Fidrmuc and Senaj 2006; Kočenda 2005; Kočenda and Valachy 2005; Fidrmuc and Horváth 2008). We study the period from 1996q1 to 2009q2. The variables are taken from the OECD (Main Economic Indicators) and Eurostat. They include real GDP, real unit labour costs (RULC), consumer prices (CPI), core inflation (defined as consumer prices excluding food and energy), the real effective exchange rate and the short-term interest rate. The variables are displayed in Figure 1. GDP, RULC and price variables are in logs and seasonally adjusted. For estimations we use first differences.

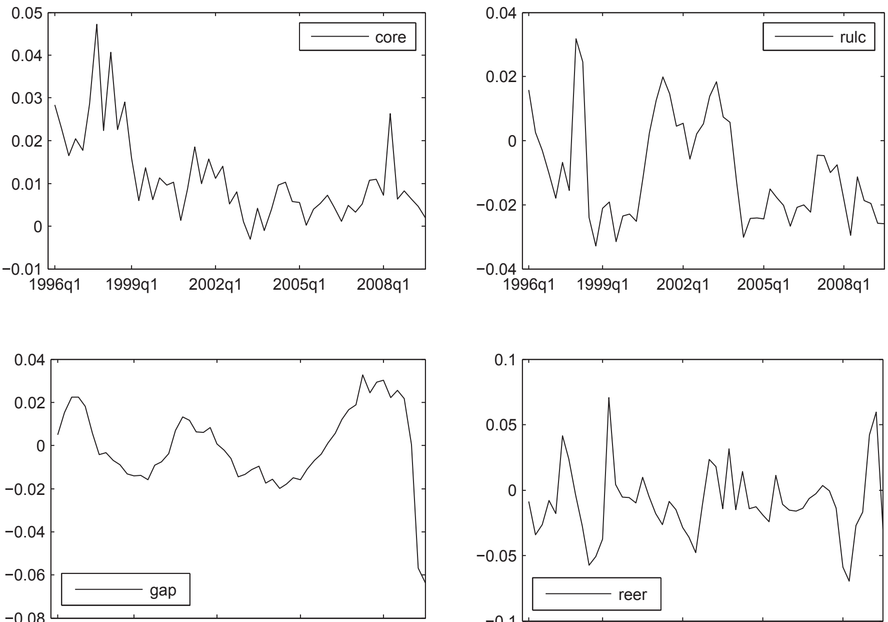

Figure 1. Selected time series for the Czech Republic, 1996–2009

We estimate the NKPC specified by equation (1) applying the iterative GMM with the starting estimates coming from the two stage least squares. Since the GMM assumes stationary time series, the variables are tested for stationarity with DF-GLS and KPSS unit root tests. The detailed results are summarized in Table 1 and stationarity may be assumed in all cases with the exception of core inflation. But core inflation is also stationary if unit root tests are performed for a sub-sample period starting in 2000q1.

Table 1. Unit root tests for selected time series, 1996–2009

	ADF-GLS H^0: unit root		KPSS H^0: stationarity	
	stat^c	stat^t	stat^c	stat^t
core	-1.1174	-2.3979	0.6043^{**}	0.1609^{*}
core^{2000+}	$\mathbf{-2.7972}^{***}$	-2.8363	**0.1227**	**0.1064**
rulc	$\mathbf{-2.4930}^{**}$	$\mathbf{-3.0431}^{*}$	**0.1491**	**0.0783**
gap	$\mathbf{-2.4928}^{**}$	-2.4144	**0.0808**	**0.0827**
reer	$\mathbf{-5.2095}^{***}$	$\mathbf{-6.0749}^{***}$	**0.0564**	**0.0563**

Notes: c – test equation includes constant; t – test equation includes constant and trend; *,**,*** denote significance at 10%, 5%, 1%, respectively; bold numbers indicate stationarity.

