# Hungarian Inflation Dynamics[1]

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30 January 2005
Preliminary Version
Work in progress, please do not quote

[1]Part of this research has been conducted at the Magyar Nemzeti Bank. I am grateful to Zsolt Darvas, Frederic Gaspart and Paul Reding for useful discussions and suggestions. I would also like to thank participants of seminars held at the Magyar Nemzeti Bank and at the University of Namur for their comments. All remaining errors are mine.

**Abstract**

This paper estimates traditional and New Phillips curves for Hungary over the sample period 1995Q1 to 2004Q1. We find that Hungarian inflation dynamics can be reasonably well described by an open economy extension of the New Hybrid Phillips curve specifying imported goods as intermediate production goods. Our estimation results indicate that Hungarian inflation is significantly more inertial than Euro area inflation. Hungarian inflation inertia appears to be the result of pervasive backward looking price setting behaviour, while prices are reset more frequently than in the Euro area.

# Contents

# 1 Introduction

This paper analyses short-run inflation dynamics in Hungary over the past ten years. The interest in the study of the sources and nature of Hungarian inflation is twofold. First, it is central to achieving the primary goal of Hungarian monetary policy, which, according the act 2001 LVIII, is to attain and to maintain price stability.[1] Second, the EU Accession Treaty prescribes Hungary's accession to the Euro area. The assessment of similarities and differences between Hungarian and Euro area inflation dynamics, pertinent to the evaluation of the effects of the common monetary policy, is hence essential to the elaboration of monetary policy strategies both prior to and after Hungary's accession to the Euro zone.

Modeling and estimating short-run inflation dynamics has been among the central issues in macroeconomics over the past decades. Lucas' critique of the traditional Phillips curve has led to the emergence of a new Phillips curve literature, which builds on seminal work by Taylor (1980) and Calvo (1983). These so called New Phillips curves (NPCs) are deduced explicitly from the staggered price setting by forward looking, monopolistically competing firms. They show current inflation to be linked to future expected inflation and to the real marginal cost. The parameters of the NPCs are directly linked to the behavior of economic agents and are thus exempt from the Lucas critique. Nevertheless, estimation of the NPCs has raised a number of new questions both of theoretical and empirical nature.

One major issue concerns the empirical sluggishness of inflation and the way this can be modeled theoretically. Gali, Gertler (GG, 1999) have extended the standard NPC to allow for inflation inertia by assuming a mixture of forward looking and backward looking firms. Backward looking price setting produces inflation inertia, the intensity of which is directly proportional to the fraction of backward looking firms.[2] At the same time, GG (1999) estimate this Hybrid Phillips curve for US postwar inflation and show that, while backward looking behavior is statistically significant, it is not quantitatively important. A similar finding is reported by Gali, Gertler, Lopez-Salido (GGL, 2001) for the Euro area.

This issue is linked to an econometric debate on the appropriate estimation technique of New Phillips curves. GG (1999) have estimated New Phillips curves using the Generalized Method of Moments technique, while others, like Rudd, Whelan (2001) and Linde (2002), claim that this technique is inappropriate and introduces a bias into the evaluation of backward looking behavior. GGL (2003) have responded to these critiques showing that they were 'plainly incorrect'.

In this paper, we estimate traditional and New Hybrid Phillips curves for Hungary over the sample period of 1995Q1 to 2004Q1 using instrumental variable techniques. Our aim is to find the model that best captures Hungarian

[1] The target of the Magyar Nemzeti Bank is a core inflation rate measured and published by the Central Statistical Office in Hungary.

[2] For other structural models of inflation inertia see e.g. Christiano et al. (2003), Smets, Wouters (2003), Woodford (2003).

inflation dynamics and to analyze the sources of these dynamics. While we tried to keep the estimated New Phillips curve models and their estimation technique as close as possible to those used in GG (1999) and GGL (2001) so as to obtain results comparable with theirs, both the models and the techniques had to be adapted to Hungarian specificities.

First of all, Hungary being a small open economy, we estimate open economy extensions of the hybrid Phillips curve. Open economy models vary in their way of specifying the type of imported goods and the exchange rate pass-through. In this paper, we follow two approaches as described in Kara, Nelson (2003) and extend them to allow for inflation inertia. The first approach considers imported goods as final consumption goods with full pass-through of changes in the nominal exchange rate to the inflation rate of the general price level. The second approach models imported goods as intermediate production goods; changes in the nominal exchange rate are hence fully passed through to import prices but only partially passed through to the general inflation rate.

Moreover, the estimation of Hungarian Phillips curves raises a number of questions in addition to the issues regarding Phillips curve estimations for developed economies. Hungary has been going through a phase of transition to market economy and convergence to the EU economy over the past 15 years. These transition and convergence processes have been characterized, among others, by important institutional changes, by deep structural changes in the economy and by a disinflation, the credibility of which is likely to have varied over time.

A shift in the MNB's monetary policy regime from exchange rate targeting to inflation targeting in 2001Q1 leads to the Lucas critique being particularly pertinent to the estimation of traditional Phillips curves on Hungarian data. But even the estimation of New Phillips curves may be influenced by such a regime switch if the parameters of the model do not properly capture deep behavioral parameters of economic agents. To test the stability of our estimates, we therefore perform recursive regressions of both the traditional and the New Phillips curve estimations.

Phillips curves are to describe short-run fluctuations of inflation around its long-run target. To tackle the problem of structural changes' and disinflations effects on key economic variables, we consider the effects of the convergence process to have exclusively influenced the series' trends but not the cycles around these trends. Modeling these effects is beyond the scope of this paper, we hence detrend the series and model cyclical deviations by the standard specifications. A very similar approach is adapted by Coenen, Wieland (2000) in Euro area estimations. The use of detrended series is discussed in more detail later on.

Considering the shortness of the sample period, the estimation technique had to be adapted to estimators' small sample properties. Instead of the efficient GMM estimator used in the estimation of developed economies' Phillips curves, we have used the Two Stage Least Squares (2SLS) estimator our New Phillips curve estimations. Unlike efficient GMM estimator, the 2SLS estimator does not make use of estimated fourth moments which may be particularly imprecisely estimated in the case of short samples. In addition, we also check the robustness of our results with respect to the use of various instrument sets.

While the aforementioned considerations leave us cautious about the interpretation of our results, the following findings appear to be quite robust.

First, we confirm the Lucas critique for the traditional Phillips curve. At the same time, we find that Hungarian inflation dynamics is reasonably well described by the open economy extension of the Hybrid Phillips curve specifying imported goods as intermediate production goods.

Second, we use this model to compare Hungarian and Euro area inflation dynamics and find Hungarian inflation to be significantly more persistent than Euro area inflation: our results indicate that the weights of past and of expected future inflation rate are roughly equal in the determination of Hungarian inflation; in contrast, as reported by GGL (2001), the weight of lagged inflation in the determination of Euro area inflation is significantly lower than that of future expected inflation.

Third, the estimation of structural parameters indicates that the persistence of Hungarian inflation is the result of pervasive backward looking behavior, while prices are reset more frequently in Hungary than in the Euro-zone. The fraction of backward looking price-setters is estimated to be between 1/3 and 1/2 in Hungary while in the Euro zone this fraction is reported to be between 0 and 1/3. At the same time, the average duration of prices in Hungary is between 1.6 to 3.4 quarters compared to the 10 to 12 quarters reported by GGL (2001) for the Euro zone. Recursive estimation of structural parameters indicates the stability of these results.

The remainder of the paper is organized as follows. Section 2 gives a detailed discussion of the data. Section 3 presents traditional Phillips curve estimates. Section 4 describes structural closed economy models of New Phillips curves and their estimation. Section 5 presents two extensions of the New Hybrid Phillips curve to open economies. Section 6 discusses the robustness of our estimates and Section 7 concludes.

# 2 Data

The data used in our estimations are quarterly series of Hungarian key economic variables taken from the database of the Magyar Nemzeti Bank (MNB) Quarterly Projection Model (NEM). The sample period is 1995Q1 to 2004Q1.

Figure 1 shows the historical path of the annualized rates of headline and core inflation, and of the logs of real unit labor cost, real exchange rate and real GDP ($x_t = 100 \ln(X_t)$).[3] The vertical line in the graphs indicates the period of the monetary policy regime switch, 2001Q3.

The core inflation rate is the current target of the MNB monetary policy. It is computed by the Central Statistical Office in Hungary from the Consumer Price Index (CPI) by the exclusion of following products and services: nonprocessed foods, market energy prices, regulated prices, privately owned housing services.[4]

[3] The series of real unit labor cost and real exchange rate are logs of indexes (1999=1).
[4] See e.g. Methodological Notes on the computation of core inflation on www.mnb.hu.

The real unit labor cost is computed from the nominal unit labor cost in the competitive sector deflated by the CPI.[5] The real exchange rate is the inverse of the real effective exchange rate in the NEM database. It expresses the rate of the foreign CPI price index, expressed in Hungarian currency, to the Hungarian CPI price index. An increase (decrease) in the real exchange rate hence indicates a real depreciation (appreciation).

Hodrick-Prescott trends of the series are displayed by dotted lines in the figure. As can be seen in the top panels of the figure, the CPI and the core inflation rate have steadily decreased over the entire sample period. While at the beginning of the period, inflation rate was at approximately 40 percent, it has been reduced to one digit levels by the end of the period. At the same time, the rate of decrease seems to have declined over the sample period. In addition, trends can also be observed in the series of real unit labor cost, real exchange rate and of the real GDP.

The trends in most of these series are specific to the political and economic situation Hungary has experienced over the past fifteen years. The transition to market economy and the convergence process to the EU economy have been characterized by fundamental institutional reforms and deep structural transformations of the economy, which can in general not be explained by standard economic models. As noted e.g. in Darvas, Vadas (2004), the economic downturn during the first half of the nineties has been matched by a massive rise in inflation, while the disinflation was accompanied by relatively high growth rates of real economic activity which is contrary to conventional wisdom. This is confirmed by the pattern of dynamic cross-correlations of the output gap with leads and lags of inflation as displayed in the top left-hand panel of Figure 2.[6] While for developed economies the cross-correlation of current output gap is reported to be positive with leads of inflation and negative with lags of inflation,[7] in Hungarian data, the cross-correlation is negative at any leads and lags of inflation. At the same time, as shown in the top right-hand panel of Figure 2, the pattern of dynamic cross-correlation of the current output gap with detrended inflation is similar to that found in developed economies. While the effects of the convergence process may have hence induced some non-conventional dynamic relationship between economic aggregates, these effects appear to have principally concerned the trends in these series. The dynamic relationships of the series' cyclical fluctuations seem to be similar to those in developed economies and can hence arguably be modeled by standard business cycle models.

On the basis of these considerations, we estimate standard models of Phillips curves using detrended series. A similar approach is adapted e.g. by Coenen, Wieland (2000) in their estimations of a small scale model for the Euro area. Coenen, Wieland (2000) only use two variables, the inflation and the output gap

[5] The NEM database does not contain any producer price index. The use of the GDP deflator does not seem to be theoretically more appropriate than the use of the CPI. Still, we checked the sensitivity of our results to this alternative deflating and found them robust.

[6] The figure shows correlations of output gap with core inflation rate. The pattern with CPI inflation is very similar.

[7] See e.g. Fuhrer, Moore (1995), GG (1999).

in their estimations. In contrast, we use various explanatory and instrument variables, the trends of all of which have been removed.

We are aware that this method is a shortcut, which may introduce various biases into our estimation results. First, as discussed in Coenen, Wieland (2000), focusing on inflation deviations from trend would only be theoretically appropriate if the 'source of the disinflation had been a credible, fully anticipated, gradually phased-in reduction in the policy makers' inflation target.' This has most likely not been the case in Hungary. However, Coenen, Wieland (2000) analyze the sensitivity of their estimation results to this implicit assumption and show that it does not imply significant distortions for the estimations.

Second, the assumption that the convergence process has influenced the long run dynamics only is a further simplification introducing another potential source of bias into our results. The changes in the inflation target as well as real convergence might indeed have effects on the short run dynamics of inflation which we exclude by our approach of applying standard models to detrended series. Modelling the effects of the convergence process on short-run inflation dynamics is however beyond the scope of the present paper. Instead, we check the sensitivity of our results by reestimating all the considered specifications with non-detrended variables. We discuss the effects of detrending later on.

As opposed to Coenen, Wieland (2000) who use a deterministic trend to detrend inflation and output, we use the Hodrick-Prescott filter with the standard smoothing parameter for quarterly series of $\lambda = 1600$.[8] The advantage of the HP filter is that it is quite flexible and does therefore not imply too long cycles. The drawback is that it may create artificial business cycles. We have conducted some sensitivity tests by using a deterministic quadratic trend instead of the HP filter for the CPI and core inflation, the log of real unit labor cost and the log of real exchange rate, and a log-linear trend for the real GDP.[9] The use of series detrended with the alternative detrending technique led to very similar conclusions. For a comparison, Figure 3 shows detrended series both using the HP filter (solid lines) and the deviations from a deterministic quadratic trend (dashed lines). As can be seen in the figure, the differences are minor.[10] Throughout the entire paper, our results will be displayed for estimations using HP detrended series. Results for non detrended series are displayed in the tables and discussed in footnotes when differences are significant.

