Modelling Inflation in EU Accession Countries: The Case of the Czech Republic, Hungary and Poland

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The Eastward Enlargement of the Eurozone

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Abstract

Inflation in Central and East European countries varied considerably over the transition phase, and econometric relationships between prices, money, wages and exchange rates are said to have been unstable during this period. In order to shed some light on the issue, this paper analyses some empirical models of the inflation process in the three earliest east European transition economies: the Czech Republic, Hungary and Poland. Since the end of the 1980s these economies have experienced high rates of inflation, although significant disinflation measures were introduced during the mid-nineties to enhance these countries' chances of joining the EU, and they succeeded in getting inflation under control without high costs in terms of lost output. Given this, the determinants of inflation need to be empirically analysed not only in order to understand the disinflation measures, but also to assess the possible effects of future pressure on prices. Price stabilisation is an essential complement to the success of transition. Policies to contain inflation are necessary for transition economies to grow and firms to restructure. In the present paper, we first look at inflation within the context of multivariate cointegration, where domestic and foreign price determinants are initially assessed in separate blocks (each single-theory based) in order to obtain a number of long-term attractors. We then formulate consumer and producer inflation equations from more general VEqCMs for each country. The importance of theory-based imbalances (from previous cointegration experiments) in explaining inflation can be assessed at this stage. Our most significant empirical findings seem to substantiate the idea that many, if not all, theoretical determinants of inflation are of importance in those countries in question: the exchange rate and the output gap would appear to be of particular importance in explaining the phenomenon.

JEL–Classification: C4, E5

Keywords : Inflation modelling, transition economies, European Union enlargement

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Modelling inflation in EU accession countries: the case of the Czech Republic, Hungary and Poland [1]

by Roberto Golinelli and Renzo Orsi[2]

Summary. Inflation in Central and East European countries varied considerably over the transition phase, and econometric relationships between prices, money, wages and exchange rates are said to have been unstable during this period. In order to shed some light on the issue, this paper analyses some empirical models of the inflation process in the three earliest east European transition economies: the Czech Republic, Hungary and Poland. In particular, we first look at inflation within the context of multivariate cointegration, where domestic and foreign price determinants are initially assessed in separate blocks (each single-theory based) in order to obtain a number of long-term attractors. Then, we put previous information in short term simultaneous vector equilibrium correcting models for each country. The importance of theory-based imbalances (attractors) in explaining inflation can be assessed at this stage. Our most significant empirical findings seem to substantiate the idea that many, if not all, theoretical determinants of inflation are of importance in those countries in question: the exchange rate and the output gap would appear to be of particular importance in explaining the phenomenon..

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1. Introduction

As a result of the decisions taken at the Copenhagen European Council meeting held in June 1993, the enlargement of the European Union (EU) to a number of applicant Central and Eastern European countries is no longer a question of *if* but *when*. From an economic point of view, EU accession will take place as soon as the applicant is able to: *(a)* cope with competitive pressure and market forces within the EU; *(b)* take on the obligations of membership, including allegiance to the aims of political unification as well as Economic and Monetary Union (EMU). The European Central Bank (ECB) points out that: “Even though the fulfilment of the Maastricht convergence criteria (including price stability, the sustainability of public finance, exchange rate stability in the framework of participation in the exchange rate mechanism and the convergence of interest rates) is not mandatory for EU accession, accession countries[3] should have macroeconomic programmes consistent with those prevailing in the Euro area in their policy agenda” (ECB, 2000, p. 44). Inflation is extremely important according to the EU policy agenda, as can be seen from the main conclusions reached by the Helsinki seminar held in November 1999: “Accession countries need to continue to implement monetary policies geared towards achieving and maintaining price stability, and to support this process with prudent fiscal policies and adequate structural reforms.” (ECB, 2000, p. 49). Reform in these countries means systematic changes in basic economic structures, with price liberalisation as a main step towards a market economy.

In this paper we formulate structural models of inflation for the Czech Republic, Hungary and Poland that allow for the stylised facts and peculiarities of each country. Hungary and Poland represent significantly different ways of switching from one economic regime to another[4], and are still characterised by persistent moderate inflation. Due to the lack of data, inflation within the Czech Republic has rarely been analysed using econometric models, while it is the transition economy with the best disinflation record. In addition, these three countries are of particular economic importance among the group of 12 accession countries. In fact, in 1998 they accounted for 66.7% of the accession countries’ total GDP (Poland alone accounting for 40%), and 55.8% of their total population (Poland 36.5%).

[3] At present, there are 12 “accession countries”, i.e. countries that have already started formal, separate negotiations for EU accession: Bulgaria, Cyprus, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Romania, the Slovak Republic and Slovenia. Turkey is only an official candidate for accession, as the country has yet to meet those pre-conditions needed before negotiations may start.

Their respective shares of GDP and population in the Euro area (4% of GDP and 20.1% of population) underlines the importance of this issue from the EU point of view as well.

During the 1990s, the accession countries achieved substantial reductions in their rates of inflation. The ECB (2000, p. 41) argues that this fact mirrors the growing importance of price stability as a statutory objective of accession countries' central banks. An important phase of negotiations involved a process of screening designed to provide a comparative analysis of the countries taking part in the project. National Programs for preparation for European Union Membership were produced. Priority has been given to programs designed, among other things, to produce adjustments to monetary policy instruments. In order to achieve greater price stability, central banks have to adjust and track the evolution of variables (intermediate targets) according to the requirements of the EU and the transmission mechanism that links the policy instrument setting to the primary objective. Monetary aggregates and exchange-rate policy are possible intermediate targets on which agents' inflation expectations are based (nominal anchors), but during the transition phase the link between money and inflation was adversely affected by financial innovation and the increasing mobility of international capital flows.[5] Hence, central banks chose either to target the inflation rate (e.g. in the Czech Republic from December 1997 onwards and in Poland from September 1998), or to move from fixed- to crawling-peg exchange rate regimes (e.g. in Hungary from March 1995)[6]. In the first case, the central bank directly targets inflation by pledging an explicit and credible commitment to maintaining the inflation rate at a given targeted level (or within a given target band); the central bank's inflation forecast is the nominal benchmark anchor. In the second case, a devaluation rate can be predicted (if economic agents believe in the pre-announced rate of crawl), and this allows for a higher rate

[4] According to W. Welfe (1995, p. 496), the case of Poland was one of an "over-night" change, while in Hungary it took several years.

[5] For a thorough discussion of these issues with reference to transition economies see NBH (2000a), NBP (1998), Coats (2000), and Kutan and Brada (2000). The debate about alternative monetary policies in EU accession countries is summarised in EBRD (2000, p. 52).

[6] The Hungarian crawling peg regime should be phased out in the near future: however, although "The Bank and the government both recognise the need to abandon the current exchange rate regime and agree on how it should be done, no agreement has been reached as yet about the timing of such a move" (NBH, 2000b, p. 4). Of course, inflation targeting is a viable option for Hungary too, as shown in Siklos and Abel (2001). Before this policy change, however, Siklos and Abel suggest that "the relative responsibilities and expectations of the central bank vis-à-vis the government need clarification and elaboration", and recognise that there are "mitigating circumstances that suggest caution in adopting this route at the present time" (p. 21).

of inflation than that of the country's main trading partners during the disinflation phase; the exchange rate is the nominal anchor in this case.[7]

In both targeting regimes, successfully targeting depends largely on relatively accurate inflation forecasts, as there can be implicit inflation targets in the rate of exchange rate crawl too. For example, the NBH (2000b, p. 4) says: "Monetary policy can reduce inflation [...] by setting the pre-announced rate of devaluation lower than *expected inflation*, i.e. via real exchange rate appreciation." (emphasis added). After all, monetary growth remains a dominant factor in explaining inflation; relative price adjustments and nominal wage shocks are also partly to blame, even if their impact on inflation can be modified by monetary and exchange rate policy.