4.2 Generalized method of moments

We estimate several specifications with the output gap and RULC as a proxy for marginal costs. We use the following orthogonality conditions to form the baseline for the GMM specification:

$$E_t\{(\pi_t - \gamma_f E_t \pi_{t+1} - \gamma_b \pi_{t-1} - \lambda x_t)\mathbf{z_{t-1}}\} = 0,$$

where x_t stands either for the output gap or RULC as a proxy variable for real marginal cost and $\mathbf{z_t}$ is a vector of instrumental variables. We assume that expectations[8] are rational,

$$\pi_{t+1} = E_t \pi_{t+1} + v_t$$

with the disturbance term v_t to be i.i.d. GG (1999) interpret the error term as a "cost-push" shock, while Neiss and Nelson (2002) refer to this component as a "price-level shock."[9] The analyzed period is characterized by a comparably stable inflationary process which was less influenced by the reform shocks (e.g. price liberalizations) than the earlier periods. We use quarterly data over the period from 1996q1 to 2009q2. Table 2 presents the GMM and OLS estimates.[10] We compare four different instrument sets. The first basic set includes three lags of inflation and the proxy for real marginal costs. Then we add three lags of the alternative proxy variable (set 2), the real effective exchange rate (set 3) and the interest rate (set 4). Similar instrument sets were used by Menyhért (2008) for Hungary. The sum of forward and backward looking behaviour is restricted to unity. Unrestricted estimates show that the sum would be less than unity in specifications with a larger number of instruments.

[8] In general, regular quarterly forecasts are not available as an alternative to rational expectations. Moreover, inflation forecasts showed a significant forecast error during the analyzed period (see Antal et al. 2008).
[9] This means that the shock permanently raises the price level, but (provided that monetary policy is non-accomodative) it increases inflation only temporarily.
[10] We set the kernel as Barlett with a fixed Newey-West bandwidth selection. The prewhitening was not applied (prewhitening in our case does not have significant impact on the coefficients and their signs).

Table 2. NKPC Estimates for the Czech Republic

	λ	γ_f	γ_b	J-stat	Partial R^2	Adj. P. R^2	F-stat	S-Y 10%/30%
GAP 1	−0.016	0.613***	0.387*	3.602	0.406	0.344	7.223	11.12/5.15
	(0.032)	(0.204)	(0.204)	(0.463)				
GAP 2	−0.035	0.756***	0.244***	3.929	0.440	0.340	5.249	11.46/4.92
	(0.023)	(0.126)	(0.126)	(0.788)				
GAP 3	−0.026	0.715***	0.285***	5.646	0.633	0.537	5.698	11.52/4.75
	(0.017)	(0.073)	(0.073)	(0.844)				
GAP 4	−0.017	0.664***	0.336***	6.404	0.826	0.764	4.305	11.49/4.63
	(0.015)	(0.065)	(0.065)	(0.930)				
GAP OLS	−0.024	0.516***	0.484***					
	(0.057)	(0.132)	(0.132)					
RULC 1	−0.004	0.638***	0.362***	3.283	0.438	0.379	6.541	11.12/5.15
	(0.025)	(0.122)	(0.122)	(0.512)				
RULC 2	0.002	0.549***	0.451***	4.996	0.531	0.448	5.249	11.46/4.92
	(0.021)	(0.106)	(0.106)	(0.660)				
RULC 3	−0.008	0.652***	0.348***	6.196	0.680	0.596	5.698	11.52/4.75
	(0.018)	(0.084)	(0.084)	(0.799)				
RULC 4	0.001	0.500***	0.500***	6.770	0.848	0.793	4.305	11.49/4.63
	(0.017)	(0.066)	(0.066)	(0.914)				
RULC OLS	−0.054	0.518***	0.482***					
	(0.038)	(0.103)	(0.103)					

Notes: Estimated period 1996q1–2009q2; standard errors are reported in parentheses below the coefficient's estimates; p-values are reported in parentheses below J-statistics; S-Y 10%/30% are Stock-Yogo critical values for Weak IV test statistics for maximal percentage bias; *,**,*** denote significance at 10%, 5%, 1%, respectively.

Instrumental Variables: GAP 1 – core(1, 2, 3), gap(1,2,3); GAP 2 – core(1,2,3), gap(1,2,3), rulc(1,2,3); GAP 3 – core(1,2,3), gap(1,2,3), rulc(1,2,3), reer(1,2,3); GAP 4 – core(1,2,3), gap(1,2,3), rulc(1,2,3), reer(1,2,3), ir(1,2,3); RULC 1 – core(1, 2, 3), rulc(1,2,3); RULC 2 – core(1,2,3), rulc(1,2,3), gap(1,2,3); RULC 3 – core(1,2,3), rulc(1,2,3), gap(1,2,3), reer(1,2,3); RULC 4 – core(1,2,3), gap(1,2,3), rulc(1,2,3), reer(1,2,3), ir(1,2,3).