# 3 Traditional Phillips Curve

In this section, we estimate traditional Phillips curves for Hungarian data and check the stability of our estimates.

[8] Coenen, Wieland (2000) use exponential trend for the inflation in the Euro area, France and Italy. For the real GDP they use a log-linear trend. See also references within Coenen, Wieland (2000).

[9] The quadratic term has been highly insignificant in the log of real GDP.

[10] For a more detailed discussion of the implications of different detrending methods for New EU Member States business cycles see Darvas, Vadas (2004).

## 3.1 Closed Economy Model

The traditional Phillips curve describes short run inflation dynamics by lagged values of inflation and some cyclical indicator. Cyclical indicators can be, among others, the output gap, unemployment rate, capacity utilization or real marginal cost. Denoting the inflation rate by $\pi_t$ and the cyclical indicator by $x_t$, the Phillips curve is commonly specified as follows:

$$\pi_t = \sum_{i=1}^{h} \beta_i \pi_{t-i} + \lambda x_{t-1} + \epsilon_t. \tag{1}$$

A real expansion (contraction) is expected to be related positively (negatively) to the inflation rate. We will use the output gap and the real marginal cost as proxied by real unit labor cost as cyclical indicators. In this case, the coefficient $\lambda$ is expected to be positive. In addition, since there is no long-run trade-off between inflation and real variables, the sum of inflation coefficients is in general expected to be unity. By using cyclical deviations from long run trends however, we eliminate the long run from our estimations. The $\sum_{i=1}^{h} \beta_i = 1$ restriction looses its relevance in this context.

The traditional Phillips curve is reported to describe post-war US and Euro-area inflation reasonably well. Estimates of equation (1) for the Euro-area and the United States can be found e.g. in Rudebusch, Svensson (1999) and Gali, Gertler, Lopez-Salido (2001). These authors include four lags of inflation and use the output gap as cyclical indicator; they confirm sign and value restrictions on the coefficients for both US and Euro area data. GGL (2001) note at the same time, that despite this apparent empirical success of the traditional Phillips curve, the Lucas critique still remains an issue.[11]

Table 1 summarizes the results of the OLS estimation of equation (1) for Hungarian data. Lag selection is based on the BIC.[12] Breusch-Pagan tests do not reject homoscedasticity of the error term in any variant of the estimation. Breusch-Godfrey serial correlation tests indicate however residual autocorrelation at ten percent. To correct for this, we use Newey-West error correction including 3 lagged values.[13]

The following points seem to be worth noting. First, the success of the traditional Phillips curve's estimation for Hungarian data depends on the cyclical indicator used. While the output gap does not appear to have sizeable effects on the inflation rate's cyclical fluctuation, the real unit labor cost turns out to be significant. The insignificance of the output gap is quite surprising and raises some doubts about our measure of the output gap.

Note also, that the weight of lagged inflation is relatively low. As already noted however, this is not a violation of the long-run independence of inflation and the cyclical indicator.[14]

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[11] Balakrishnan, Lopez-Salido (2002) report similar findings for UK data.

[12] The AIC would have implied the inclusion of further lags. Estimations including additional lags yielded very similar conclusions.

[13] This is implied by the standard formula considering the sample size.

[14] Note, that $\beta$ is relatively low in the estimation with non-detrended variables as well. In

## 3.2 Open Economy Model

Hungary being a small open economy, we have some reason to assume that external shocks may have an influence on the evolution of the inflation rate in addition to standard closed economy variables. We have therefore estimated an open economy extension of the traditional Phillips curve which includes some measure of external shocks:

$$\pi_t = \sum_{i=1}^{h} \beta_i \pi_{t-i} + \lambda x_{t-1} + \sum_{i=1}^{h} \gamma_i open_{t-i} + \epsilon_t, \tag{2}$$

where $open_t$ denotes the external variable. The open economy variable is usually the change of real exchange rate, the change in the terms of trade or the change in real import prices.[15] Defining the external prices as foreign prices over domestic prices, the sign of the open economy variable's coefficient is expected to be positive. It might however be difficult to isolate the relationship between external variables and the inflation rate from the closed economy relationship: real depreciation may increase domestic prices directly thereby decreasing aggregate demand; at the same time real depreciation can stimulate exports, hence increasing demand and thereby have an increasing effect on prices. Balakrishnan, Lopez-Salido (1999) estimate specification (2) for UK data using the change in terms of trade as external variable and argue that it is an important control for external shocks. In contrast, Kara, Nelson (2003a) estimate the same specification for UK data using the real exchange rate and claim that 'the coefficient of the exchange rate in the UK data seems to be very small'.[16]

Table 2 displays our estimation results for Hungary using various external variables: $dq_{t-1}$ denotes the change in the real exchange rate, $drpm_{t-1}$ stands for the change in real import prices and $dtot_{t-1}$ for the change in the terms of trade. As can be seen, real depreciation is highly insignificant across the different variants of the estimation. We have equally run estimations including more than one lags of inflation and the external variable if the AIC has supported this choice.[17] The results of the exclusion tests of external variables are collected in Table 2a. The insignificance of external variables is confirmed across all estimations with the notable exception of the specification using the change in real import prices when 3 lags of inflation and the external variable have been included.[18]

the estimation using the output gap, the coefficient of the output gap is highly insignificant. $\beta < 1$ might in this case, indicate the decreasing trend of the inflation rate. Unit labor cost is however significant while $\beta < 1$, which seems to be a violation of the long-run verticality. This confirms our intuition that the convergence process has induced some long-run relationship between variables.

[15] For a similar specification see e.g. Balakrishnan, Lopez-Salido (2002) and Kara, Nelson (2003).

[16] These estimations do not use the same sample period. However, Kara, Nelson (2003) report subsample estimations in support of the robustness of their claim over various periods.

[17] The BIC implied the inclusion of one lag for all specifications.

[18] Findings are somewhat different when non-detrended data are used. In this case, both the sum of coefficients of the change in real exchange rate and of the change in real import

### 3.3 Stability of Estimations

As already noted, due to the regime change in 2001$Q$3, one needs to be particularly cautious about structural breaks in the estimations.

Formal Chow structural break tests performed for specification (1) indicate a structural break in the parameter estimates for the range of 1999$Q$4 to 2001$Q$3 at less than 1 percent significance when the real unit labor cost is used as cyclical indicator. This range moves a bit when the output gap is used: in this case, structural break is indicated for all periods after 2001$Q$1. This entire range for which structural breaks are indicated might be explained first, by the pure statistical fact that the effect of the new structure starts to dominate if less values of the old structure are included in the second sub-period. On the other hand, the change in the policy regime might have been anticipated by agents before the change already.[19] There have indeed been political discussions in Hungary about a monetary policy regime change for quite a long time before 2001$Q$3.

To see how sensitive our coefficient estimates really are, we performed recursive regressions. Figure 4.a displays the point estimates ± standard errors of the coefficients $\beta, \lambda$ for specification (1) with the cyclical indicator being real unit labor cost. Figure 4.b shows the same coefficients using the output gap; and figure 4.c shows point estimates with one standard error bands for $\beta, \lambda$ and $\gamma$ in specification (2) using the real unit labor cost and the change in the real exchange rate. Estimations included one lag of each variable. Terminal dates go from 2000$Q$2 to 2004$Q$1. The following points should be noted.

First, as can be seen in the figures, the estimation of the lagged inflation coefficient is reasonably stable around 0.5. The point estimates of $\beta$ remain within the one standard error bands in all specifications during the entire period.

Second, the point estimate of $\lambda$ seems to be decreasing over the entire range of terminal estimation points when real unit labor cost is used. The coefficient of the output gap seems to decrease for endpoints going from 2003$Q$1 to 2004$Q$1. Although the changes in the estimates are not statistically significant, overall, the graphs seem to indicate a decreasing intensity of the relationship between the inflation rate and the cyclical indicator. It is interesting to note, that while the output gap is not significant over the entire sample period, there is a whole range of estimations with terminal dates going from 2000$Q$3 to 2003$Q$3 in which the coefficient of the output gap is significant at least ten percent.

Third, the coefficient of real depreciation, $\gamma$ appears to be negative and more than one standard error different from 0 for endpoints going from 2000$Q$2 to 2001$Q$2. These estimates result however of very small samples and should be treated with caution. For all longer samples, the coefficient of real depreciation is positive and not significant.

Finally, a structural break is also supported by subsample estimations. The

prices is significant when more than one lags are included. The change in real prices seems thus to have affected the trend rather than the cycles in the inflation rate. The effect remains however quantitatively small: the sum of these coefficients remains less than 0.2 in any case.

[19] To do: Chow forecast test

adjusted $\bar{R}^2$ indicates a better fit of both closed and open economy traditional Phillips curves for the period of the exchange rate targeting than for the entire sample period. Table 3 shows the results of these estimations including one lag of inflation and with the external variable being $open_t = dq_t; drpm_t$ respectively $dtot_t$.

In sum, evidence for the traditional Phillips curve's success in describing short-run Hungarian inflation dynamics over the past ten years is rather mixed.

First, the closed economy specification appears to describe Hungarian data relatively well when real unit labor cost is used as cyclical indicator. Using the output gap yields less promising results.

Second, the exchange rate channel is shown to be in most cases insignificant or at best of minor importance when cyclical fluctuations of the inflation rate are modeled by open economy extensions of the traditional Phillips curve.

Third, and perhaps most importantly, the Lucas critique seems to be supported by our stability tests: the monetary policy regime change seems to have influenced the parameters of the model. This fact should warn of using this specification for the evaluation of changes in the monetary policy.

We therefore estimate New Phillips curves explicitly deduced from structural models in the following section.

# 4 New Phillips Curve

A number of recent studies have tried to solve instability problems arising in the traditional Phillips curve estimations for US and Euro-area data by elaborating structural models of short run inflation dynamics. In this section we present estimation results of some of these models for Hungarian data. In this section, we estimate a closed economy model. The next section estimates two possible open economy extensions.[20]

## 4.1 Closed Economy Model

The New Phillips curve is based on individual firms' price setting behavior. The model that will be estimated is a version of the Calvo (1983) staggered price setting model extended to incorporate backward looking price setting by a fraction of firms. This model has first been presented by GG (1999). At this place, we only present the outlines of the model, for a detailed discussion see GG (1999).

There is a continuum of monopolistically competing firms in the economy, whose size is normalized to 1. As in the baseline Calvo model, each firm faces a probability $\xi$ of not being able to readjust its price. This probability is constant across firms and constant over time. In addition, GG (1999) assume two types of firms: a fraction $1-\omega$ who adjust their prices in a forward looking way, as in

[20]For earlier New Phillips curve estimates for EU accession countries see Arratibel et al. (2002). For Hungarian estimations see Hornok, Jakab (2003). These authors only estimate reduced form coefficients.

the baseline Calvo model, and a fraction $\omega$ which follow instead some backward looking rule-of-thumb in price readjustment.

These assumptions imply that the average price level, $p_t$ can be expressed as:[21]

$$p_t = \xi p_t^{fix} + (1 - \xi) p_t^* \tag{3}$$

where $p_t^{fix}$ stands for the average of fixed prices, i.e. prices that have not been readjusted in period $t$ and $p_t^*$ denotes the average of newly set prices.

As in Calvo (1983), the average of fixed prices equals the average of the previous period general price level:

$$p_t^{fix} = p_{t-1} \tag{4}$$

while the average of newly set prices is the weighted average of prices readjusted in a forward looking way, $p_t^f$, and those readjusted following the rule of thumb, $p_t^b$:

$$p_t^* = (1 - \omega) p_t^f + \omega p_t^b. \tag{5}$$

Forward looking firms set their price to maximize their future flow of profits subject to the price setting rules. Denoting nominal marginal costs by $mc_t^n$ and the time discount factor by $\beta$, the optimally readjusted price is:

$$p_t^f = (1 - \beta\xi) \sum_{k=0}^{\infty} (\beta\xi)^k E_t(mc_t^n). \tag{6}$$

Backward looking firms follow the a rule-of-thumb according to which they set their prices to the previous period average of newly set prices updated by the previous period inflation rate:

$$p_t^b = p_{t-1}^* + \pi_{t-1}. \tag{7}$$

Although the assumption of backward looking price setting might be criticized, note that it can be motivated by some costs of information gathering which are exogenous to this model. Rule-of-thumb behavior can in this case be considered as a useful shortcut. In addition the rule-of-thumb as specified in equation (7) has several appealing features as pointed out in GG (1999). First, it implies no long-run deviation of backward looking prices from the reoptimized price if inflation is stationary. Second, the rule-of-thumb is not entirely backward looking in that, by the previous period newly set price index, it takes into account previous expectations about the future.