The aim of the present paper is to model consumer price (CPI) and producer price (PPI) inflation rates in the Czech Republic, Hungary and Poland, and to provide quantitative evidence for the results given by alternative economic explanations in the different countries. The focus on CPI inflation results from the emphasis placed by transition economies' governments and central banks on inflation measures based on percentage variations in CPI levels (untransformed, net or smoothed). In Poland, the inflation target focuses on CPI, since there is a deep-rooted public belief that this is the fundamental measure of inflation. Since 2001, the Czech Republic targets year-on-year CPI growth, after three years of net inflation targeting (the inflation rate net of changes in administratively regulated prices, indirect taxes and fees). The analyses of the (non inflation targeting) Hungarian Central Bank are mainly based on core CPI inflation, which is generally considered a good index of inflation (see Valkovszky and Vincze, 2000). CPI inflation is also going to be of importance for future policymaking in the enlarged EU, as the harmonised index of consumer prices (HICP) is used by the ECB to assess risks to price stability in the Euro area, and to check for inflation convergence among EMU countries (see ECB, 1999). However, in order to get a better understanding of the dynamics of inflation we have extended our analysis to cover PPI inflation modelling as well.

The paper is sub-divided as follows. Section 2 surveys empirical studies of inflation in transition countries, and formulates a modelling strategy designed to cope with a number of

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7 Despite the present crawling peg regime, many economic observers agree that, prior to membership of the European Exchange Rate Mechanism ERM2, a period of freely floating exchange rate is advisable, in order to bring the market rate closer to the equilibrium rate. Corker, Beaumont, van Elkan and Iakova (2000) survey the main issues in EMU accession of five "advanced" transition economies.

methodological and data drawbacks. Given that we acknowledge the importance of several potential determinants of inflation, so that no "single-cause" explanation is going to suffice, we have devoted section 3 to the measurement of a number of long-term relationships which may help to explain the path taken by inflation in the transition economies during the 1991-2000 period. Section 4 analyses the feedback-effect that temporary imbalances ("gaps") between actual and long-term target levels have on domestic inflation rates. Parsimonious models of inflation for the three accession countries under scrutiny are constructed, and the relevance of different individual causes is compared within and between countries. Section 5 presents our principal findings and conclusions.

2. Stylised facts and the methodological framework of analysis

Inflation modelling is an inherently difficult task that is further complicated in the case of transition economies due mainly to the short time-span during which free-market prices have existed. Furthermore, as a result of the major structural changes that have taken place over the last ten years, the econometric relationships between inflation, money, wages, exchange rates and other indicators have been susceptible to instability. Finally, as far as the measurement of economic variables is concerned, empirical studies sometimes raise doubts about the reliability of certain official statistics.[8]

Given such difficulties, central banks forecast inflation using a variety of approaches, and publish and discuss inflation forecasts in detail without any reference to the technical instruments employed. On the other hand, it is widely acknowledged that important pre-requisites for the management of monetary policy, such as relatively accurate inflation forecasts and the assessment of the impact had by changes in monetary instruments on inflation, are also very important for successful targeting, and as such they can benefit from the use of suitable econometric tools.

The econometric modelling of the inflation rate can be accomplished by following a multivariate approach, whereby a number of alternative indicators are used to define a set of core variables that may influence inflation. Recent IMF studies represent a first step towards inflation model building based on indicators for the Czech Republic, Hungary and Poland (see Laursen, 1998; Doyle and Nyberg, 1999; Christoffersen and Wescott, 1999; Christoffersen, Slok and Wescott, 2001). Their main findings can be summarised as follows:

[8] Filer and Hanousek (2000) argue that the upward bias in the inflation measurement for the 1990-99 period can partly explain the so-called transition recession. However, the bias measurement issue in transition economies is controversial, as is clear from the discussion in Brada, King and Kutan (2000).

(a) there is statistical evidence in support of good inflation predictability using key macroeconomic determinants within an atheoretical VAR model; *(b)* the set of inflation determinants is country specific;[9] *(c)* dynamic relationships are mainly measured by impulse-response profiles that are in turn affected by the choice and order of the variables appearing in the model; and *(d)* the statistical robustness of the links between inflation and indicators is not firmly established due to a lack of statistical information.[10] The same countries have been analysed within a VAR framework by Brada and Kutan (1999), who find that import price changes play the most important role in explaining inflation dynamics, while nominal wage growth and money supply are quantitatively unimportant. The heterogeneity of outcomes and policy implications, along with the difficulty of comparing results, are typical problems of studies based on the VAR approach. Overfitting, the loss of parsimony and the lack of an economic interpretation of the results constitute serious drawbacks.[11]

The methodological approach we follow in the paper should be able to deliver results that are more interpretable in the light of the underlying economic theories. Then, in order to reduce the degree of arbitrariness in the choice of variables and to achieve more parsimonious specifications, our model-specification strategy consists in a two-stage approach based on theoretical structural economic considerations (a very similar approach is followed in Juselius, 1992). Due to the considerable variety of potential inflation determinants, the lack of data, and the possible regime shift over the sample period, a multivariate cointegration analysis including both inflation and the whole set of variables in question is not feasible. Our two-stage approach can be seen as a multivariate VAR extension of the two single-equation stages advocated by Engle and Granger (1987).

At the first stage (described in section 3), the potential determinants of inflation are grouped together in blocks, each item representative of the most popular inflation theories: cost push determinants (as labour cost); foreign prices and exchange rates; money overhang. We start with a VAR representation of the variables of interest in each block, and then develop sub-models simplifying the VAR, where restrictions are interpretable in the light of

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[9] Wages, import prices in national currency, and money for the Czech Republic; import prices, nominal and real exchange rates, and budget balance for Hungary; import prices in national currency, effective exchange rates, and retail sales for Poland.

[10] As shown by the US inflation forecasting exercise in Stock and Watson (1999), overfitting and parameter instability risks can be limited by long time spans, which unfortunately are not available in the present context.

[11] Alternatively, the classical structural modelling approach is applied to the Russian Federation by Basdevant (2000). Merlevede, Plasmans and van Aarle (2001) report preliminary results about macroeconometric models for six accession countries.

economic theory. In this manner, we are able to establish a connection (where it exists) between structural parameters and theoretical determinants of inflation. The representation of the whole model in terms of a number of blocks, is in keeping with the idea of testing for relationships that are unaffected by substantial economic changes, as suggested by Hall, Mizon and Welfe (2000). For this reason, particular care is given to assessing the stability of the identified long-term relationships.

At the second stage (in section 4), we use a VEqCM specification capable of representing short-term dynamics by means of stationary variables and past imbalances (the "gaps" detected by previous cointegrated relationships), and extend the single equation approach found in Ericsson, Hendry and Prestwich (1998), Gerlach and Svensson (2000), and Hendry (2000), in a multivariate direction At this point, demand-pull and cost-push inflation determinants are put together and their relevance assessed. We then introduce a number of parameter restrictions in order to improve efficiency. Since different restrictions are imposed in different equations, the model parameters are SURE estimated.

3. The block structure of inflation models

In this section we evaluate a number of long-term relationships which may help to explain the inflation path in the transition economies during the 1991-2000 period. In particular, section 3.1 measures the labour cost determinants of inflation, section 3.2 analyses the exchange rate issue, and section 3.3 is devoted to monetary determinants.

3.1 Labour markets and the wage gap

Empirical studies have extensively analysed the relationships between labour market, wages and prices in the Czech Republic, Hungary and Poland, and in Qin and Vanags (1996) the three countries are modelled using the same theoretical structure. Despite the fact that Poland has been more extensively analysed than Hungary,[12] a number of studies cover both Poland and Hungary.[13] The Czech Republic has not been so extensively analysed because of a lack of data due to the break-up of the CSFR at the end of 1992. Previous studies often used data covering different regimes (planned-transition-market), and the modelling strategies proposed in these studies can be broadly classified into two alternative methodological

[12] Welfe (1991, 1996, 2000), Blangiewicz and Bolt (1993), Golinelli and Orsi (1994), W. Welfe (1995), Osiewalski and Welfe (1997, 1998), Marcellino and Mizon (2000) refer to Poland, while Beaumont (1999) refers to Hungary.
[13] See Commander and Coricelli (1992) and Golinelli and Orsi (1998, 2000).

approaches: *(a)* switching parameter models;[14] *(b)* linear dynamic models with stable long-term relationships.