The results for the gap specification imply a high weight of forward looking behaviour. The corresponding coefficients are between 0.613 and 0.756, while the coefficients for backward looking behaviour are between 0.244 and 0.387. The coefficient of the output gap is, as in the previous literature, negative and insignificant. The specifications with the real unit labour costs proxy put lower weight on future inflation. Both forward and backward looking parameters are significant while the coefficient for the real marginal cost proxy are again negative (and insignificant) except for two cases. These figures are slightly above values from earlier studies. Vašíček (2009b) for example estimated the reduced-form NKPC for the Czech Republic and obtained γ_f equal to 0.56 for the gap specification and 0.43 for the RULC specification.

The GMM results are surprisingly similar to the OLS estimates.[11] Actually, the GMM results for the forward looking coefficient are higher than the OLS coefficient. This is especially true if we include a parsimonious set of instruments. However, using

[11] OLS estimates are also restricted to $\gamma_f + \gamma_b = 1$, the unrestricted estimates are significantly less than one.

the GMM with weak instruments often leads to an overestimation of the forward looking coefficient and misleading sampling errors biased towards the probability limit of the OLS estimator, which may be the case also in our results. The weak instrument problem can arise if future inflation is not sufficiently correlated with the selected instruments. To test the quality of instruments we employ the Stock and Yogo (2002) test based on the concentration parameter. For instrument set (including a constant) $\mathbf{Z}$ and normally distributed error terms, ε and ν, the concentration parameter μ^2 is defined as

$$\mu^2 = \delta' \mathbf{Z}' \mathbf{Z} \delta / \sigma_\nu^2,$$

where δ is the vector of coefficients estimated in

$$\pi_{t+1} = \mathbf{z}_\mathbf{t}' \delta + \nu_t.$$

With $\mu^2 \to \infty$, the GMM sampling distribution converges to the normal distribution with zero mean. However, for small values of μ^2, the distribution is nonstandard. To decide whether the instruments are weak we test $H_0 : \delta_{nc} = 0$.[12] High F-statistics indicate high relevance of instrument set and higher μ^2. Their values range between 5.2 and 7.2 (see Table 2) in all specifications and they are decreasing with the number of instruments in the instrument set. There are various interpretations of these F-statistics. Stock and Yogo (2002) report critical values for a GMM bias. If F-statistics are higher than the reported crticial value for 10% or 30% significance levels, the maximum bias of the GMM will be less than 10% or 30% of the OLS bias. Thus, in our estimations the GMM exceeds 10% in three cases and even 30% in one case for both proxy specifications. The conclusion is that the GMM is only partially able to improve the OLS estimates.

Another approach to detect instrument relevance is the examination of partial R^2 suggested by Shea (1996). Low values of this indicator discredit selected instruments to predict the endogenous variable. In our case, partial R^2 values increase with the number of instruments, as can be expected. Moreover R^2 depends positively on the number of observations. Therefore we compute also adjusted partial R^2 that takes into account the number of instruments and observations. Both these values suggest the largest instrument set. Comparing with the F-statistics, hence, partial R^2 leads to the opposite results and it does not take into account the growing GMM bias.

Finally, J-statistics show Hansen's (1982) J-statistics of overidentifying restrictions. These test statistics are equal to the value of the GMM objective function multiplied by the number of observations. Reported p-values are all above the 10% significance level and suggest the validity of overidentifying restrictions. However, the disadvantage of J-statistics is that the asymptotic distribution provides only a poor approximation for the finite-sample distribution of estimators.

Overall, the results are not very encouraging since the GMM contains more than 10%, and in one case more than 30%, of the OLS bias. Also estimated coefficients vary with the employed instrument set. The bias of GMM estimates is found by Lindé

[12] The subscript "nc" stands for "no constant" because the constant term is not included in the null hypothesis.