By combining equations (3) to (7), the New Hybrid Phillips curve can be expressed as:

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E_t \pi_{t+1} + \lambda mc_t. \tag{8}$$

[21] Lower case variables denote log deviation from steady state.

The coefficients $\gamma_b, \gamma_f, \lambda$ are functions of deep parameters:

$$\begin{aligned}\gamma_b &\equiv \frac{\omega}{\phi} \\ \gamma_f &\equiv \frac{\beta\xi}{\phi} \\ \lambda &\equiv \frac{(1-\omega)(1-\xi)(1-\beta\xi)}{\phi} \\ \text{with } \phi &= \xi + \omega[1-\xi(1-\beta)].\end{aligned} \tag{9}$$

Note first, that the backward looking price leads to the inertia of the inflation rate in this setting. When there is a positive fraction of backward looking firms in the economy, the coefficient of lagged inflation rate is bigger than zero. At the extreme where all firms are forward looking, the Hybrid Phillips curve (8) reduces to the standard pure forward looking New Phillips curve.

Note further, that in this model, the sum of the coefficients of past and expected future inflation rates, $\gamma_b$ and $\gamma_f$, is related to the time discount factor $\beta$. Mathematically, if $\beta = 1$, then $\gamma_b + \gamma_f = 1$. Since time units are measured in quarters, the discount factor is expected to be very close but not equal to 1. However, for plausible levels of $\xi$ and $\omega$, the sum of $\gamma_b$ and $\gamma_f$ remains reasonably close to 1.[22]

Another interesting feature of all New Phillips curves is that they give an explicit indication for the cyclical indicator to be used. Deduced from the structural model, short run inflation dynamics is directly linked to the real marginal cost. The use of the output gap instead of the marginal cost as cyclical indicator requires additional assumptions assuring the proportionality of these variables, i.e. $mc_t = \kappa y_t$, with $\kappa > 0$ being the constant output elasticity of real marginal cost.

## 4.2 Empirical Issues

This subsection discusses some empirical issues concerning the estimation of New Phillips curves in general and their estimations for Hungarian data in particular. We first discuss the choice of the estimator, then the choice of the cyclical indicator and last the choice of the instrument set used in our estimations.

### 4.2.1 Estimator

Most previous studies have used single-equation Generalized Method of Moments (GMM) to estimate New Keynesian Phillips curves. Rudd, Whelan (RW, 2001 and Linde (2002) however criticize this method. RW (2001) argue, that by estimating a reduced form of the hybrid Phillips curve, their conclusions concerning the relative weight of backward looking price setting is very different

[22] If e.g. $\beta = 0.9$, the value of $\gamma_b + \gamma_f \in (0.95, 1)$.

from GG (1999)'s conclusion: according to RW, GG underestimate the importance of backward looking behavior. Linde (2002) agrees with RW in that limited information estimations yield in general biased estimates of Phillips curves. In contrast to RW however, Linde (2002) concludes that single-equation GMM tends to overestimate inflation inertia. In addition, as shown by Linde (2002) the GMM method implies changes in structural parameters when the monetary policy changes. Linde suggests the use of full information estimation techniques to avoid these problems.

In a response, GGL (2003) show that the claims of both RW (2001) and Linde (2002) concerning the bias introduced by the GMM estimator are 'plainly incorrect'. They further argue that full information methods have their own drawbacks when some of the equations are not correctly specified.

In the case of Hungary, the use of full information methods seems to be particularly problematic. Especially, the specification of a monetary policy equation appears to be difficult first, because as already noted, there has been a shift in the monetary policy regime in 2001; and also because monetary policy does not seem to have followed any stable underlying systematic reaction function over the sample period. Therefore, the more widely used single-equation technique will be used throughout this paper. We will at the same time try to check the stability of our estimates by performing recursive regressions.

While we stick to the instrumental variables technique, the Two Stage Least Squares (2SLS) estimator has been preferred to the more widely used efficient GMM estimator because of the shortness of available series. The efficient GMM estimator can be shown to be consistent and to achieve the lower bound of estimates' asymptotic variance. However, the small sample properties of this estimator are likely to be poorer than those of methods not using fourth moment estimates, like the 2SLS. In addition, in the case of homoskedasticity, the 2SLS estimator corresponds to the efficient GMM.[23]

#### 4.2.2 Cyclical indicator

As already noted in the previous section, the cyclical indicator is implied to be the real marginal cost by the structural model. The real marginal cost depends in turn on the specification of the production function. As shown in GG (1999), based on the Cobb-Douglas production technology, the real marginal cost can be expressed as $MC_t = \frac{S_t}{\alpha_n}$ where $S_t \equiv \frac{W_t N_t}{P_t Y_t}$ is the real unit labor cost, and $\alpha_n$ is the labor's share parameter in the production function. In percentage deviation from steady state, this can be expressed as $mc_t = s_t$ which suggests the use of real unit labor cost's deviation from its steady state as a proxy for real marginal cost in the estimation of the New Phillips curve. As shown by GG (1999) and GGL (2001), real unit labor cost enters the New Phillips curve significantly and with the correct sign for both US and Euro area data. In contrast, the output gap enters the New Phillips curve with the opposite negative sign.

As discussed e.g. in Neiss, Nelson (2002) the different performance of the

[23] See e.g. Hayashi (2000) Chapter 3.

output gap and the real unit labor cost in New Phillips curve estimations may have two potential explanations. First, the output gap as measured by the real GDP's deviation from a trend might not be a good measure of the theoretical output gap defined as the deviation of real GDP from its potential level which would prevail if prices were fully flexible. Second, the assumption of the proportionality of output gap and real unit labor cost, which is implicit in the use of output gap as cyclical indicator, does not seem to be confirmed by data. Gali, Gertler (1999) have reported for US data e.g. that real marginal cost, as proxied by real unit labor cost, lags the output gap over the cycle while it co-moves contemporaneously with the inflation rate.[24]

A comprehensive discussion of this issue is beyond the scope of the present paper. It shall simply be noted at this place, that in Hungarian data, the real unit labor cost seems to be more synchronized with the output gap than in US data. As can be seen in the right-hand panels of Figure 2, the cross-correlations of the alternative cyclical indicators with leads and lags of inflation seem to be quite similar. The relatively synchronized co-movement of output gap and real unit labor cost in Hungarian data is also supported by Figure 5 which displays the cross-correlations of current output gap with leads and lags of the real unit labor cost. The dynamic relationship between these variables does hence have no clear indications for whether the unit labor cost might perform better in estimations compared to the output gap. We will present the results of estimations using real marginal cost and discuss the robustness of these results to replacing marginal cost by the output gap.

#### 4.2.3 Instrument set

The performance of the 2SLS estimator, as that of any other instrumental variable estimator, crucially depends on the relevance of instruments.[25] As discussed in Shea (1996), 'relevance in a multivariate context requires that the instrument set have components important to the endogenous explanatory variable that are linearly independent of those important to exogenous variables included into the regression'. This means that the instrument set is relevant if it can explain a large enough fraction of the endogenous explanatory variables', and thereby of the dependent variable's, variance directly and not only indirectly, i.e. by explaining the exogenous explanatory variables' variation. Shea suggests to use the partial $R^2$ and the adjusted partial $R^2$ statistics to check for the relevance of instruments.[26]

In most Phillips curve estimations, the instrument set contains lags of explanatory variables. Such an instrument set turns out, however, to perform poorly on Hungarian data. Therefore, the instrument set chosen contains the following variables in addition to a constant term: two lags of inflation, one lag

[24] The choice of output gap vs. real unit labor cost has been extensively debated in recent literature. See e.g. Sbordone (2002), GG (1999), Neiss, Nelson (2002) for some examples.

[25] See e.g. Nelson, Startz (1990).

[26] For a detailed description of the computation of partial $R^2$ see Shea (1996). For an alternative test for weak instruments see e.g. Stock, Yogo (2003).

of real unit labor cost, the contemporaneous growth rate of real GDP, current and lagged ratio of budget deficit to the GDP, and current and lagged non processed food price inflation rate.[27] This instrument set has a partial $R^2$ of 0.56 and an adjusted partial $R^2$ of 0.37 in the closed economy framework.[28] The robustness of our results to the choice of the instrument set will be discussed in Section 6.

### 4.3 Reduced Form

This subsection reports estimates of coefficients $\gamma_b, \gamma_f$ and $\lambda$ in equation (8). Following Gali, Gertler (1999), these estimates will be referred to as 'reduced form' since they are estimated directly without the identification of the underlying deep parameters $\beta, \xi, \omega$.[29]

Assuming expectations are rational, the error term $E_t(\pi_{t+1}) - \pi_{t+1} = \epsilon_t$ is uncorrelated with information dated $t$ or earlier. The following orthogonality condition can then be estimated by instrumental variables:

$$E_t\{(\pi_t - \gamma_f \pi_{t+1} - \gamma_b \pi_{t-1} - \lambda mc_t)\mathbf{z}_t\} = 0, \tag{10}$$

where $\mathbf{z}_t$ denotes the vector of instruments containing contemporaneous or past values of variables.

The first two lines of Table 4 show the results of our estimations of equation (10) using detrended core inflation and detrended real unit labor cost as a proxy for real marginal cost's deviation from the steady state.

The following diagnostic tests have been performed.

First, the model's overidentifying restrictions were tested. Hansen's J test does not reject the overidentifying restrictions, and hence the specification of the model. ($J = 5.973$, with $p = 0.650$). Note, that while the reported estimation was estimated by the weighting matrix $(ZZ')^{-1}$, this test is performed using the efficient weighting matrix. As noted above, the two estimators correspond in case of homoskedasticity. Testing for homoskedasticity in the case of 2SLS estimator is not straightforward.[30] We have therefore re-estimated the equation with the optimal weighting matrix and checked for the differences in point estimates. The differences turned out to be of minor importance. The estimates of $\gamma_b$ and $\gamma_f$ are equal up to the third digit. The estimate of the slope coefficient $\lambda$ is slightly different, with $\lambda = 0.029$ $(0.052)$ . This seems to support conditional homoskedasticity. In addition to this informal comparison, a more

[27] All variables in deviations from HP trend except the real GDP growth rate.

[28] For estimations using non detrended core inflation rate, we used the same variables without detrending them. Partial $R^2 = 0.48$, Adjusted Partial $R^2 = 0.24$. Estimations using detrended CPI inflation rate use the following instrument set: constant, one lag of CPI inflation, two lags of core inflation, one lag of ulc, current wage inflation, current and lagged budget deficit/GDP ratio and current and lagged non processed food price inflation. Partial $R^2 = 0.57$, Adjusted Partial $R^2 = 0.37$.

[29] We do not show separately the results of the estimations assuming purely forward looking inflation dynamics. As already noted however, the hybrid Phillips curve nests the purely forward looking Phillips curve for $\omega = 0 \Longleftrightarrow \gamma_b = 0$.

[30] See e.g. Hayashi (2000) Ch.3. p. 234.

formal modified Breusch-Pagan test was conducted. The Breusch-Pagan test requires the auxiliary regression of the squared residuals on the regression's exogenous variables to be insignificant. The above regression includes however an endogenous variable. The squared residuals have therefore been regressed on all included and excluded exogenous variables. This test confirmed our intuition: conditional homoskedasticity of residuals cannot be rejected ($p = 0.626$).

Tests of residuals' serial correlation find significant autocorrelation. The Ljung-Box test as well as a Breusch-Godfrey test modified in the same way as the Breusch-Pagan test previously, reject the null hypothesis of no serial correlation at any usual significance level. Gali, Gertler, Lopez-Salido (2001) encounter the same problem. They argue, that the serial correlation of residuals might be due to the fact that the hybrid Phillips curve model does not fully capture all the dynamics in the present data. One reason for this might be that the backward looking price adjustment takes into account more than on lagged value of inflation. We therefore tried to include up to eight lags of inflation into the Phillips curve equation. As it turned out, the additional lags were often not significantly different from zero, while both the Ljung Box and the modified Breusch-Godfrey test continued to reject the null hypothesis of no serial correlation at least five percent.[31] Instead of the modification of the specification, we corrected for serial correlation in the original specification by the use of a 3 lag Newey-West estimate of the covariance matrix.

Finally, we have tested for the exogeneity of unit labor cost and of lagged inflation. The test of a subset of orthogonality conditions does not reject the exogeneity of these variables at any usual significance level.

The $R^2$ of the estimation is 0.62. Note however, that in the case of the 2SLS estimator, this does not have a straight forward interpretation.

Overall, the estimation results are quite encouraging. All the variables enter the equation with the expected sign although the coefficient of real unit labor cost is not significant. At the same time, as shown in the last two columns of table 4, the hypothesis that the inflation coefficients sum to 1 cannot be rejected. In this framework, this suggests, that the discount factor is very close to 1, as can be expected in a quarterly model.