Empirical inflation models are generally linked to the cost-push theory of inflation and to the conventional wage equation, since such price- and wage-setting behaviour is presumed to be valid for both market and transition economies.[15] Hence, the list of potential variables that may be used to explain consumer (or producer) prices is quite similar from one study to another, and includes production cost components, such as labour cost, import prices and exchange rates. Since unit labour cost is of importance in explaining prices, in this section we check for the existence of stable, long-term relationships between its main determinants, namely real wage, labour productivity and labour demand-supply imbalances. Deviations of actual values from target (long-term) real wages will be labelled as "wage gaps". The concept of wage gap was first introduced by Sachs (1983) in order to define the departure of historical labour shares in output from an appropriate (target) level, measured on the basis of a given economic theory of interest. In what follows we propose a wage gap measure designed to evaluate the usefulness of these "supply side" imbalances in explaining the path taken by inflation in the Czech Republic, Hungary and Poland.

Real wage data are obtained by deflating per capita nominal wages with either CPI (consumer wage, *wpc*) or production prices (producer wage, *wpp*); labour productivity is measured using industrial output-employment ratios (*prod*). The unemployment rate (*ur*) and the vacancy-unemployment ratio (*vur*) are alternative measures of labour market imbalances, though the former has hardly ever turned out to be of significance during transition. All previous variables were measured on a monthly basis over the 1990-2000 period, are expressed in logarithms, and are seasonally adjusted (details are listed in the data appendix).

We established an unrestricted VAR model for the variables of interest, under the hypothesis of no switch between regimes. Potential parameter non-constancy due to important shifts was dealt with by modelling data over the post-1990 period, when any major changes are presumed to have already taken place.[16] The preliminary univariate analysis,

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[14] Either linear or non-linear models with either deterministic or endogenous switching; see Osiewalski and Welfe (1997, 1998), Golinelli and Orsi (2000), and Welfe (2000).

[15] Among others, Rutkowski (1994) finds that the principles of wage bargaining and efficiency wages were operating during pre-transition phases (see also Welfe, 1991). It is widely acknowledged that mark-up price setting and the link of domestic to foreign prices were also operational in transition economies; see for example Commander and Coricelli (1992), Qin and Vanags (1996), Osiewalski and Welfe (1997), Beaumont (1999), and Welfe (2000).

[16] Welfe (1996), Beaumont (1999), and Marcellino and Mizon (2000) are examples of this kind of modelling approach.

conducted on the basis of graphical inspection, correlograms and unit root tests, suggests that the variables listed above are either I(0) or I(1), and are not affected by important deterministic breaks over the period under scrutiny. We are thus able to perform the Johansen (1995) cointegration analysis, the main results of which are given in Table 1 below.

The lag length of the UVAR models is chosen on the basis of both AIC criterion and residual diagnostics. As many of the variables in question display trending paths, we always choose to include the intercept in the models (in the Hungarian VAR model, a linear trend restricted in the cointegration space is also included, as discussed below). We never specify impulse or step dummies in order to use the conventional critical values of cointegration tests (see the discussion in Doornik, Hendry and Nielsen, 1998).

Different sets of variables correspond to different countries. In the Polish VAR we use the most suitable variables in order to estimate the long-term relationship between labour's share of output (equal, in logs, to the real per capita producer wage minus labour productivity) and the labour market imbalances, as required by our modelling framework described above. As far as the Czech Republic's VAR is concerned, we use consumer, instead of producer, real wages since *Dwpp* was seen not to be stationary. The Hungarian model proved the most controversial since the available data did not provide us with much information given the issue we are analysing here. The problems evidenced by preliminary univariate and multivariate analyses (not given here) mirror those experienced in Golinelli and Orsi (2000): a lack of statistical information, probably due to the relevant policy changes, during the mid-1990s. Some tentative cointegration (at a 10% significance level) is achieved by ex ante restricting real wage and labour productivity in labour's share of output, and by performing a Johansen analysis at the bivariate level using (*wpp-prod*) and *vur* variables (plus the trend in the long-term relationship to proxy regime changes).

Tab. 1 – Cointegration analysis of wage models

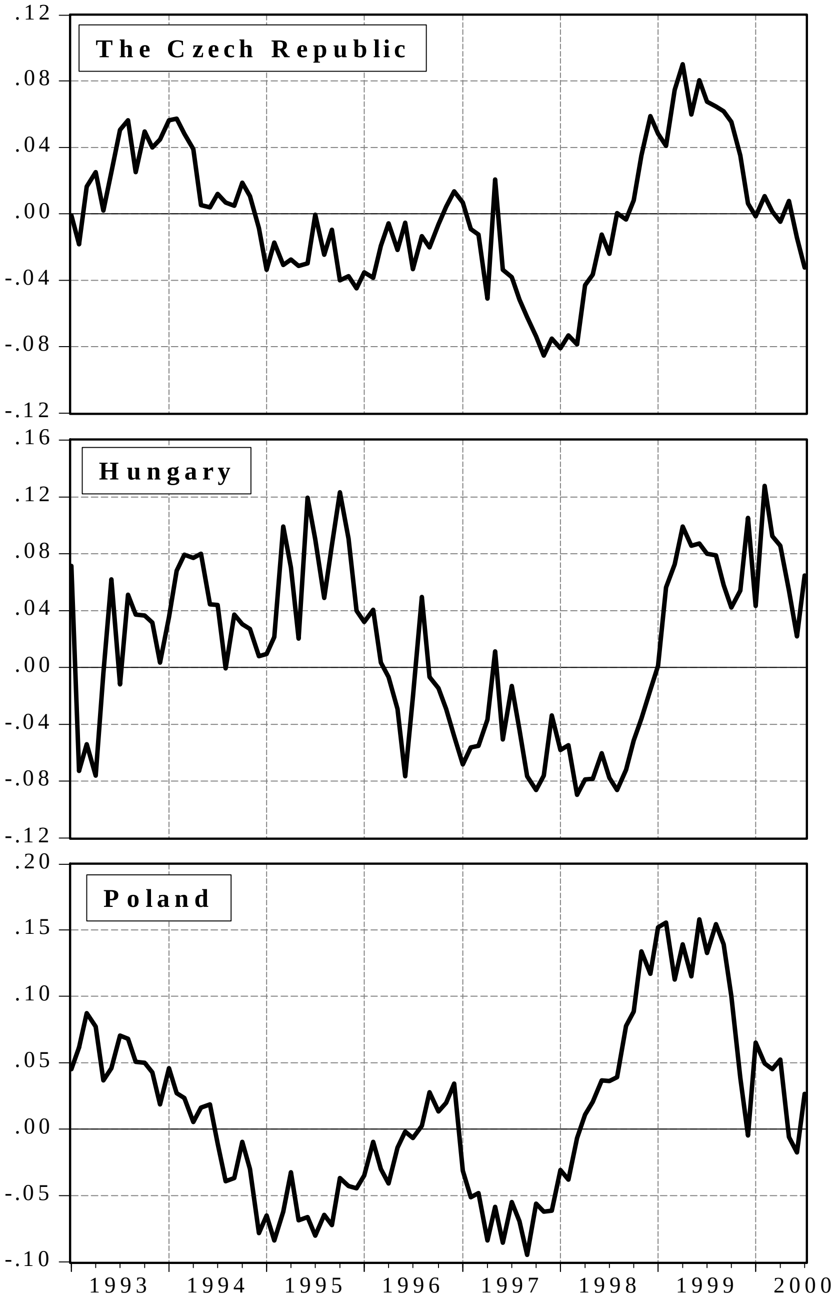
	Czech Republic	Hungary	Poland
Period	1993.1-2000.7	1991.1-2000.7	1991.1-2000.7
UVAR lags	11	3	3
Deterministic components	*c*	*c*, *t*	*c*
Cointegration rank (significance)	2 (5%)	1 (10%)	1 (5%)
Long run relationships:			
wpp	-	-	1
wpc	1	-	-
prod [s.e.]	-1 [-]	-	-1 [-]
(wpp-prod)	-	1	-
vur [s.e.]	-0.075 [0.009]	-0.407 [0.142]	-0.278 [0.044]
t [s.e.]	-	0.011 [0.002]	-
Loading factors:			
Dwpp [s.e.]	-	-	0 [-]
Dwpc [s.e.]	0 [-]	-	-
Dprod [s.e.]	0.208 [0.065]	-	0.119 [0.024]
D(wpp-prod) [s.e.]	-	-0.195 [0.056]	-
Dvur [s.e.]	0.467 [0.177]	0 [-]	0 [-]
Identification restrictions:			
statistic χ^2 (d.f.)	0.881 (2)	0.637 (1)	3.877 (3)
p-value	0.644	0.425	0.275

Notes: *c* (unrestricted intercept); *t* (restricted trend); *wpp* (real producer wage); *wpc* (real consumer wage); *prod* (labour productivity); *vur* (vacancy-unemployment ratio). "-" means "excluded from the analysis"; "[-]" means "restricted estimate".