(2005) and Rudd and Whelan (2005) but they differ in its direction. While Lindé (2005) shows that forward looking behaviour is downward biased, Rudd and Whelan (2005) favour the backward looking specification of NKPC. Menyhért (2008) also analyzed the two stage least squares and the continuous-updating GMM estimator and concluded that also these methods are likely to produce biased estimates. Therefore, Lendvai (2005) and Menyhért (2008) recommend using the full information maximum likelihood (FIML) estimator.

4.3 Full information maximum likelihood

The FIML estimator belongs to the class of full information methods. Specification requires a multiple-equation model formed by a complete system of simultaneous equations formulated for each endogenous variable.[13] The advantage of FIML is that it is consistent also for models where the error term does not follow a normal distribution. Moreover, the FIML exploits the full information available in the complete system of simultaneous equations.[14] Our approach follows Menyhért (2008) and Fuhrer et al. (1995). It is formed by the NKPC equation (1) and a vector autoregressive model (VAR) containing the endogenous variables collected in a K-dimensional vector $\mathbf{z_t}$

$$\mathbf{z_t} = \mathbf{c} + \mathbf{M}(L)\mathbf{z_{t-1}} + \mathbf{m}(L)\pi_{\mathbf{t-1}} + \xi_{\mathbf{t}},$$

where $\mathbf{M}(L) = \mathbf{M_0} + \mathbf{M_1}L + \mathbf{M_2}L^2 + \cdots + \mathbf{M_I}L^I$, $\mathbf{m}(L) = \mathbf{m_0} + \mathbf{m_1}L + \mathbf{m_2}L^2 + \cdots + \mathbf{m_I}L^I$, I equals the number of lags and L is the lag operator. $\mathbf{M}_i$ and $\mathbf{m}_i$ are $K \times K$ matrices of coefficients and $\xi_{\mathbf{t}}$ is a vector of residuals. Our set of endogenous variables includes RULC, output gap and the real exchange rate. For each specification we also consider three different lag lengths. The results for gap specification are reported in Table 3.[15]

The first three rows contain the estimations of VAR with initially only the output gap of different lag lengths. In further specifications the VAR model is extended by RULC and the real effective exchange rate. The central part of the table presents the estimates of structural parameters θ, ω, β.[16] Finally, the right part of the table shows the reduced-form parameters calculated from estimated structural parameters. The estimates of forward and backward looking behaviour are close to unity in all cases, therefore no restrictions are applied. Forward looking behaviour receives less weight than in the GMM estimates. The estimated coefficients range between 0.580 and 0.617 and they are highly significant. The output gap coefficient is in nearly all cases positive but insignificant. A more encouraging result can be found for the estimated structural coefficients, which are highly significant. The discount factor, β, is close to one as predicted by the reduced form estimates. The share of subjects with constant prices is estimated at 0.708 to 0.898. The average duration of constant prices is calculated as

[13] Details on FIML can be found e.g. in Hayashi (2000).

[14] This can be turned into a disadvantage in case part of the system is misspecified. In that case, selecting limited-information maximum likelihood is preferable.

[15] We are grateful Jeffrey C. Fuhrer from the Federal Reserve Bank of Boston for providing us the Matlab code for FIML estimations.

[16] Due to the FIML's invariance to normalization we can explore both structural and reduced-form estimates.

$1/(1-\theta)$ and varies from 3.4 quarters to 9.8 quarters. Approximately a half of the firms are backward looking. The implied reduced-form parameters for inflation parameters are similar to estimated parameters and the coefficient for real marginal costs, λ, is close to 0.010 on average.