Note, that the coefficients of future expected inflation rate and of the lagged inflation rate are roughly equal to 0.5; although the weight of expected future inflation rate is slightly higher than that of the lagged inflation rate, their importance in the determination of current inflation is more or less equal. This is opposed to results reported for US and Euro area data by GGL (2001) according to which, the forward looking term is dominant in the short-run inflation dynamics in both of these regions.

Lines 3 to 8 in table 4 show the robustness of these results across different specifications. We performed the same estimation using the CPI inflation rate, using non-detrended core inflation rate, and using detrended core inflation rate and the output gap instead of the real unit labor cost. As can be seen, the

[31] Similar findings are reported in both Gali, Gertler (1999) and GGL (2001).

above results stand out to be robust to the choice of the inflation rate, to the decision of detrending and also to the cyclical indicator. Note especially that, as could be expected following our discussion of the previous subsection, the output gap did not yield results significantly different from those of the real unit labor cost.[32] In what follows, we restrict our description to estimations with real unit labor cost.

## 4.4 Structural form

This section tries to recover the New Hybrid Phillips curve's deep parameters, $\beta$, $\xi$ and $\omega$. These parameters mainly concern the price setting behavior of agents and thereby allow to draw conclusions with regard to mechanisms leading to inflation inertia. As shown by the high weight of lagged inflation in the reduced form estimations of the Hybrid Phillips curve, this issue is particularly relevant to Hungarian inflation dynamics.

Formally, we substitute the structural parameters for the reduced form parameters in the New Hybrid Phillips curve (8) according to the functions (9). It is known that small-sample nonlinear instrumental variable estimations are sensitive to the precise specification of orthogonality conditions.[33] We therefore follow Gali, Gertler (1999) and GGL (2001) to estimate two alternative variants of the orthogonality condition. In the first specification, the coefficient of current inflation is not normalized:

$$E_t\{(\phi\pi_t - \beta\xi\pi_{t+1} - \omega\pi_{t-1} - (1-\omega)(1-\xi)(1-\beta\xi)mc_t)\mathbf{z}_t\} = 0. \qquad (11)$$

In the second specification, the coefficient of current inflation is normalized:

$$E_t\{(\pi_t - \phi^{-1}\beta\xi\pi_{t+1} - \phi^{-1}\omega\pi_{t-1} - \phi^{-1}(1-\omega)(1-\xi)(1-\beta\xi)mc_t)\mathbf{z}_t\} = 0. \qquad (12)$$

The first four lines in Table 5 summarize our estimation results with the detrended core inflation as dependent variable and the detrended real unit labor cost as cyclical indicator. The first three columns in the table show the estimates of the structural parameters. Columns 4 to 6 give the estimates of the resulting reduced form parameters. Columns 7 and 8 show Hansens's J test resp. its significance level. The last two columns indicate results of the Wald restriction test with the null hypothesis $\beta = 1$. The test statistic follows a chi-square distribution under the null hypothesis.

The specification test does not reject the overidentifying restrictions. The restriction $\beta = 1$ cannot be rejected either at any usual significance level in neither variant of the orthogonality condition. Given the results of this test, and considering the shortness of the sample, it seems preferable to impose the restriction $\beta = 1$ to reduce the number of parameters to be estimated. We discuss estimation results based on our preferred specification (lines 3 and 4).

[32] This holds for all the following estimations, too.
[33] See e.g. Fuhrer, Moore, Schuh (1995).

The estimations of the deep parameters are estimated at a significance level of less than 1 percent in all specifications. As can be seen however, standard errors are relatively large. Two features seem interesting to note.

First, the estimates support the importance of backward looking price setting behavior in Hungary as measured by the fraction of backward looking firms, $\omega$. This parameter is estimated to be in the interval of $(0.31; 0.55)$. Despite the large error bands, both specifications imply significantly greater fractions of backward looking price setters than those reported for the Euro area by GGL (2001).[34] Note, that according to specification 2, more than half of the firms follow a rule-of thumb behavior.

At the same time however, the probability of fixed prices $\xi$ is relatively low compared to Euro area estimates. Even after taking into account the standard errors, Hungarian estimates are lower than those found for the Euro area.

Resulting estimates of reduced form coefficients support the relatively high degree of inflation inertia in Hungary. At the same time, as also discussed e.g. in GG (1999), our estimations might overstate the degree of price and inflation sluggishness. To see why, note that the coefficient of real unit labor cost is not significant in our estimations. This might be due to a downward bias introduced by the poor quality of our proxy for real marginal costs. Since the slope coefficient is inversely related to the parameters of price and inflation rigidity $\xi$ and $\omega$, if the slope coefficient is biased downward, the nominal rigidity parameters will automatically be biased upward in our estimations.

Comparing our estimation results to microeconomic evidence yields some further support in favor of the overestimation problem. Ratfai (2000) studies the price setting behavior of various Hungarian stores over the period of 1993 to 1996 and finds that stores keep their prices fixed on an average for 3.42 months, i.e. a little longer than one quarter. The average price duration can be calculated in the New Phillips curve model as $D = \frac{1}{1-\xi}$. The resulting average duration of specification 1 of the benchmark estimation is 1.92 quarters. Specification 2 implies $D = 2.64$. Hence, both specifications imply longer price durations than microeconomic evidence. It should be noted however, that, as discussed in Ratfai (2000), the higher the inflation rate in a country, the lower the duration of prices tends to be.[35] Since the findings of Ratfai are reported for a period when the average inflation rate has been higher than the sample period of the current estimation, we have reason to assume, that the average duration of prices may have somewhat increased since Ratfai has established his results. Thus, the overestimation bias does not seem to be too big.

Moreover, the aforementioned problems concerning the estimation of fixed prices probability and of the importance of backward looking price setting are however by no means specific to Hungarian estimations. Similar remarks apply to estimations for developed economies. Note especially, that the slope

[34] Estimates reported by GGL (2001) for the same specification on euro area data are: Specification 1: $\omega = 0.024$ (0.122), $\xi = 0.907$ (0.015). Specification 2: $\omega = 0.335$ (0.129), $\xi = 0.922$ (0.031).

[35] For a state-contingent price setting model see also Dotsey, King, Wolman (1999). These authors arrive to similar conclusions.

coefficient in the Euro area hybrid Phillips curve, as reported by GGL (2001) for the same specification, is not significantly different from zero either[36] and that the average price duration implied by these specifications is also relatively high (10 to 12 quarters).[37] Hence, both Euro area and Hungarian estimations tend to overestimate price and inflation inertia. Since the estimations use the same technique the *relative* degree of price and inflation inertia in these regions is likely to be correctly indicated. Our estimations therefore seem to support that Hungarian inflation is significantly more sluggish than Euro area inflation. The major reason leading to Hungarian inflation inertia, seems to be pervasive backward looking behavior, while prices are more often reset than in the Euro area.

To check for the robustness of our results, equations (11) and (12) have been reestimated first, using the detrended CPI inflation rate, and second, using non-detrended variables. Results are summarized in rows 5 to 12 of Table 5. As can be seen, the use of these variables yields less promising results: some of the estimations were not convergent, while one (line 6) was converging to theoretically implausible estimates. This seems to be once more due to the insignificance of the real unit labor cost in these specifications. At the same time, the results of the remaining specifications are very similar to those of the benchmark estimations. Note especially that the coefficient estimate of $\lambda$ is significant in both the estimation of Specification 1 using detrended CPI inflation rate and the estimation of Specification 2 using non detrended variables. The estimates of the other coefficients are not statistically different from the estimates of the corresponding specification when using the detrended core inflation rate.

# 5 Open Economy New Hybrid Phillips Curves

In this section we extend the New Hybrid Phillips curve to settings for open economies. The relationship between the inflation rate and real exchange rate can be modeled in different ways depending on assumptions made on the type of imported goods and on the exchange rate pass-through. The first model we discuss considers imported goods as final consumption goods and assumes that changes in the exchange rate are fully passed through to the general price level. In contrast, in the second model, we specify imported goods as intermediate consumption goods. In this setting, there is full pass-through of changes in the exchange rate to import prices, the pass-through to the general price level is however incomplete.

[36] GGL (2001) report the following estimates. Specification 1: $\lambda = 0.018$ (0.012). Specification 2: $\lambda = 0.006$ (0.007). (See Table 2, lines 3 and 4 in their paper).

[37] It should be noted, that this is not the preferred specification of GGL (2001). These authors have indeed modified their model to allow for decreasing returns to scale, which yielded more realistic results. However, the same modification for Hungarian estimations did not change our conclusions.

## 5.1 Imported Final Consumption Goods

The description of this model is principally based on Kara, Nelson (2003). We extend their specification to the Hybrid Phillips curve setting.

Let us consider imported goods as final consumption goods. In addition, let us assume, that domestic price setters behavior can be captured by the Hybrid Phillips curve model. In this case, equation (8) describes the evolution of the domestic inflation rate, $\pi_t^d$:

$$\pi_t^d = \gamma_b \pi_{t-1}^d + \gamma_f E_t \pi_{t+1}^d + \lambda mc_t. \tag{13}$$

Following from households' optimal choice between domestic and imported goods, the overall inflation rate, $\pi_t$ can be expressed as a weighted average of domestic and imported goods inflation rate.[38] Assuming full pass-through, the overall inflation rate can be written as:

$$\pi_t = (1-s)\pi_t^d + s(\pi_t^f + \Delta e_t),$$

where $\pi_t^f$ is the inflation rate of imported prices in foreign currency, $\Delta e_t$ is the depreciation rate of the domestic currency and $s$ stands for the share of imported prices in the inflation rate of the general price level. Defining the real exchange rate as $q_t = p_t^f + e_t + p_t^d$ and rearranging we get:

$$\pi_t^d = \pi_t - s\Delta q_t, \tag{14}$$

where $\Delta q_t$ is the rate of change of the real exchange rate. Defined in this way a rise (decrease) in $q_t$ shows the real depreciation (appreciation) of the domestic currency. Restricting $\gamma_b + \gamma_f = 1$ and substituting expression (14) into the Phillips curve (13) yields the cumbersome expression:

$$\pi_t = \gamma_f E_t \pi_{t+1} + \gamma_b \pi_{t-1} - s\gamma_f (E_t \Delta q_{t+1} - \Delta q_t) + s\gamma_b (\Delta q_t - \Delta q_{t-1}) + \lambda mc_t. \tag{15}$$

The inflation rate hence depends on the current and the future expected change of the real depreciation rate. The sign of the current change of real depreciation rate is expected to be positive, while that of the future expected change in the real depreciation rate is expected to be negative.

The orthogonality condition that has been estimated takes the following form:

$$E_t\{(\pi_t - \gamma_f \pi_{t+1} - \gamma_b \pi_{t-1} + s\gamma_f(\Delta q_{t+1} - \Delta q_t) - s\gamma_b(\Delta q_t - \Delta q_{t-1}) - \lambda mc_t)\mathbf{z}_t\} = 0. \tag{16}$$

The 2SLS estimator with 3 lag Newey-West correction has been used once more. Note that the expected change of real depreciation is an additional endogenous

[38] This expression can be deduced assuming monopolistic competition at home and abroad with an equal elasticity of substitution between differentiated goods within a region and a CES utility function over domestic and imported goods. See e.g. Monacelli (2004).

variable, which had to be instrumented. This required the modification of our instrument set. The new instrument set includes: an additional lag of unit labor cost, two additional lags of budget deficit over GDP ratio while dropping the real GDP growth rate, the current deficit / GDP ratio and the lag of nonprocessed food price inflation. The partial $R^2$ for the inflation expectation is 0.213 and partial $R^2$ for the expectation of the real depreciation is 0.146.[39]

The first line in table 6 shows estimation results with detrended core inflation as dependant variable. The test of the overidentifying restrictions does not reject the specification ($J = 4.810$ with a p value of 0.683). Note, however, that the terms with the change of real depreciation rate are highly insignificant; moreover, the current change of real depreciation enters the equation with the incorrect negative sign. The results suggest that the open economy model assuming imported final consumption goods is a poor description of the Hungarian inflation dynamics. To compare, a purely forward looking version of equation (15) has been estimated by Kara, Nelson (2003b) on UK data. The results reported by Kara and Nelson are very similar to ours: they do not find evidence supporting this open economy model for the UK inflation.

## 5.2 Imported Intermediate Production goods

As an alternative, we follow McCallum, Nelson (1999) to model imported goods as intermediate production goods while all final goods are produced domestically. This specification will modify the Hybrid Phillips curve in the expression of real marginal cost. The real exchange rate will enter the new equation in levels and not in differences.