Cointegration relationship results are quite clear cut: in the long term, the real wage develops a positive relationship with both labour productivity and the excess demand on the labour market (in keeping with theory). All cointegrating vectors are normalised in the wage variable: the equilibrium correction estimates refer to the labour share's actual and target deviations, that can be interpreted as an updated, statistically grounded version of the Sachs (1983) wage gap concept. Although we must be careful of the above-mentioned caveats, of the three, Hungarian wages seem the most elastic in relation to the vacancy-unemployment indicator.

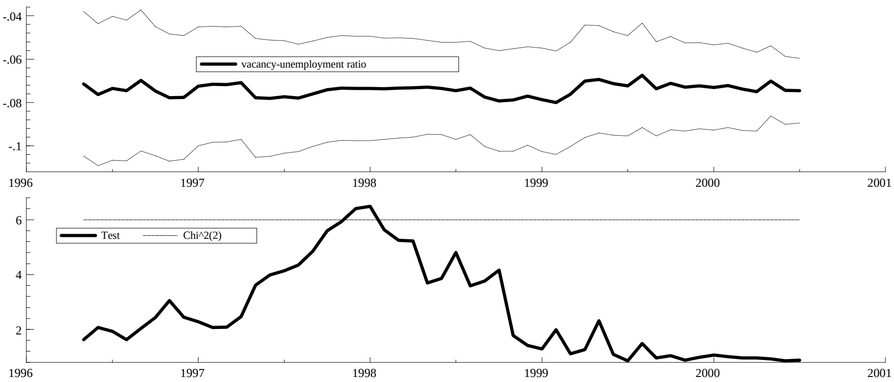
As far as loading factor estimates and weak exogeneity are concerned, out-of-target real wages affect short-term labour productivity in all countries (the efficiency wage effect). In Hungary and Poland, labour share imbalances do not feedback to the vacancy-unemployment rate, while in the Czech Republic, positive wage gaps lead to a short-term increase in labour demand (due to expansion of aggregate demand) that is higher than the increase in labour supply (due to higher real wages). The wage gaps, which correspond to the cointegration relationships reported in Table 1, are illustrated in Figure 1.

Fig. 1 - Estimated wage-gaps

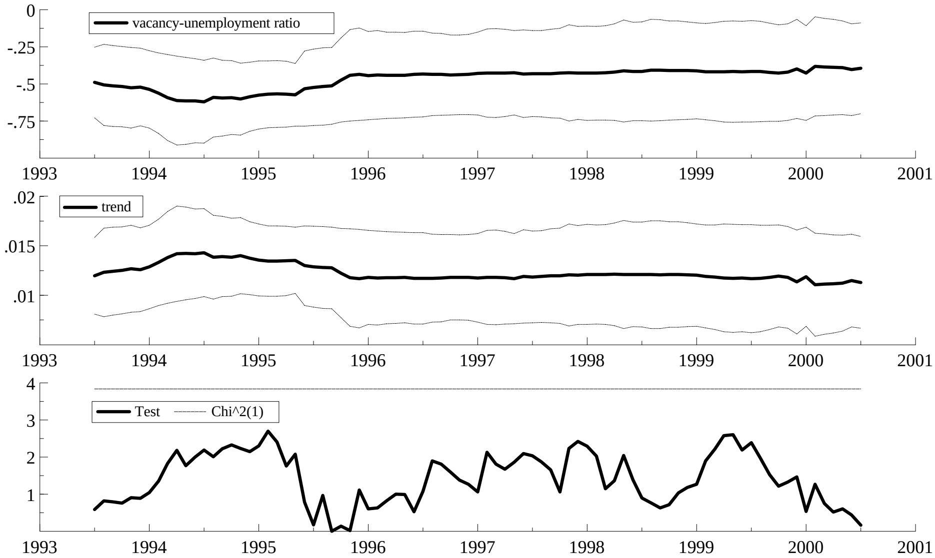


In order to make comparisons, the gaps are plotted over the same January 1993 – July 2000 period, despite the fact that they are available from January 1991 in the case of both Hungary and Poland. Overall, the three wage gaps show quite a similar path; in particular, note the marked tendency to shift from negative gaps (at the end of 1997) to positive ones (at the beginning of 1999).

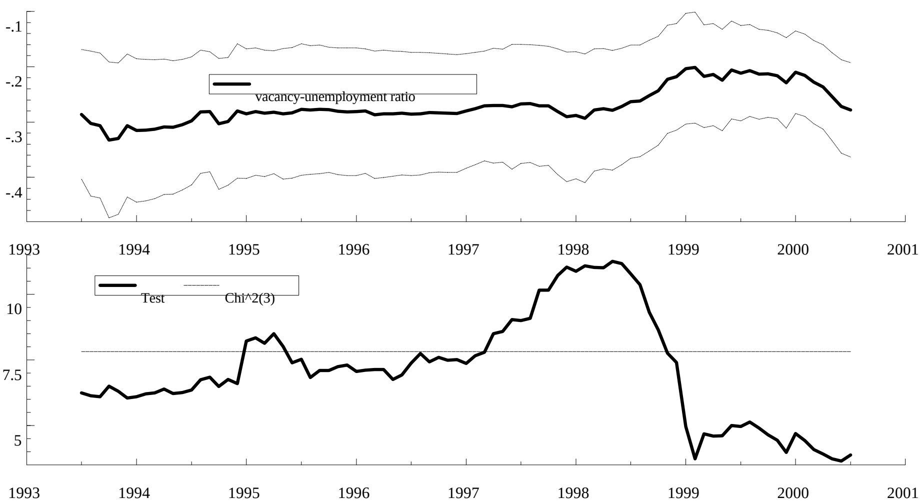
Fig. 2 - Recursive long-term analysis of the wage models



(a) Czech Republic



(b) Hungary



(c) Poland

The Hungarian wage gap is characterised by a more marked month-to-month variability than that of the other two countries, and in the period from 1996 to 1998 is persistently negative, probably due to the 1995 stabilisation measures (including a stringent incomes policy). All series are I(0) in the light of further unit root test and correlogram inspection.

An important issue is that of the stability of the wage-gap-generating relationships. The recursive long-term parameter estimates of the restricted models are summarised in Figure 2. The initial sample consists of the first 30 observations (40 in the Czech Republic case) designed to assess structural stability over a relatively lengthy time-span. Our findings point to stable long-term parameter estimates, despite the many economic policy changes introduced during the period of analysis. Plots of the overidentifying restriction tests in the lower panels of Figure 2 show that we hardly ever reject these restrictions at 5%, and never at 1% during the recursion period.