Table 3. The FIML estimates of the NKPC with output gap proxy

GAP	Reduced-form estimates			Structural estimates					
	λ	γ_f	γ_b	θ	ω	β	λ	γ_f	γ_b
V1L1	0.004	0.615***	0.391***	0.876***	0.556***	1.000***	0.005	0.612	0.388
	(0.007)	(0.054)	(0.055)	(0.218)	(0.118)	(0.101)			
V1L2	0.018	0.602***	0.398***	0.827***	0.406***	0.999***	0.014	0.671	0.329
	(0.011)	(0.085)	(0.072)	(0.063)	(0.024)	(0.013)			
V1L3	0.005	0.605***	0.395***	0.860***	0.564***	0.996***	0.006	0.602	0.397
	(0.009)	(0.065)	(0.065)	(0.107)	(0.122)	(0.011)			
V2L1	0.005	0.603***	0.404***	0.875***	0.545***	0.999***	0.0002	0.641	0.359
	(0.017)	(0.089)	(0.105)	(0.184)	(0.117)	(0.079)			
V2L2	0.015	0.605***	0.397*	0.802***	0.516***	0.997***	0.015	0.607	0.392
	(0.024)	(0.189)	(0.210)	(0.056)	(0.094)	(0.017)			
V2L3	0.003	0.604***	0.396***	0.898***	0.592***	0.999***	0.003	0.602	0.397
	(0.003)	(0.048)	(0.048)	(0.072)	(0.099)	(0.002)			
V3L1	0.006	0.617***	0.389***	0.862***	0.524***	0.999***	0.007	0.621	0.378
	(0.006)	(0.042)	(0.044)	(0.112)	(0.105)	(0.094)			
V3L2	0.016	0.598***	0.405***	0.708***	0.514***	0.982***	0.036	0.572	0.423
	(0.010)	(0.062)	(0.061)	(0.030)	(0.085)	(0.024)			
V3L3	−0.001	0.580***	0.419***	0.754***	0.597***	0.990***	0.003	0.612	0.388
	(0.002)	(0.035)	(0.035)	(0.000)	(0.101)	(0.002)			

Notes: Estimated period 1996q1–2009q2; standard errors are reported in parentheses below the coefficient's estimates; VKLI stands for VAR(K) and LAG(I); $I = 1,2,3$; $K = 1,2,3$ contains the output gap and is extended with rulc and reer; *,**,*** denote significance at 10%, 5%, 1%, respectively.

The results for the RULC specification can be found in Table 4. The share of forward and backward looking firms is similar to the output gap version. While the forward looking component is between 0.598 and 0.617, the backward looking component ranges from 0.381 to 0.405. Contrary to the gap specifications, λ is marginally significant and correctly signed in two models, V1L1 and V3L3, with values of 0.015 and 0.004. The structural estimates are again highly significant with average time prices remaining unchanged from 3 to 11 quarters. GG (1999) report θ above 0.8 for the USA and 0.9 for the euro area, which implies price durations from 5 to 10 quarters. Menyhért (2008)[17] reports only 3 to 4 quarters for Hungary. Using micro-data, Coricelli and Horváth (2010) estimate an average duration of price spells between 3.7 and 4.2 months in Slovakia.

The average implied value for λ from the structural estimates is 0.018, which is close to the V1L1 specification. Comparing our results to the results obtained by

[17] Menyhért (2008) and GG (1999) estimated structural NKPC only for a specification with the real unit labour costs as a proxy variable.

Menyhért (2008) for Hungary, the structural parameters are quite similar with slightly lower θ and higher ω. However, his reduced-form estimates lead significant estimates for λ in nearly all cases.

Table 4. The FIML estimates of the NKPC with RULC proxy

RULC	Reduced-form estimates			Structural estimates					
	λ	γ_f	γ_b	θ	ω	β	λ	γ_f	γ_b
V1L1	0.015*	0.598***	0.403***	0.793***	0.537***	0.996***	0.015	0.594	0.404
	(0.008)	(0.035)	(0.036)	(0.105)	(0.147)	(0.010)			
V1L2	−0.000	0.616***	0.384***	0.779*	0.528**	0.993***	0.018	0.593	0.405
	(0.003)	(0.095)	(0.092)	(0.418)	(0.264)	(0.101)			
V1L3	−0.006	0.610***	0.384***	0.807***	0.584***	0.999***	0.011	0.580	0.420
	(0.035)	(0.054)	(0.048)	(0.104)	(0.113)	(0.006)			
V2L1	0.019	0.606***	0.405***	0.789***	0.513***	0.998***	0.017	0.605	0.395
	(0.016)	(0.115)	(0.036)	(0.070)	(0.094)	(0.011)			
V2L2	−0.000	0.615***	0.384***	0.775***	0.513***	0.996***	0.020	0.600	0.399
	(0.006)	(0.071)	(0.067)	(0.151)	(0.102)	(0.016)			
V2L3	−0.013	0.606***	0.381***	0.910***	0.590***	0.998***	0.002	0.605	0.394
	(0.044)	(0.056)	(0.052)	(0.107)	(0.082)	(0.004)			
V3L1	−0.000	0.613***	0.386***	0.785***	0.518***	0.997***	0.017	0.601	0.398
	(0.000)	(0.043)	(0.043)	(0.111)	(0.109)	(0.089)			
V3L2	−0.000	0.616***	0.384***	0.681***	0.458***	0.997***	0.048	0.597	0.402
	(0.003)	(0.045)	(0.045)	(0.070)	(0.096)	(0.001)			
V3L3	0.004*	0.617***	0.383***	0.807***	0.584***	0.999***	0.011	0.579	0.420
	(0.002)	(0.010)	(0.008)	(0.104)	(0.113)	(0.006)			