Formally the model can be deduced as follows. To keep things simple, we assume a Cobb-Douglas production technology:[40]

$$y_t = \alpha l_t + (1 - \alpha) y_t^m$$

where $y_t^m$ is an index of imported differentiated intermediate production goods. Assuming the price of one unit of the imported composite good is $p_t^f + e_t$, the real marginal cost can be expressed as:

$$mc_t = \alpha ulc_t + (1 - \alpha) q_t, \tag{17}$$

where $ulc_t$ stands for real unit labor cost and $q_t$ is the real cost of a unit of the imported good. Substituting this expression into the closed economy Hybrid Phillips curve (8) gives:

$$\pi_t = \gamma_f E_t \pi_{t+1} + \gamma_b \pi_{t-1} + \lambda^l ulc_t + \lambda^m q_t. \tag{18}$$

[39] Values of estimation using detrended data. While these values seem acceptable, the adjusted partial $R^2$ is negative in both cases. It was however impossible to find an instrument set which instruments both variables at the same time so that it remains reasonably parsimonious as well.

[40] Variables expressed in deviation from steady state.

The expressions linking $\gamma_b$ and $\gamma_f$ to the structural parameters are as in expression (9), $\phi$ remains unchanged as well. The coefficient of real unit labor cost, and real exchange rate are:

$$\begin{aligned}
\lambda^l &\equiv \frac{\alpha(1-\omega)(1-\xi)(1-\beta\xi)}{\phi}; \\
\lambda^m &\equiv \frac{(1-\alpha)(1-\omega)(1-\xi)(1-\beta\xi)}{\phi}.
\end{aligned}$$

#### 5.2.1 Reduced Form

As in the case of the closed economy model, we first perform 2SLS estimates of the model's reduced form. The corresponding orthogonality condition is:

$$E_t\{(\pi_t - \gamma_f \pi_{t+1} - \gamma_b \pi_{t-1} - \lambda^l ulc_t - \lambda^m q_t)\mathbf{z}_t\} = 0. \tag{19}$$

To estimate this orthogonality condition we use the benchmark instrument set.[41]

The results of the estimation of equation (19) using detrended variables are displayed in the first two lines of table 7. As can be seen, the results of the closed economy models' estimations remain relatively stable. The restriction of the inflation coefficients summing to unity can again not be rejected, with the lagged and future expected inflation rate receiving a roughly equal weight once more. The real exchange rate takes the expected sign, although its estimate is again not statistically different from zero. Interestingly, Kara, Nelson (2003b) again report similar findings with regard to the UK inflation. They similarly recover estimates of the real unit labor cost and of the real exchange rate with the expected sign but neither of these variables is found to be significant.

#### 5.2.2 Structural Form

Last, we reestimate the open economy Hybrid Phillips curve equation with the underlying structural parameters using the real effective exchange rate as a proxy for the real cost of imported goods. As in the closed economy framework, we estimate two alternative specifications according to the normalization of the current inflation coefficient.

Specification (1):

$$\begin{aligned}
&E_t\{(\phi\pi_t - \beta\xi\pi_{t+1} - \omega\pi_{t-1} - \alpha(1-\omega)(1-\xi)(1-\beta\xi)ulc_t \\
&\quad -(1-\alpha)(1-\omega)(1-\xi)(1-\beta\xi)q_t)\mathbf{z}_t\} = 0.
\end{aligned}$$

Specification (2):

$$\begin{aligned}
&E_t\{(\pi_t - \phi^{-1}\beta\xi\pi_{t+1} - \phi^{-1}\omega\pi_{t-1} - \phi^{-1}\alpha(1-\omega)(1-\xi)(1-\beta\xi)ulc_t \\
&\quad -\phi^{-1}(1-\alpha)(1-\omega)(1-\xi)(1-\beta\xi)q_t)\mathbf{z}_t\} = 0.
\end{aligned}$$

In addition to the closed economy structural parameters, this specification includes the additional parameter $\alpha$ of the production technology. We have first

[41] In the open economy model, the partial $R^2 = 0.44$ and adjusted partial $R^2 = 0.16$.

estimated the above specifications without restricting $\alpha$ to a particular value. The estimate of $\alpha$ has been around the value of 0.45 in all specifications.[42] However, since we are more interested in recovering the parameters of the price setting itself, and since restricting $\alpha$ to this particular value significantly reduced the standard error in our estimates, we only report results for estimations where $\alpha$ is calibrated to 0.45. The first two lines in table 8 show estimation results for $\beta$, $\xi$ and $\omega$ using detrended data. Lines 3 and 4 restrict $\beta = 1$ and estimate only $\xi$ and $\omega$. As can be seen in table 8.a, this restriction cannot be rejected at any usual significance level. The restriction increases at the same time the stability of our estimations, we hence prefer this specification to the unrestricted one.

The estimation of the open economy specification supports previous results. The importance of backward looking behavior is confirmed to be relatively high and the probability of fixed prices relatively low compared to Euro area estimates. According to our preferred specification, the fraction of firms which set prices according to a backward looking rule-of-thumb lies between 1/3 and 1/2. The probability in any period that a firm cannot reset its price is indicated to be between 40 to 50 percent. This corresponds to an average duration of fixed prices of 1.6 - 2.1 quarters. Note that the open economy estimates allow us to refine the interval of $\omega$ compared to the closed economy estimates. At the same time, the estimate of $\xi$ is indicated to be somewhat lower in the open economy model than those implied in the closed economy model; the open economy estimates are also closer to microeconomic evidence as reported by Ratfai (2000).

These improvements might indicate, that the open economy proxy for real marginal cost, as expressed by equation (17) does a better job, than the simple real unit labor cost. Note, that both real unit labor cost and the real exchange rate have the expected signs in both specifications and that the estimates of $\lambda^l$ and of $\lambda^m$ in specification (1) are at more than one standard error distance from zero. Overall the results seem encouraging in favor of this open economy model.

# 6 Robustness

In this subsection, we check the robustness of our estimates in two ways. First, we test the stability of our parameter estimates with regard to structural breaks. Second, we show how the choice of different instrument sets influences our results.

## 6.1 Structural stability

As mentioned earlier, Linde (2002) criticized the GMM estimation method claiming that the coefficient estimates were sensitive to switches in the monetary policy. We therefore check the stability of our estimates to the considered

[42] Formal restriction tests do not reject the $H_0$ of $\alpha = 0.45$ in any of our estimates.

sample period. Doing so, we perform recursive estimations. Since our conclusions are not affected by the particular specifications, we restrict our discussion to the outcome of the stability check for our preferred specification, i.e. the open economy model with imported goods specified as intermediate production goods. The terminal date goes from $T = 2000Q1$ to $2004Q1$ as in recursive estimates of the traditional Phillips curve. Figure 6a shows coefficient estimates of the structural parameters $\omega$ and $\xi$ for specification (1), figure 6b displays those for specification (2). The upper line in each figure displays results of unrestricted estimations, the bottom line those of the restricted ones.

The coefficient estimates implied by specification (1) appear to be very stable both in the unrestricted and in the restricted estimations: all estimates remain well within the one standard error band. Note especially, that estimates on the subsample of the exchange rate targeting regime, i.e. with $T = 2001Q2$, are very close to the estimates of the full sample: while over the subsample, $\omega$ is estimated to be 0.32, its value is estimated to be 0.34 over the full sample; by the same token, the subsample estimate of $\xi$ is 0.38, the full sample estimate is 0.4.

The estimation of specification (2) also appears to be relatively stable apart from a couple of terminal dates around the year 2000, where the estimates seem to converge to implausible values. Note that estimates on the entire subsample of the exchange rate targeting regime ($T = 2001Q2$), are relatively close to the full sample estimates in this specification, too. Although the point estimates of both $\omega$ and $\xi$ decrease from the subsample to the full sample by about 0.1, the standard errors are also relatively large, so that these differences remain within the error band of the estimations. We hence conclude that our estimates are reasonably stable.

## 6.2 Instrument sets

As already discussed, the appropriate choice of the instrument set is essential in instrumental variable estimations. So far, we have described our results for estimations using always the same set of instruments. At this place, we discuss the way in which the use of different instrument sets influences our estimation results.

New instrument sets either include additional variables or additional lags of already used variables. Each instrument set is described in Appendix A with the respective $R^2$ and partial $R^2$ indicators. The benchmark instrument set is Set 4. The partial $R^2s$ of alternative instrument sets lie in the range of 0.15 (for Set 2) to 0.49 (for Set 1). As a reminder, partial $R^2$ of the benchmark instrument set 4 is equal to 0.56.

Our conclusions were once more independent of the particular specification estimated. Figure 7 shows estimation results of Specification (1) of the open economy model with imported intermediate production goods with the time discount factor restricted to 1. In each panel, we shaded the area between the minimum and the maximum value of the point estimates.

As can be seen in the figure, neither estimates of the deep structural parameters, nor the implied reduced form parameters are too much influenced by the choice of the instrument set. In particular, all estimates of each coefficient lie within one standard error distance from other estimates of the same coefficient. Note, that the point estimates of $\omega$ and $\gamma^f$ implied by the benchmark instrument set, lie within the range of estimates with different instrument sets. At the same time, the estimates of $\xi$ are at the lower extreme of estimates using different instrument sets, while those of the slope coefficients $\lambda^l$ and $\lambda^m$ lie at the upper edge. This might again be linked to the aforementioned overestimation problem in the case of weaker instrument sets: hence, when the coefficients of real marginal cost are low, this will automatically imply high values of the price rigidity coefficients. Thus, the relatively low value of $\xi$ implied by estimates using the best performing instrument set may well be closer to reality than other estimates.

Overall, the main results, namely the relatively high degree of inflation inertia, the relatively big size of backward looking firms and the relatively low average price duration, are not altered by the use of any other instrument set.

# 7 Conclusion

In this paper, we estimated different Phillips curve models to describe Hungarian inflation dynamics over the period 1995Q1 to 2004Q1. Our results suggest, that, while estimates of traditional Phillips curve are subject to the Lucas critique, an open economy extension of the New Hybrid Phillips curve, specifying imported goods as intermediate production goods, can reasonably well describe Hungarian inflation dynamics. We use this model to compare Hungarian and Euro area inflation dynamics and find that inflation in Hungary is significantly more inertial than Euro area inflation. The source of Hungarian inflation is indicated to lie in substantial backward looking price setting, while prices seem to be reset more often than in the Euro area.

It should be noted that New Hybrid Phillips curves, explaining inflation inertia by assuming some backward looking price setting behavior, are deduced from models based on the assumption of rational expectations. As shown in Roberts (1997) however, the same reduced form of the Phillips curve can be deduced assuming pure forward looking price setting, where inflation persistence is explained by agents' not-fully rational expectations. Estimating the New Hybrid Phillips curve we therefore exclude by assumption a possible source of inflation persistence: adaptive expectations. This may then lead to an overestimation of the importance of the other potential source of inflation inertia, embodied in the New Hybrid Phillips curve: backward looking price setting behavior.

In addition, estimates for Hungary are based on data which stem from a period of transition and convergence characterized by fundamental institutional and structural economic changes. Under these circumstances, the evolution of key economic indicators can hardly be considered as being purely cyclical variations around a constant trend. Instead, transition and convergence have arguably affected economic indicators in a way which is not captured by standard economic models. In our estimations, we followed the approach to consider all the influence of the convergence process to have been captured in the series' trends. Based on this simplifying assumption, we have applied standard Phillips curves to detrended Hungarian data. This approach is also adopted e.g. by Coenen, Wieland (2000) in the estimation of a small-scale Euro area model. At the same time, this approach does not take into account the potential impact of the convergence process on business cycle dynamics, which might be relevant in some aspects. Indeed, as discussed in Coenen, Wieland (2000), inflation dynamics appears to exhibit a higher degree of persistence in countries experiencing convergence processes. This might suggest that some characteristics of the convergence process might have an influence on short-run inflation dynamics as well. Investigating the potential ways in which this influence is exerted should be the topic of future research.