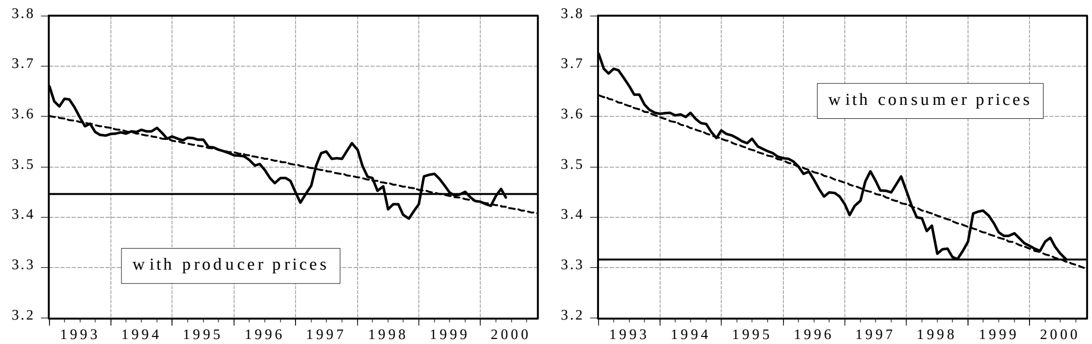
3.2 Foreign markets, the PPP, and the Balassa-Samuelson model

Since the Czech Republic, Hungary and Poland are all small, open economies, exchange rate behaviour is important to an understanding of their macroeconomic performance. As far as nominal figures are concerned, purchasing power parity (PPP) theory can be used to analyse the contribution made by foreign markets and exchange rates to domestic inflation. PPP is viewed as a long-term equilibrium condition of the real exchange rate, and may be empirically examined by defining deviations in historical real exchange rates from the constant equilibrium PPP: if such deviations are stationary over time, PPP holds in the long run, and the corresponding imbalances can be defined as “PPP gaps” (note that in the next multivariate long term analysis we will label as “PPP gap” the residuals from the Balassa-Samuelson long term relationships). PPP price determination predicts that positive PPP gaps exert positive effects on domestic inflation (*e.g.* through imported inflation). Any evaluation of the performance of alternative exchange-rate regimes is beyond the aims of the present work, as the issue is influenced by too complex a set of actions.[17] Instead, by testing for the existence of long-term PPP, we are simply trying to shed some light on the behaviour of real exchange rates in the Czech Republic, Hungary and Poland.

[17] Hochreiter (2000) provides a thorough evaluation in historical terms of the alternative exchange rate policies potentially available to the accession East European countries; the Hungarian case is also presented by Szapary and Jakab (1998). The choice of the exchange rate regime in future phases of EU accession by the East European countries is discussed by Szapary (2000).

The main drawback of the PPP empirical approach (through real exchange rate stationarity tests) is the time-span of the data used in the analysis: that can be too short to give enough power to the testing procedure. However, the extension of the sample over time can be inappropriate because of the risk of regime changes. Our analysis focuses on monthly exchange rate averages from 1991 onwards (1993 in the case of the Czech Republic), as they can be presumed free from any significant breaks.[18] The measurement problems associated with PPP theory are also widely acknowledged: PPP tests can use either bilateral or effective exchange rates, and relative prices can be measured using either production or consumer prices. In empirical studies, traded goods prices are usually represented by manufacturing price indexes, while CPI is generally considered a mixture of traded and non-traded goods prices. However, producer price indexes of different countries are measured using different weights, and often contain some form of non-tradable component. These factors, together with others, can mislead PPP tests on real exchange rates. We take Czech crown, Hungarian forint and Polish zloty exchange rates against the Euro, (*euro*), as these countries have close trade relationships with the EU,[19] and are following a phased process towards adoption of the Euro. Real exchange rates (in logs) have been computed using both producer (*ppp_pp*) and consumer (*ppp_pc*) prices, and plotted in Figure 3 (data sources and transformation are given in the appendix).

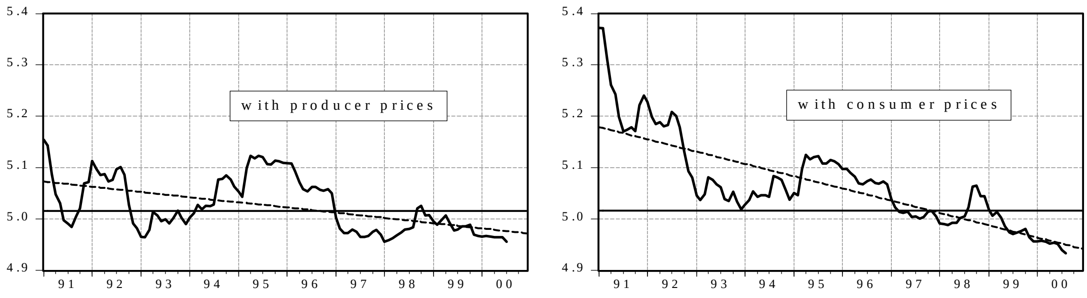
Fig. 3 – Real effective exchange rates against the Euro



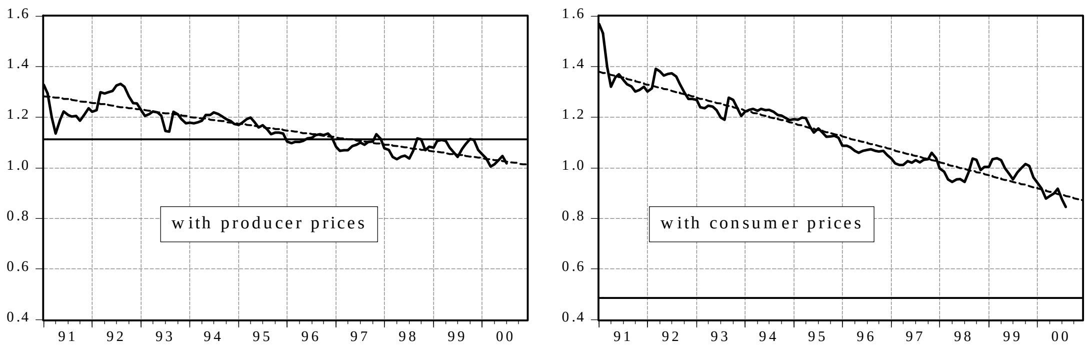
(a) Czech Republic

[18] Kocenda (1999) tackles the issue of testing for the existence of structural breaks in exchange rate series during the 1991-1997 transition. Results show that many East European countries (including the Czech Republic, Hungary and Poland) either did not experience a structural break in their exchange rates, or its effect was quite limited.

[19] After COMECON ceased to exist in 1991, most accession countries reoriented their external trade to Western Europe; as a result, Czech Republic, Hungary and Poland exported around 60% to 70% of their total exports to the Euro area.



(b) Hungary



(c) Poland

In order to improve readability, each graph features two lines: the (dotted) downward trend and the constant line which represent, respectively, estimates of the deterministic components of univariate Dickey-Fuller (DF) real exchange rate models with and without trend.[20]

[20] In some cases, as in the Czech Republic and Poland, *ppp_cpi* models without a trend are very imprecisely estimated; this is the reason why the horizontal lines are so far away from the historical paths.

Tab. 2 – Dickey-Fuller tests of real exchange rates against the Euro

	Model	Statistic	k	T
The Czech Republic:				
- *ppp_ppi*	c, t	-3.451 **	10	90
- *ppp_ppi*	c	-2.873 **	16	90
- *ppp_cpi*	c, t	-4.552 ***	1	92
- *ppp_cpi*	c	-3.863 ***	18	92
- *prod_diff*	c, t	-2.049	17	92
- *Dprod_diff*	c	-3.370 **	16	92
Hungary:				
- *ppp_ppi*	c, t	-2.757	1	115
- *ppp_ppi*	c	-2.324	1	115
- *Dppp_ppi*	c	-8.382 ***	0	115
- *ppp_cpi*	c, t	-2.361	10	116
- *ppp_cpi*	c	-1.769	10	116
- *Dppp_cpi*	c	-3.157 **	9	116
- *prod_diff*	c, t	-4.893 ***	18	108
Poland:				
- *ppp_ppi*	c, t	-3.411 **	16	110
- *ppp_ppi*	c	-3.066 **	11	115
- *ppp_cpi*	c, t	-2.153	11	116
- *ppp_cpi*	c	-0.631	18	109
- *Dppp_cpi*	c	-3.232 **	7	116
- *prod_diff*	c, t	-3.255 *	16	111

Notes: c = only the constant; *c, t* = model with constant and trend; *, **, *** represent 10%, 5%, and 1% rejection; k = order of augmentation of DF's tests; T = observations.