Notes: Estimated period 1996q1–2009q2; standard errors are reported in parentheses below the coefficient's estimates; VKLI stands for VAR(K) and LAG(I); $I = 1,2,3$; $K = 1,2,3$ contains rulc and is extended with output gap and reer; *,**,*** denote significance at 10%, 5%, 1%, respectively.

Overall, our results show that forward looking behaviour is close to 0.6 in both specifications. More importantly, forward looking behaviour can also act as a shock stabilizer (Menyhért 2008). An interesting outcome follows from the structural estimates which provide significant structural parameters of the NKPC. They imply that slightly more than half of the firms is backward looking, which is shown by a relatively high coefficient ω. We have to keep in mind that a high value of γ_f and a high share of backward looking firms, ω, are not in contradiction because the backward looking subjects also use information from forward looking firms.

Finally, the estimated results are also similar to those obtained by other authors for the early member states of the European Union. Jondeau and LeBihan (2006) estimated the forward looking coefficients in a RULC specification at 0.6 for the EU, 0.54 for France, USA and Italy, 0.56 for Germany and 0.71 for the UK.

5. Conclusions

Price liberalization was a substantial part of economic reforms in Central and Eastern European countries. Market prices allowed an efficient allocation of resources in transition economies, but they also resulted in high inflation persistence and a loss of international competitiveness in some countries. Macroeconomic policies often focused on disinflation, but also on the reduction of unemployment. These two aims of economic policy were often seen as contradictory, although this view was contradicted by modern macroeconomic theory.

So far, there are not many deeper analyses on the relationship between inflation dynamics and aggregate output. Short time series, structural breaks, and external inflationary factors make the analysis of inflation dynamics in Eastern Europe especially difficult. Therefore, we concentrate on the Czech Republic between 1996 and 2009. The Czech Republic completed major macroeconomic reforms before other Eastern European countries. The Czech economy was not subject to a deep currency crisis and reform reversals, which were observed in some neighboring countries.

We compare two methods for the estimation of the NKPC. First, we use the GMM which dominates the previous literature. However, these results have been strongly criticized by several authors. Therefore, we apply also the FIML, which was proposed more recently. Our results support the critical conclusions formulated by Fuhrer et al. (1995), Menyhért (2008) and others. Their and our results show that the GMM results are likely to be biased. Furthermore, we demonstrate that the results may depend strongly on the selected instruments and a proxy for real marginal costs. If real unit labor costs are applied as a proxy, as recommended in the literature, the size of the bias depends on the choice of instrumental variables. If the output gap is used as a proxy, we can see that the bias is generally rather high.

Our preferred results are comparable to those reported for other countries of the EU. In particular, real activity is correctly signed in some specifications, but the coefficients are generally low or even insignificant. This implies that the New Keynesian Phillips Curve is flat in the Czech Republic in comparison to other countries. In turn, we find a relatively high share of the forward looking firms of about 60%. Thus, we can conclude that the monetary features of Czech Republic has converged to those characterized for the advanced economies.

It is rather difficult to derive implications from the Czech Republic to other transition economies in Eastern Europe. The initial conditions including the tradition of conservative monetary policy were better in the Czech Republic than in the neighbouring countries, and much better than those of the more distant East European countries. Nevertheless, we can see that these economies are converging to developed economies if their institutional settings are reformed.

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