# 8 Appendix A: Alternative Instruments

Set 1
instruments: constant corehp{2 to 3} ulchp{1 to 3} rpmhp{0 to 3}
$R^2 = 0.61385$
Partial $R^2 = 0.49178$

Set 2
instruments: constant corehp{1} ulchp{0} yhp{1} rpmhp{0} y_growth{0} dneerhp{1} dnpfhp{0 to 1}
$R^2 = 0.65004$
Partial $R^2 = 0.14602$

Set 3
instruments: constant corehp{1 to 2} yhp{0} ulchp{0} dq{2} dnpfhp{0 to 2}
$R^2 = 0.70528$
Partial $R^2 = 0.28109$

Set 4
instruments: constant corehp{1 to 2} ulchp{0 1} winflhp{0} budhp{0 to 1} y_growth{0} dnpfhp{0 to 1}
$R^2 = 0.70016$
Partial $R^2 = 0.56494$

Set 5
instruments: constant corehp{1 to 4} ulchp{0 to 4}
$R^2 = 0.40313$
Partial $R^2 = 0.23708$

Set 6
instruments: constant corehp{1 2} ulchp{0 to 2} rpmhp{1 to 2} dq{1 to 2} budhp{1 to 2} u_rate{1 to 2} y_growth{1 2}
$R^2 = 0.6074$
Partial $R^2 = 0.4666$

Note: number of lags between brackets.
Abbreviations:
...hp = deviation from HP trend
budhp = budget deficit / GDP
corehp = core inflation
dneerhp = change in nominal exchange rate
dnpfhp = change in non processed food prices
dq = change in real exchange rate (not detrended)
rpmhp = real import price (computed as nominal import price/CPI)

ulchp = real unit labor cost
u_rate = unemployment rate (not detrended)
winflhp = wage inflation
yhp = output gap
y_growth = growth rate of real GDP (not detrended)

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# TABLES

Table 1: Traditional Phillips curve for closed economy

| CLOSED | Specification | $const$ | $\pi_{t-1}$ | $x_{t-1}$ | $\bar{R}^2$ |
|---|---|---|---|---|---|
| Detrended | ulc | 0.0226<br>(0.2371) | 0.5523<br>(0.0967) | 0.1754<br>(0.0801) | 0.43 |
| | output gap | −0.0024<br>(0.2285) | 0.6291<br>(0.1466) | 0.2312<br>(0.2601) | 0.44 |
| Non detrended | ulc | 1.6075<br>(0.6796) | 0.8209<br>(0.0520) | 0.0944<br>(0.0526) | 0.95 |
| | output gap | 0.6986<br>(0.4869) | 0.8929<br>(0.0385) | 0.0221<br>(0.2612) | 0.95 |

Note: OLS estimation with Newey-West correction for serial correlation including 3 lags. Standard errors are shown in brackets.

Table 2: Traditional Phillips curve for open economy (full sample)

| OPEN | Specification | $const$ | $\pi_{t-1}$ | $open_{t-1}$ | $ulc_{t-1}$ | $\bar{R}^2$ |
|---|---|---|---|---|---|---|
| Detrended | $dq_{t-1}$ | 0.1463<br>(0.3009) | 0.5607<br>(0.1050) | 0.0332<br>(0.0417) | 0.1744<br>(0.0753) | 0.43 |
| | $drpm_{t-1}$ | 0.0099<br>(0.2599) | 0.5537<br>(0.0925) | −0.0034<br>(0.0319) | 0.1750<br>(0.0788) | 0.41 |
| | $dtot_{t-1}$ | 0.0222<br>(0.2407) | 0.5526<br>(0.1010) | 0.0012<br>(0.0589) | 0.1750<br>(0.0923) | 0.41 |
| Non detrended | $dq_{t-1}$ | 1.8342<br>(0.5808) | 0.8163<br>(0.0471) | 0.0473<br>(0.0363) | 0.0917<br>(0.0540) | 0.95 |
| | $drpm_{t-1}$ | 1.6283<br>(0.6254) | 0.8200<br>(0.0487) | 0.0027<br>(0.0325) | 0.0947<br>(0.0519) | 0.95 |
| | $dtot_{t-1}$ | 1.5964<br>(0.7137) | 0.8215<br>(0.0543) | 0.0096<br>(0.0595) | 0.0931<br>(0.0584) | 0.95 |

Note: OLS estimation with Newey-West correction for serial correlation. Lag selection for explanatory variables based on BIC. Standard errors are shown in brackets.

Table 2a: Traditional Phillips curve - Exclusion tests for external variables

| OPEN | Specification | $lags$ | $\varkappa^2$ | $p(\varkappa^2)$ |
|---|---|---|---|---|
| Detrended | $dq_{t-i}$ | 5 | 0.699 | 0.403 |
| | $drpm_{t-i}$ | 3 | 5.828 | 0.016 |
| | $dtot_{t-i}$ | 1 | 0.0004 | 0.984 |
| Non detrended | $dq_{t-i}$ | 5 | 35.759 | 0.000 |
| | $drpm_{t-i}$ | 6 | 132.365 | 0.000 |
| | $dtot_{t-i}$ | 1 | 0.022 | 0.882 |

Note: Lag selection based on AIC. The $H_0$ of the test is $\sum_{i=1}^{h} \gamma_i = 0$. The test statistic follows a $\varkappa^2$ distribution.

Table 3: Traditional Phillips curve - Subsample estimations

| Specification | sample | $const$ | $\pi_{t-1}$ | $open_{t-1}$ | $ulc_{t-1}$ | $\bar{R}^2$ |
|---|---|---|---|---|---|---|
| CLOSED | full sample | 0.0226 (0.2371) | 0.5523 (0.0967) | | 0.1754 (0.0801) | 0.43 |
| | pre-2001$Q$2 | 0.1330 (0.2764) | 0.4810 (0.1061) | | 0.2590 (0.0785) | 0.50 |
| OPEN with: | | | | | | |
| $dq_{t-1}$ | full sample | 0.1463 (0.3009) | 0.5607 (0.1050) | 0.0332 (0.0417) | 0.1744 (0.0753) | 0.43 |
| | pre-2001$Q$2 | 0.0398 (0.3227) | 0.4536 (0.1117) | −0.0453 (0.0396) | 0.2888 (0.0815) | 0.50 |
| $drpm_{t-1}$ | full sample | 0.0099 (0.2599) | 0.5537 (0.0925) | −0.0034 (0.0319) | 0.1750 (0.0788) | 0.41 |
| | pre-2001$Q$2 | 0.0824 (0.2987) | 0.4829 (0.0917) | −0.0367 (0.0303) | 0.2836 (0.0773) | 0.51 |
| $dtot_{t-1}$ | full sample | 0.0222 (0.2407) | 0.5526 (0.1010) | 0.0012 (0.0589) | 0.1750 (0.0923) | 0.41 |
| | pre-2001$Q$2 | 0.1734 (0.2808) | 0.4666 (0.1038) | −0.0400 (0.0520) | 0.2823 (0.0831) | 0.50 |

Note: OLS.estimator with Newey West correction for serial correlation. Standard errors are shown in brackets.

Table 4: Estimates of reduced form New Hybrid Phillips Curve

| Specification | $\gamma_f$ | $\gamma_b$ | $\lambda$ | $J$ | $p(J)$ | $H_0 : \beta = 1$ | $p(\varkappa^2)$ |
|---|---|---|---|---|---|---|---|
| **CoreHP** | | | | | | | |
| UR | 0.536 (0.073) | 0.470 (0.079) | 0.054 (0.084) | 5.973 | 0.650 | 0.002 | 0.960 |
| R | 0.533 (0.048) | 0.467 (0.048) | 0.056 (0.071) | | | | |
| **CPIHP** | | | | | | | |
| UR | 0.567 (0.085) | 0.466 (0.059) | −0.010 (0.065) | 6.567 | 0.584 | 0.078 | 0.778 |
| R | 0.546 (0.041) | 0.454 (0.041) | 0.003 (0.046) | | | | |
| **Core** | | | | | | | |
| UR | 0.662 (0.088) | 0.361 (0.078) | 0.060 (0.023) | 5.141 | 0.742 | 2.392 | 0.122 |
| R | 0.561 (0.059) | 0.439 (0.059) | 0.057 (0.023) | | | | |
| **CoreHP,** $y_t$ | | | | | | | |
| UR | 0.565 (0.063) | 0.475 (0.080) | 0.083 (0.176) | 5.685 | 0.771 | 0.128 | 0.720 |
| R | 0.549 (0.045) | 0.451 (0.045) | 0.119 (0.144) | | | | |

Note 1: CoreHP = detrended core inflation; CPIHP = detrended CPI inflation; Core = non-detrended core inflation; UR = unrestricted; R = restricted, referring to the restriction of $\gamma_f + \gamma_b = 1$. Unless otherwise indicated, the cyclical indicator is real unit labor cost.

Note 2: 2SLS estimator with Newey-West correction.(3 lags). Standard errors are shown in brackets.

Table 5: Estimates of New Hybrid Phillips curve - Structural form

| Specification | $\beta$ | $\omega$ | $\xi$ | $\gamma_f$ | $\gamma_b$ | $\lambda$ | $J$ | $p(J)$ | $H_0: \beta = 1$ $\varkappa^2$ | $p(\varkappa^2)$ |
|---|---|---|---|---|---|---|---|---|---|---|
| **CoreHP** | | | | | | | | | | |
| UR (1) | 0.691 (0.199) | 0.282 (0.106) | 0.545 (0.106) | 0.483 (0.086) | 0.362 (0.156) | 0.262 (0.223) | 7.235 | 0.512 | 2.404 | 0.121 |
| UR (2) | 1.025 (0.528) | 0.553 (0.233) | 0.615 (0.199) | 0.536 (0.174) | 0.470 (0.166) | 0.054 (0.159) | 5.973 | 0.650 | 0.002 | 0.962 |
| R (1) | 1 | 0.308 (0.090) | 0.481 (0.105) | 0.610 (0.072) | 0.390 (0.072) | 0.236 (0.156) | 7.135 | 0.623 | | |
| R (2) | 1 | 0.547 (0.141) | 0.621 (0.157) | 0.532 (0.053) | 0.468 (0.053) | 0.056 (0.072) | 6.054 | 0.735 | | |
| **CPIHP** | | | | | | | | | | |
| UR(1) | 0.862 (0.108) | 0.342 (0.079) | 0.569 (0.074) | 0.555 (0.076) | 0.387 (0.093) | 0.163 (0.069) | 44.327 | 0.000 | 1.613 | 0.204 |
| UR (2) | 2.637 (8.084) | 0.915 (3.189) | 0.422 (0.661) | 0.565 (1.935) | 0.465 (1.002) | −0.003 (0.103) | 6.779 | 0.561 | 0.041 | 0.840 |
| R (1) | 1 | 0.357 (0.054) | 0.533 (0.027) | 0.599 (0.043) | 0.401 (0.043) | 0.158 (0.024) | 7.333 | 0.603 | | |
| R (2) | Non convergent | | | | | | | | | |
| **Core** | | | | | | | | | | |
| UR (1) | Non convergent | | | | | | | | | |
| UR (2) | 1.048 (0.038) | 0.518 (0.090) | 0.678 (0.073) | 0.586 (0.036) | 0.427 (0.034) | 0.037 (0.026) | 5.082 | 0.749 | 1.622 | 0.203 |
| R (1) | Non convergent | | | | | | | | | |
| R (2) | 1 | 0.593 (0.108) | 0.677 (0.095) | 0.533 (0.049) | 0.467 (0.049) | 0.034 (0.028) | 4.896 | 0.843 | | |

Note: Abbreviations like in table 4. (1) resp. (2) refer to the estimated specification. Non-linear 2SLS estimator with 3 lag Newey-West correction for standard errors.

Table 6: Open economy: imported final consumption goods. Reduced form.

| Core inflation | $E_t\pi_{t+1}$ | $\pi_{t-1}$ | $(E_t\Delta q_{t+1} - \Delta q_t)$ | $(\Delta q_t - \Delta q_{t-1})$ | $mc_t$ | J* | p(J) |
|---|---|---|---|---|---|---|---|
| detrended | 0.614<br>(0.058) | 0.386<br>(0.058) | −0.004<br>(0.026) | −0.010<br>(0.039) | 0.037<br>(0.057) | 4.810 | 0.683 |
| not detrended | 0.586<br>(0.056) | 0.414<br>(0.055) | 0.002<br>(0.024) | −0.022<br>(0.040) | 0.045<br>(0.026) | 3.847 | 0.797 |

Note: 2SLS estimator with Newey-West correction. Standard errors are shown in brackets. The J statistic displayed is the one of the estimation not restricting $\gamma_f + \gamma_b = 1$.

Table 7: Open economy New Hybrid Phillips curve - Imported intermediate goods; Reduced form

| Specification | | | | | | | $H_0 : \beta = 1$ | |
|---|---|---|---|---|---|---|---|---|
| | $\gamma_f$ | $\gamma_b$ | $\lambda^l$ | $\lambda^m$ | $J$ | $p(J)$ | $\varkappa^2$ | $p(\varkappa^2)$ |
| **Core, detrended** | | | | | | | | |
| Unrestricted | 0.495<br>(0.104) | 0.481<br>(0.072) | 0.052<br>(0.010) | 0.071<br>(0.085) | 5.973 | 0.650 | 0.032 | 0.859 |
| Restricted | 0.510<br>(0.055) | 0.490<br>(0.055) | 0.041<br>(0.075) | 0.061<br>(0.065) | | | | |
| **Core, not detrended** | | | | | | | | |
| Unrestricted | 0.593<br>(0.050) | 0.418<br>(0.048) | 0.031<br>(0.025) | 0.010<br>(0.011) | 4.192 | 0.757 | 1.358 | 0.244 |
| Restricted | 0.572<br>(0.047) | 0.428<br>(0.047) | 0.030<br>(0.025) | 0.010<br>(0.011) | | | | |

Note: 2SLS estimator with 3 lag Newey-West correction. Standard errors are shown in brackets.