DF test results are reported in Table 2, and require some explanation: in the Czech Republic the real exchange rates are stationary, while in Hungary they are I(1). In Poland the producer price exchange rate is always stationary (with or without trend), while the CPI-based one is I(1).

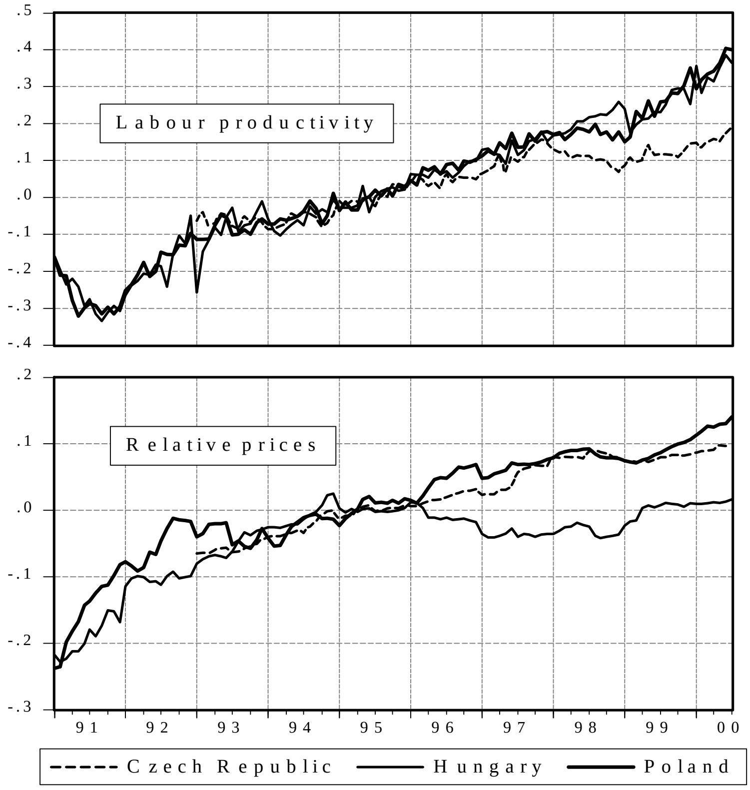
Given the heterogeneity of results, it is hard to get any useful information from them; in other words, in order to interpret real exchange rate, DF tests alone are not enough: unit root outcomes need to be supported by further evidence. The exchange rates given in Figure 3 seem to behave more homogeneously than the test results in Table 2. Real appreciation of exchange rates is evident, though it follows different paths: in terms of producer prices, real rates seem more stable over time than those obtained using consumer prices. This fact is partly due to significant tradable / non-tradable price adjustments over the analysed period.

The tendency of real exchange rates to appreciate (mainly those calculated in terms of CPI) deserves further explanation. The Balassa-Samuelson (BS)[21] model may be helpful here, as it explains real exchange rate movements in terms of sectoral labour productivity differences: if tradables (compared to non-tradables) labour productivity grows faster in Eastern European countries than in the Euro area, then their currencies should appreciate against the Euro. The BS prediction can be broken down into two sub-statements *(a)* the PPP between East European and EU countries holds for traded goods; *(b)* if labour productivity grows more in East European than in EU countries, the relative price of non-traded goods grows faster in East European than in EU countries. Despite previous reservations concerning the measurement of tradable goods prices using manufacturing indexes, statement *(a)* is confirmed by the stability over time of real exchange rates in terms of producer prices, as can be seen from Figure 3 and from the results in Table 2 (in the Czech Republic and Poland, *ppp_ppi* is 5% mean reverting, whereas in Hungary it is quite close to the 10% critical value). Statement *(b)* is analysed on the basis of Figure 4, where both relative manufacturing labour productivity and relative consumer-producer prices (in terms of the Euro area figures) are reported on a country-by-country basis.

The overall picture is in line with BS model sub-prediction *(b)*: in the three East European countries, labour productivity grows at a faster pace than in the Euro area (in Hungary and Poland the growth rate is similar, less pronounced in the Czech Republic), and the relative consumer-producer price increases are higher than in the Euro area (with the sole exception of increases in Hungary during the 1996-1998 period).

[21] See Balassa (1964), and Samuleson (1964).

Fig. 4 – Labour productivity and relative prices (with respect to the Euro area)



Given that this preliminary analysis seems to favour the BS model, we try to formally identify the BS effect in the countries in question by testing for the existence of a negative long-term relationship between real exchange rates and relative labour productivity. As in Alberola *et al* (2001), the index of relative sectoral productivity (*prod_diff*) is measured by the log-difference of productivity in the Czech Republic, Hungary and Poland compared to productivity in the Euro area (details concerning data sources can be found in the appendix). As far as the unit root tests in Table 2 are concerned, *prod_diff* in Hungary and in Poland is trend stationary, while it is difference stationary in the Czech Republic. As in the previous section, the long-term analysis is conducted in the field of Johansen's cointegration, since all variables are at most I(1).[22] The results are shown in Table 3, with reference to both producer and consumer price real exchange rates for each country.

In the Hungarian case we chose a UVAR specification with longer lags than in other countries due to autocorrelation. Overall, there is a significant long-term BS effect which, as was expected, is more pronounced in consumer price real exchange rate cases. In Hungary,

[22] Over the whole period of analysis, the series do not exhibit relevant deterministic breaks (results not reported).

the long-term response of the real exchange rate to productivity is smaller than in the Czech Republic or Poland (about half the size). High Czech elasticity is partly due to a lower productivity trend than in the other two countries (see Figure 4).[23]

Tab. 3 – Cointegration analysis of PPP Balassa-Samuelson models

	Czech Republic		Hungary		Poland	
Period	93.1-00.6	93.1-00.8	91.2-00.7	91.1-00.7	91.1-00.7	91.1-00.8
UVAR lags	3	3	13	6	3	4
Cointegration rank (significance)	1 (10%)	1 (5%)	1 (1%)	1 (10%)	1 (1%)	1 (1%)
Long-term relationships:						
ppp_cpi	-	1	-	1	-	1
[s.e.]	-	[-]	-	[-]	-	[-]
ppp_ppi	1	-	1	-	1	-
[s.e.]	[-]	-	[-]	-	[-]	-
prod_diff	0.628	1.238	0.135	0.337	0.402	0.796
[s.e.]	[0.138]	[0.143]	[0.052]	[0.121]	[0.040]	[0.056]
Loading factors:						
Dppp_cpi	-	-0.136	-	-0.080	-	-0.203
[s.e.]	-	[0.031]	-	[0.032]	-	[0.033]
Dppp_ppi	-0.132	-	-0.174	-	-0.285	-
[s.e.]	[0.038]	-	[0.042]	-	[0.044]	-
Dprod_diff	0	0	0	0	0	0
[s.e.]	[-]	[-]	[-]	[-]	[-]	[-]
Identification restrictions:						
statistic χ^2 (d.f.)	0.257 (1)	0.392 (1)	0.239 (1)	4.54 (1)	0.007 (1)	4.14 (1)
p-value	0.612	0.531	0.625	0.033	0.953	0.042

Notes: see Table 2.

These results confirm our previous graphical analysis: the real appreciation of the CPI is very important, since the growth in labour productivity (relative to the Euro are one) leads to an increase in the inflation differential of the CPI compared with the Euro area. In the Czech Republic and Poland, the speed of adjustment of the CPI and PPI real exchange rates is roughly the same, while in Hungary the *ppp_cpi* loading factor entails much slower adjustment (about 8% of the previous month's imbalance). The assumption of the weak exogeneity of relative productivity is never 1% (in many cases 5%) rejected.

A quantitative assessment of the CPI inflation rate differences for each country compared with the Euro area is obtained as a by-product of the long-term estimates given in

[23] By using a different (and much more detailed) methodological approach, Simon and Kovacs (1998) estimate a 0.66 appreciation elasticity coefficient as a response to the change of productivity in

Table 3. The resulting figures, calculated using labour productivity trend-rates during the 1993-2000 period, are reported in Table 4 below.