Table 8: Open economy New Hybrid Phillips curve - Imported intermediate goods; Structural form

| Specification | $\beta$ | $\omega$ | $\xi$ | $\gamma_f$ | $\gamma_b$ | $\lambda^l$ | $\lambda^m$ | $J$ | $p(J)$ |
|---|---|---|---|---|---|---|---|---|---|
| **CoreHP** | | | | | | | | | |
| UR (1) | 0.595 (0.268) | 0.304 (0.099) | 0.452 (0.107) | 0.384 (0.010) | 0.435 (0.215) | 0.178 (0.179) | 0.219 (0.219) | 5.805 | 0.669 |
| UR (2) | 0.909 (0.415) | 0.477 (0.189) | 0.539 (0.147) | 0.494 (0.132) | 0.480 (0.159) | 0.056 (0.114) | 0.068 (0.139) | 4.808 | 0.778 |
| R (1) | 1 | 0.337 (0.085) | 0.383 (0.110) | 0.532 (0.087) | 0.468 (0.087) | 0.158 (0.094) | 0.193 (0.115) | 5.637 | 0.776 |
| R (2) | 1 | 0.496 (0.126) | 0.522 (0.155) | 0.513 (0.057) | 0.487 (0.057) | 0.051 (0.055) | 0.062 (0.067) | 4.842 | 0.848 |
| **Core** | | | | | | | | | |
| UR(1) | 0.506 (0.373) | 0.369 (0.104) | 0.332 (0.121) | 0.262 (0.102) | 0.577 (0.257) | 0.247 (0.330) | 0.301 (0.403) | 5.591 | 0.693 |
| UR (2) | 1.035 (0.038) | 0.525 (0.104) | 0.712 (0.058) | 0.590 (0.040) | 0.420 (0.039) | 0.013 (0.011) | 0.016 (0.014) | 4.597 | 0.800 |
| R (1) | 1 | 0.331 (0.122) | 0.669 (0.045) | 0.669 (0.078) | 0.331 (0.078) | 0.033 (0.017) | 0.040 (0.021) | 5.625 | 0.777 |
| R (2) | 1 | 0.572 (0.105) | 0.705 (0.064) | 0.552 (0.050) | 0.448 (0.050) | 0.013 (0.008) | 0.016 (0.010) | 4.541 | 0.872 |

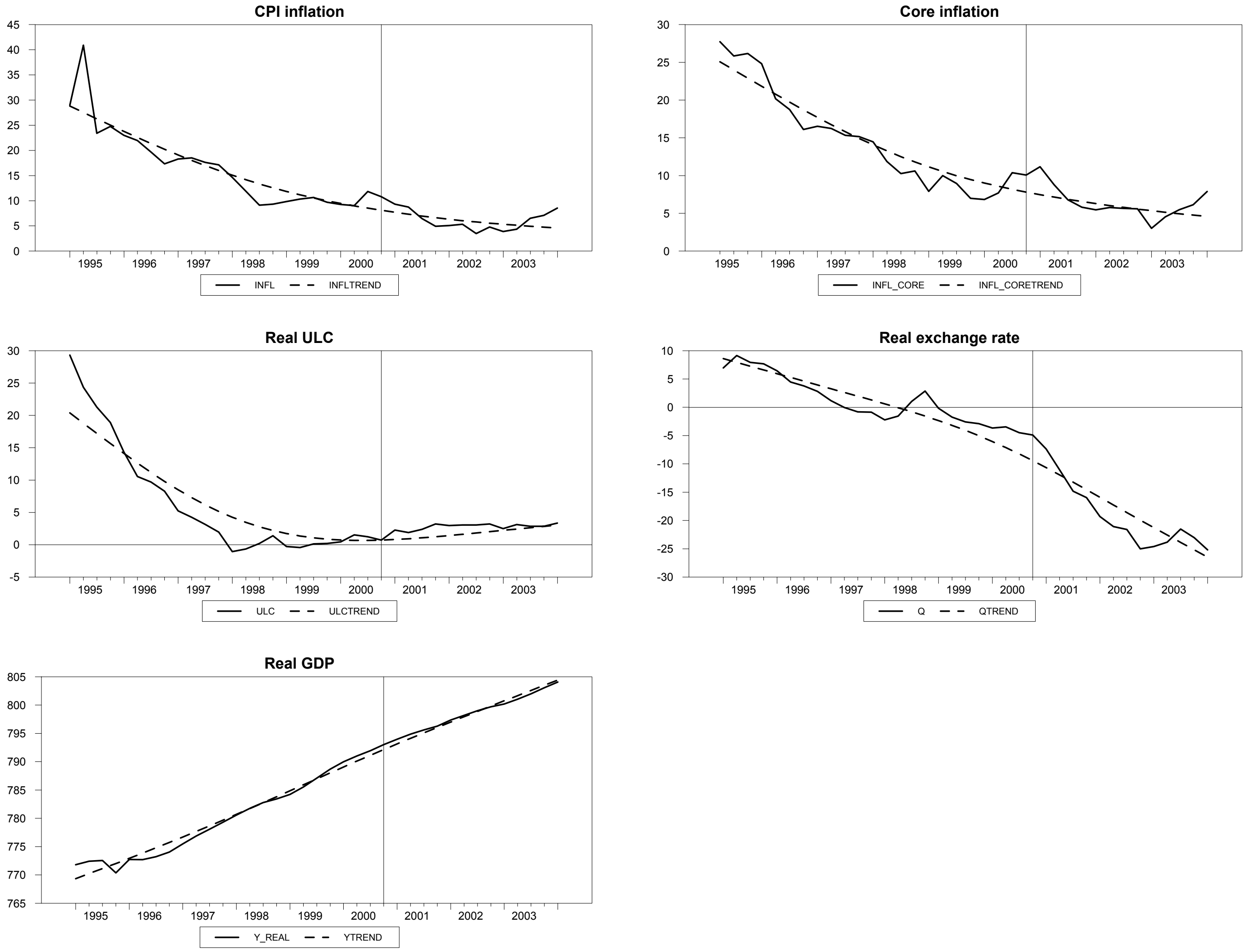
Note: Notations like previously. Non-linear 2SLS estimator with 3 lag Newey-West correction. Results of estimations with $\alpha = 0.45$.

Table 8.a: Open economy New Hybrid Phillips curve - structural form. Test results of restriction $\beta = 1$.

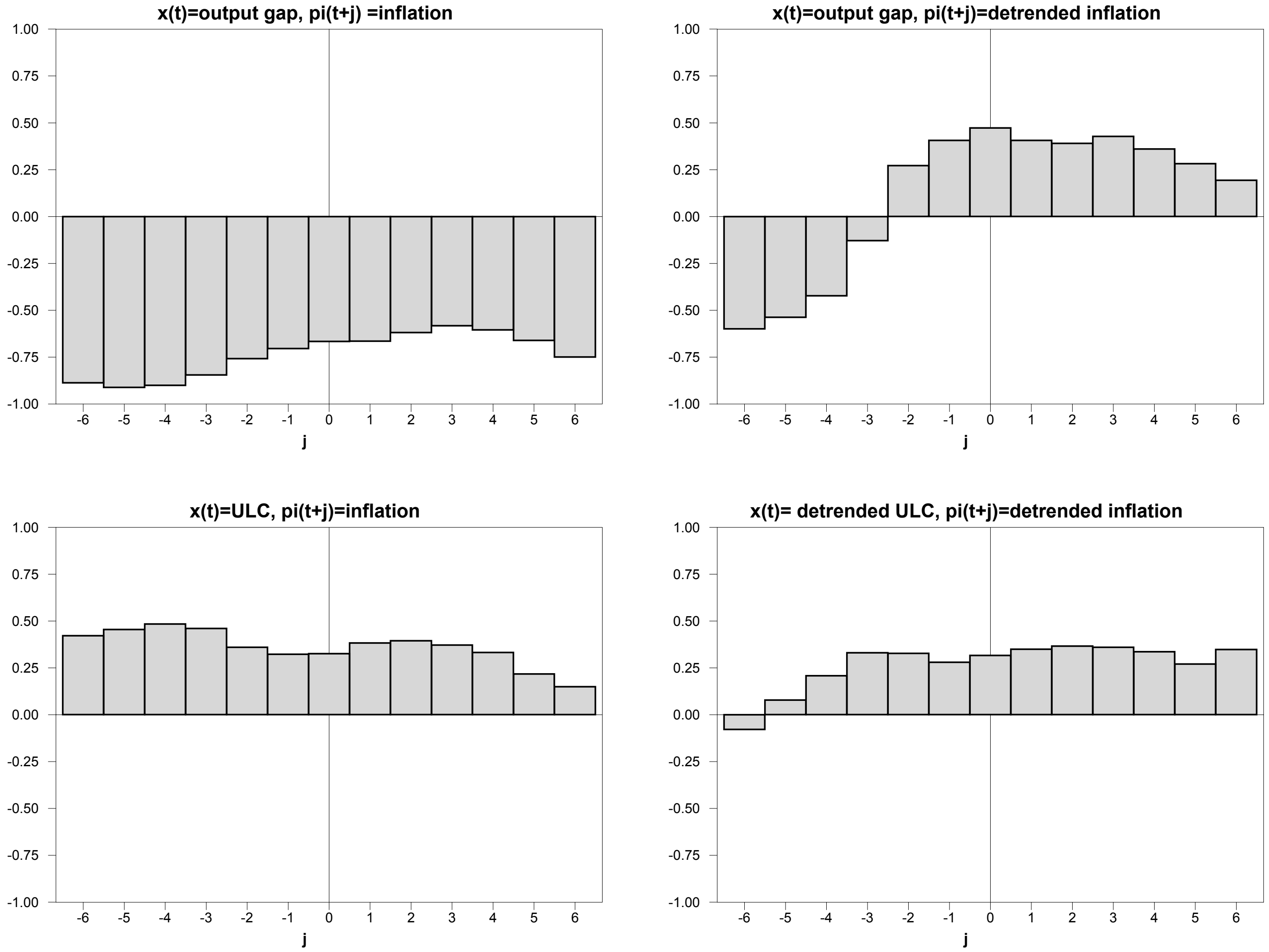
| Specification | $H_0 : \beta = 1$ | |
|---|---|---|
| | $\varkappa^2$ | $p(\varkappa^2)$ |
| **CoreHP** | | |
| UR (1) | 2.272 | 0.132 |
| UR (2) | 0.048 | 0.827 |
| **Core** | | |
| UR (1) | 1.755 | 0.185 |
| UR (2) | 0.857 | 0.355 |

Note: Complementary to table 8.