The point estimates of the BS effect shown in Table 4 are above the level of deviation allowed under the Maastricht CPI inflation criterion, and this result is in line with the empirical results given by Szapary (2000, pp. 5-6). Moreover, they confirm the finding of Corker *et al* (2000, p. 11), whereby "the contribution of the BS effect to CPI inflation might be expected to be larger than the 1-2 percentage points per annum estimated for EU countries experiencing productivity catch up".[24] However, the wide confidence intervals (especially in Hungary) suggest that we ought to be careful about drawing policy implications from such estimates.

Tab. 4 – The contribution to annual CPI inflation made by the BS effect (%)

	Czech Republic	Hungary	Poland
Productivity growth	3.5	6.4	6.4
BS effect (point estimate)	4.3	2.1	5.1
BS effect (95% upper bound)	5.3	3.7	5.8
BS effect (95% lower bound)	3.3	0.6	4.4

Note: 95% bounds are obtained using long-term s.e. reported in Table 3.

Stability analysis, in terms of recursive estimation and testing, is quite satisfactory (see the plots in Figure 5). As you can see, parameter estimates are fairly stable over time, and when restrictions are 5% rejected at the end of the period, the 1% line always lies well above the statistic plots.

Given that long-term results are quite robust to a number of alternative options (the inclusion of the trend in the cointegration space, the exclusion of rejected weak exogeneity restrictions, etc.), the distances between actual real exchange rates and the target long-term PPP in Table 3 are considered to be "PPP gaps", and are plotted in Figure 6. As opposed to the wage gaps in Figure 1, the PPP gap patterns are quite country-specific. On the other hand, the individual countries' gaps when measured using different price indexes prove to be very similar, with the exception of the Hungarian gap in the first half of the sample.

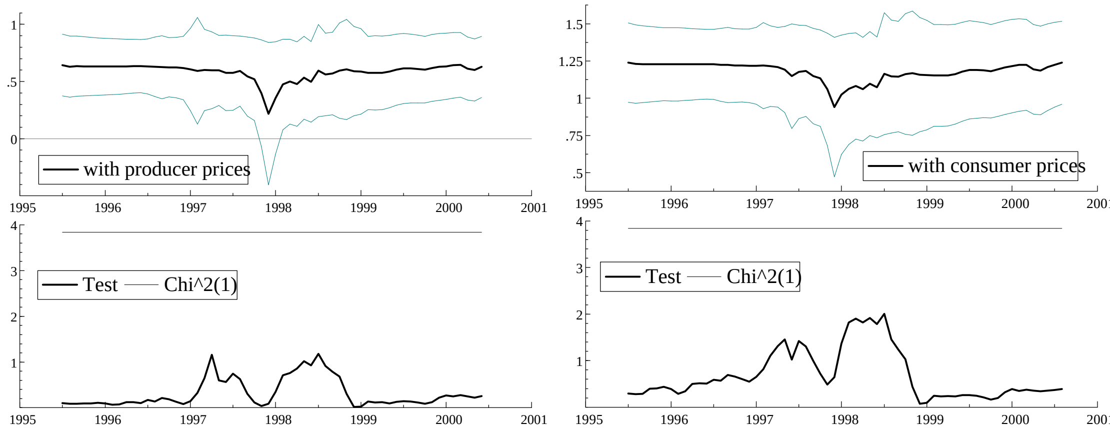
Hungary. However, such a measure is only slightly 5% significantly different from our long-term estimate.

[24] Note that one relevant conclusion of the Vienna Seminar (December, 2000) was that "[...] price adjustments are part of the process of transition and catching-up, and will typically entail inflation rates that will, for some time, range above those prevailing in the euro area [...].", ECB (2001, p. 108).

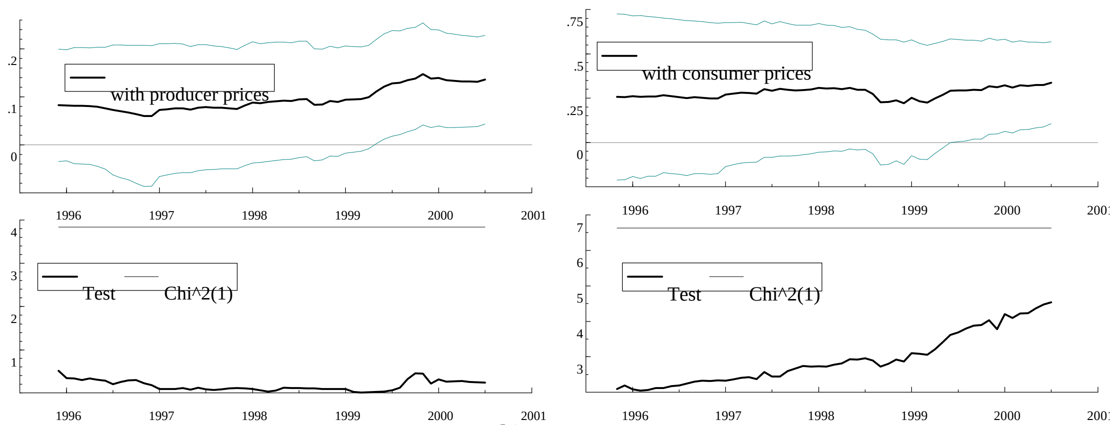
3.3 Financial markets and the money gap

According to quantity theory of money, the monetary determinants of inflation are based on the assumption that price levels are caused by money. In order to measure the excess money that potentially leads to inflationary pressures, we need to analyse the long-term relationship between money, output and its opportunity-cost. Alternatively, a simple indicator of inflationary pressure can be found in the so called P-star approach (see Hallman, Porter and Small, 1991).

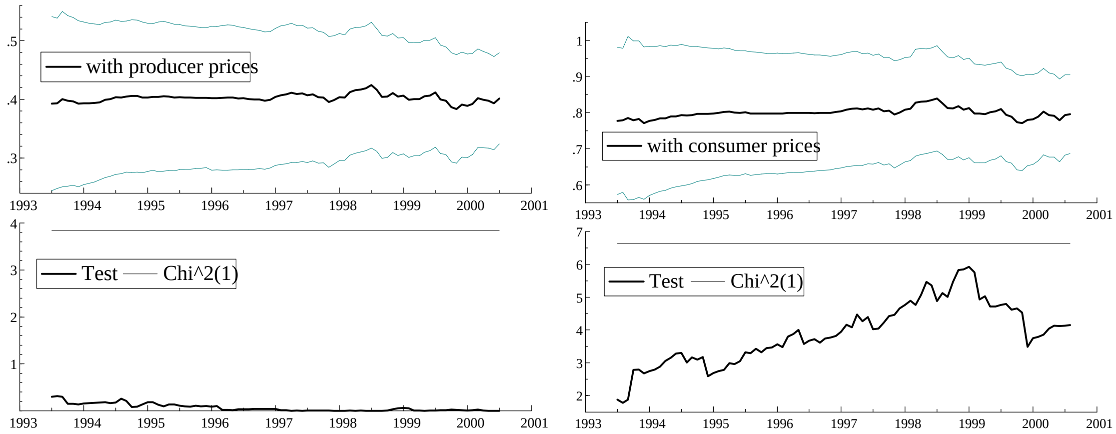
Fig. 5 - Recursive long-term analysis of the PPP Balassa-Samuelson models