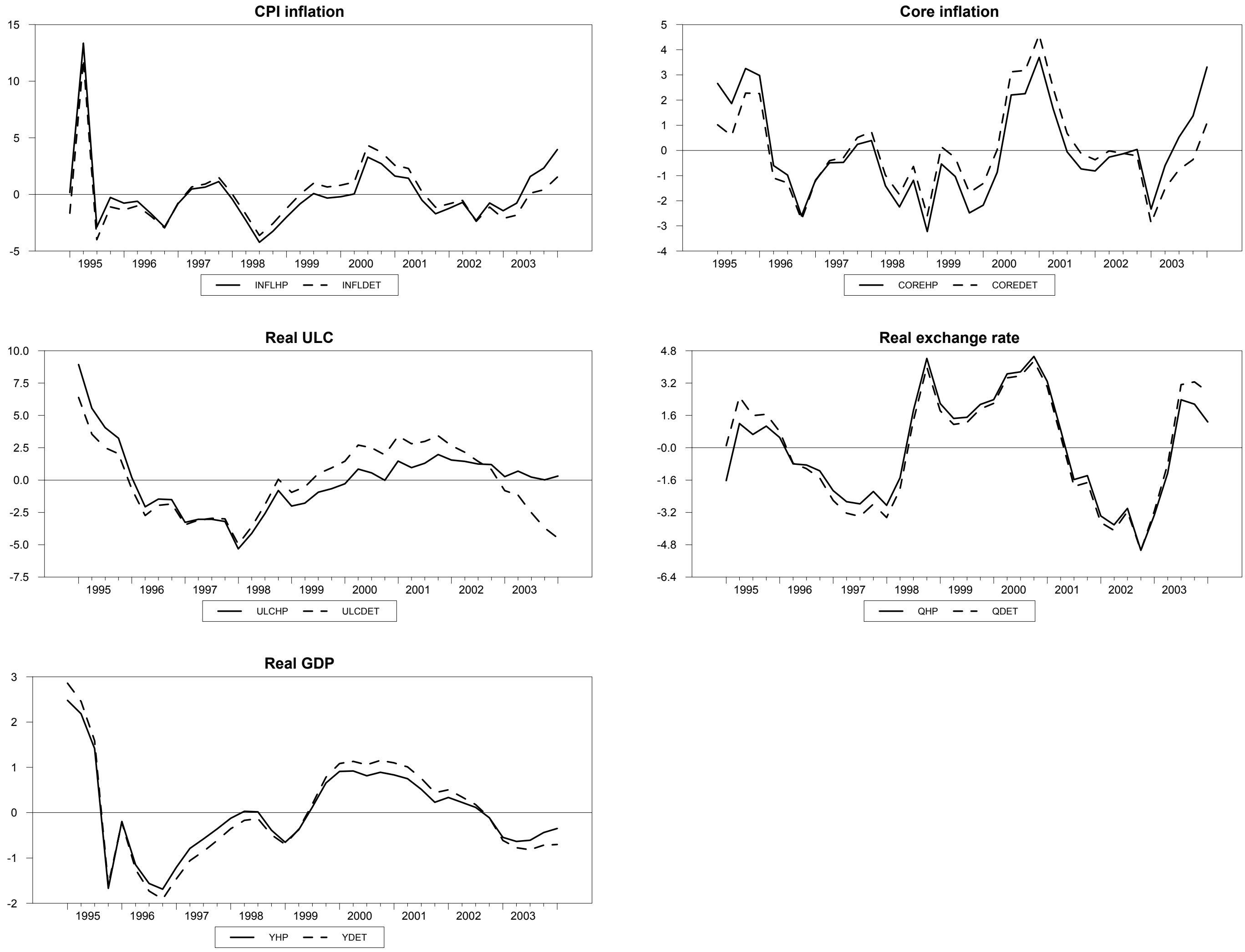
# Figure 1: Hungarian Data



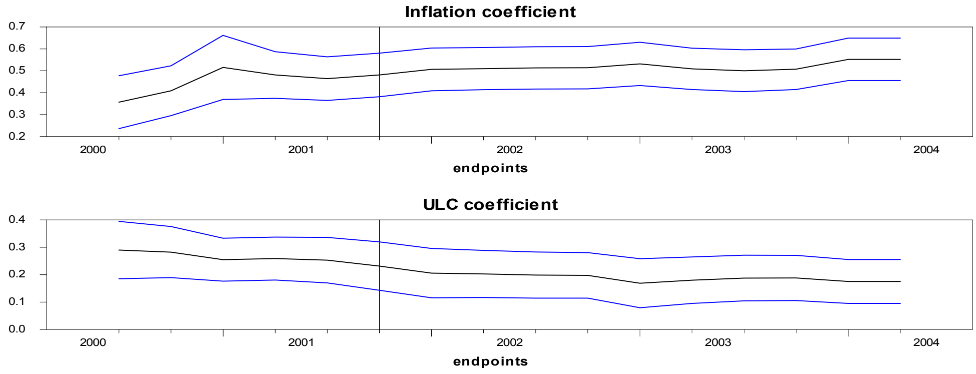
# Figure 2: Corr(x(t),pi(t+j))



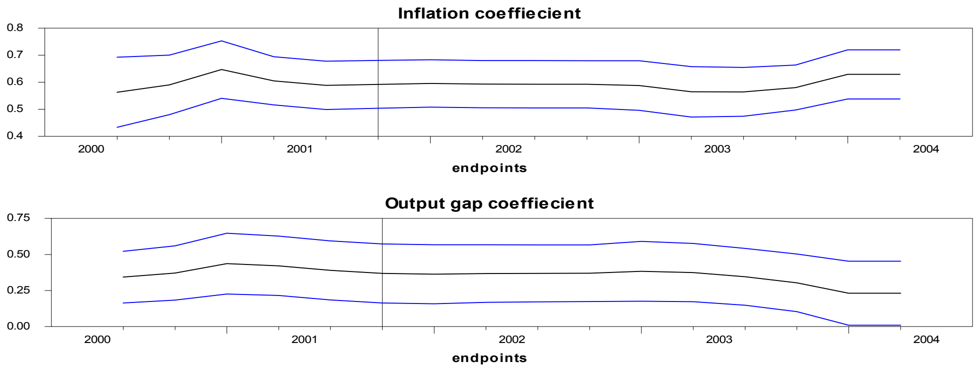
# Figure 3: Detrended data - HP filter vs. Deterministic Quadratic Trend



## Figure 4a: Recursive estimates - Closed economy with ULC




## Figure 4.b: Recursive estimates - Closed economy with output gap




## Figure 4.c: Recursive estimates - Open economy

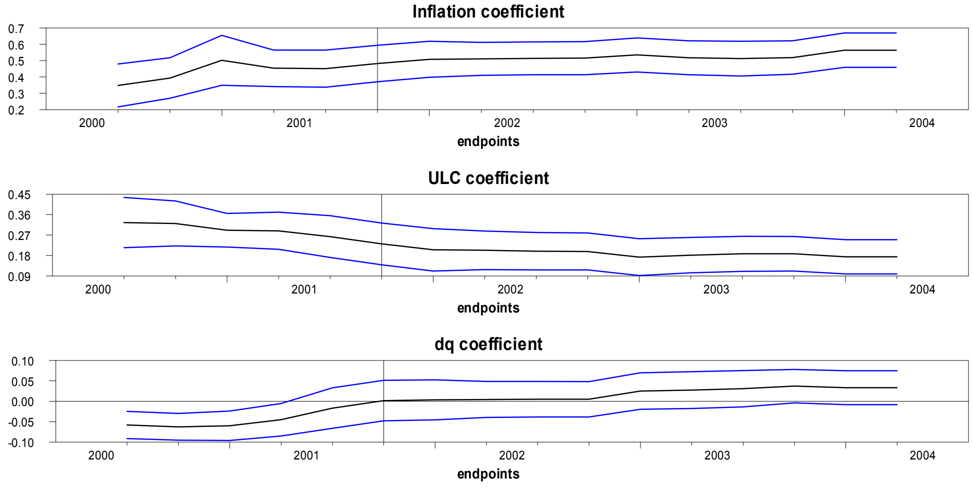
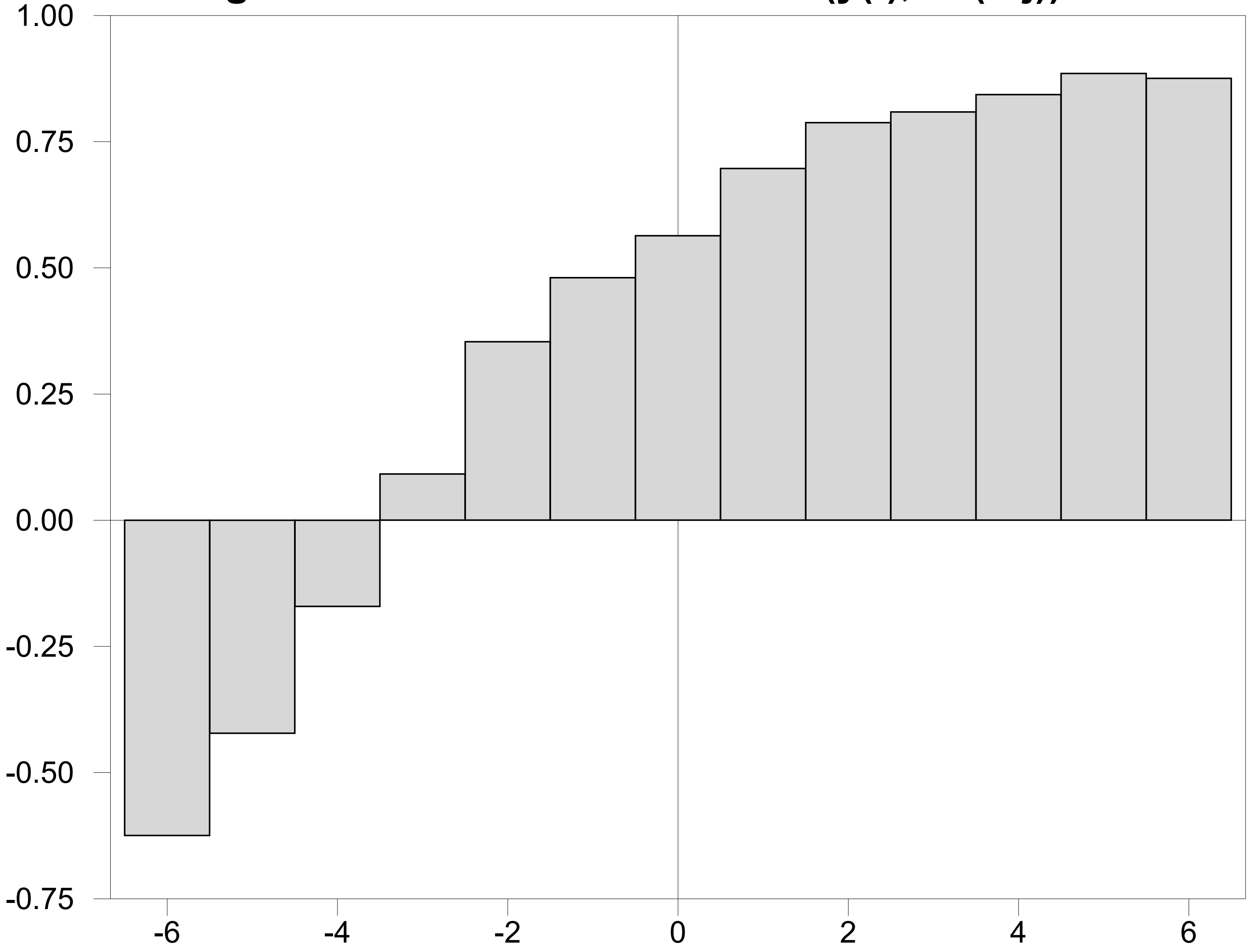
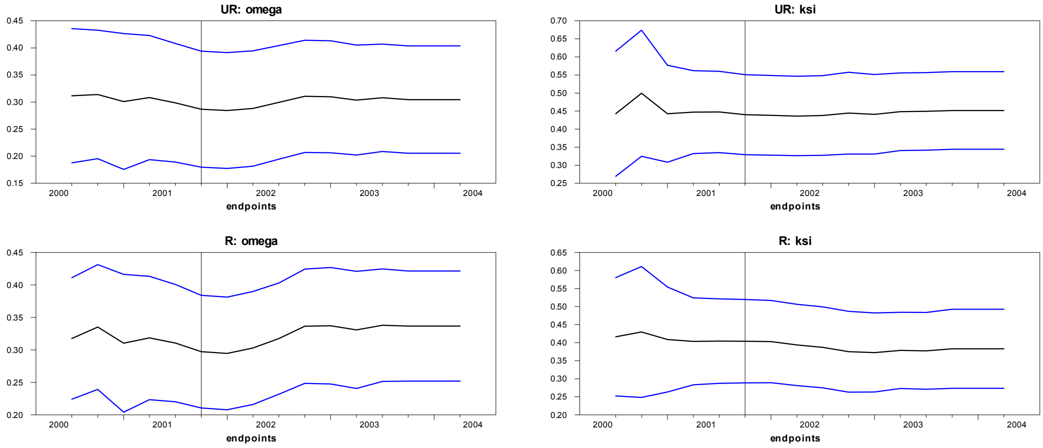


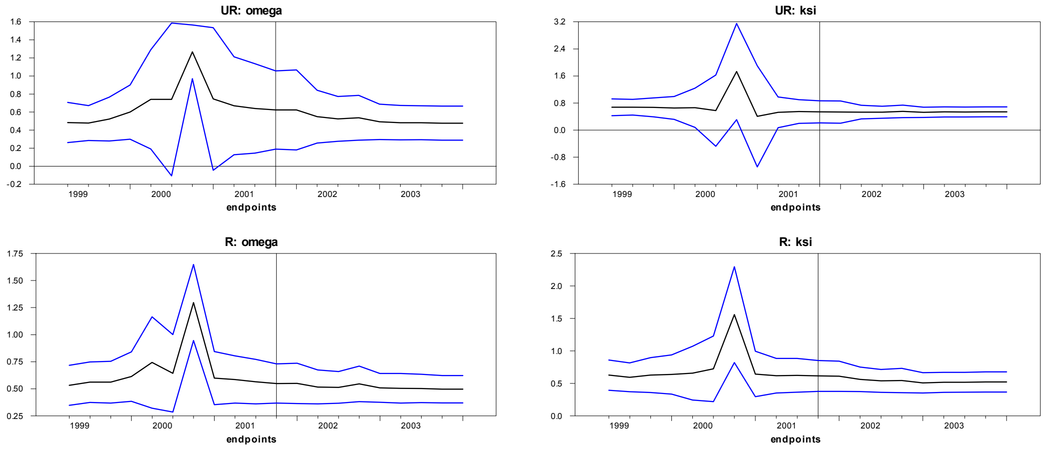
**Figure 5: Detrended series: Corr(y(t),ulc(t+j))**



## Figure 6a: Recursive estimates: Open economy - Specification (1)




## Figure 6b: Recursive estimates: Open economy - Specification (2)



# Figure 7: Check Instrument Set Stability