(a) Czech Republic



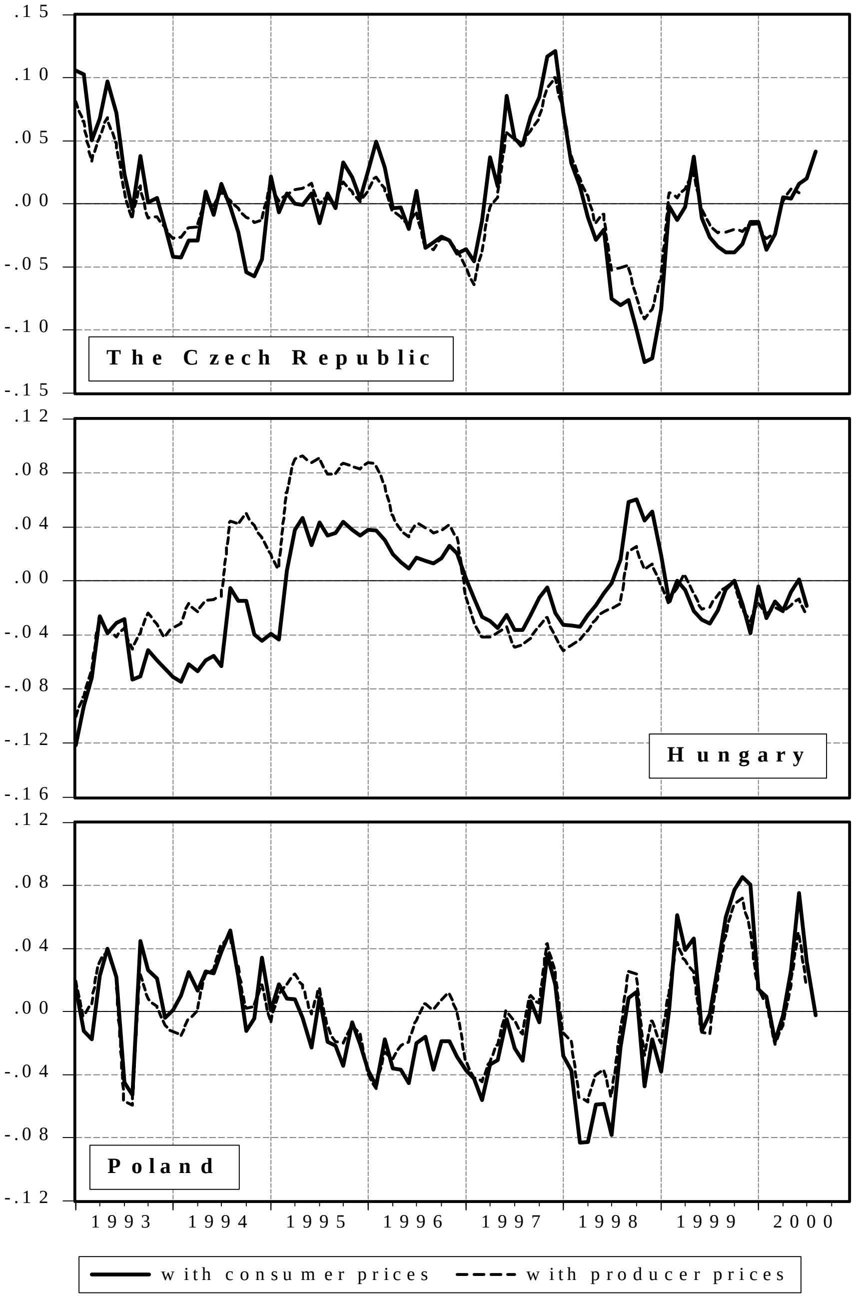
(b) Hungary



(c) Poland

The existence of such monetary indicators of inflation depends upon the empirical assessment of the causal link between money and prices, as shown by Hall and Milne (1994). In analysing this issue in the case of the three transition economies in question, Qin and Vanags (1996) recursively regress inflation to money growth and vice-versa, over a period covering the 1980s and the early 1990s: they found little evidence to support the existence of such a causal link. Nonetheless, it should be noted that Qin and Vanags' analysis covers a mixed period featuring both planned and transition economies, where "the underdevelopment of financial markets inevitably limits the scope for an independent money supply policy implemented through open market operations" (p. 154).

Fig. 6 - Estimated PPP-gaps



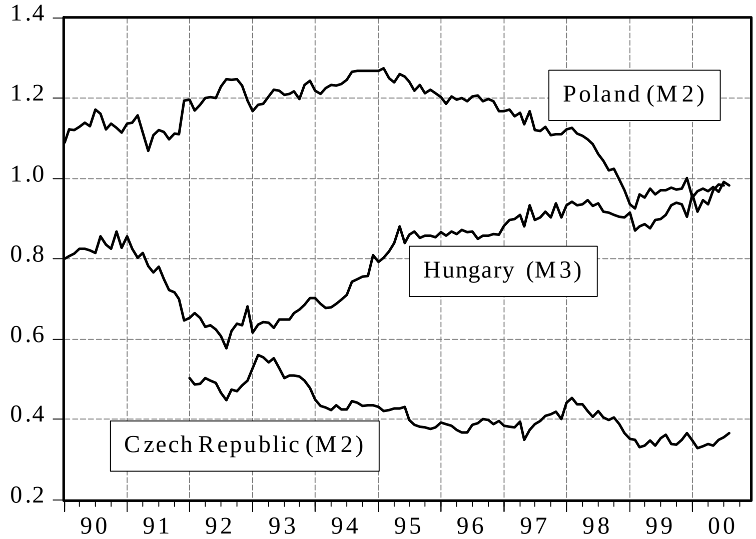
In this section we are going to address the first part of the previously-mentioned issue by looking for stable long-term money relationships as well as alternative indicators based on the P-star approach, and by subsequently measuring the corresponding "money gaps". Then in section 4, we will be looking at the question of the direction of causality between money gaps and inflation. Thus the present section is exclusively devoted to converting information about money velocity, output, prices, interest rates and other relevant variables into money gaps, *i.e.* into potential monetary explanations of inflation. This modelling strategy starts from the question of measurement.

As far as the choice of the more appropriate monetary aggregate is concerned, Brada and Kutan (2001) study the convergence of monetary patterns between East European and EU countries by focusing on narrow money ("reserve money" according to the IMF's International Financial Statistics), as they think it closely reflects the policy stance taken by the monetary authorities. On the other hand, many authors prefer to use broader measures of money in order to avoid the effects of portfolio reallocation due to financial innovation.[25] We tested for stable, long-term relationships by using alternative money aggregates; however, the results we give are based on broad, real money balances in logs *mp* (deflated using CPI), as we failed to identify any stable relationship involving narrower aggregates. The output measure *y* is the log of GDP at constant 1995 prices (annual GDP data were Chow-Lin interpolated by using monthly indexes of industrial production); the opportunity cost *r3* is measured by the nominal short-term interest rate. Data sources and transformations are listed in the appendix.

The simplest way to measure the real money gap is to calculate it using money velocity figures: if money velocity is stationary around a constant value, or follows a linear trend, then the resulting velocity gaps can be used to forecast future inflation. The broad money velocity figures in Figure 7 show trending paths, and they seem quite reasonable in the light of the effects of continuous financial market innovation during the period of analysis in question.

[25] Nijsse and Sterken (1996) analyse the household money demand function in Poland over the 1969-1995 period, and find a stable long-term cointegrated relationship for M2, but not M1, money demand. In a paper on inflation in the Baltics, Russia and other countries of the former Soviet Union, Ghosh (1997, p. 10) claims that "the economic concept of a money demand function might be expected to apply to broad money".

Fig. 7 – Broad money velocity



We can see that in Hungary, money M3 velocity slightly increased over this period, while money M2 decreased in the other two countries. It should be pointed out that the limited availability of data prevented us from using an entirely homogeneous definition of money.

Tab. 5 – Unit root test results of the broad money velocity

	T	Model	DF	k	PB	k
The Czech Republic:						
- *y-mp*	89	c, t	-2.17	14	-2.30	18
- *D(y-mp)*	89	c	-4.42 **	13	-	-
Hungary:						
- *y-mp*	115	c, t	-3.81 *	9	-	-
- *D(y-mp)*	115	c	-4.83 **	2	-	-
Poland:						
- *y-mp*	115	c, t	-1.50	12	-3.57	17
- *D(y-mp)*	115	c	-3.18 *	11	-	-

Notes: *T* = number of observations; *c* = only the constant; *c, t* = model with constant and trend; *DF* = Dickey-Fuller statistic; *PB* = Perron model 3 min t-α statistic; *, ** are 5%, and 1% rejections; *k* = order of augmentation of the test;

Tab. 6 – Johansen's cointegration analysis of broad money models

	Czech Republic	Hungary	Poland
Period	1993.1-2000.8	1991.1-2000.7	1991.5-2000.8
UVAR lags	2	4	4
Deterministic components	*c, t, d*	*c, t*	*c, t*
Cointegration rank (significance)	1 (5%)	1 (5%)	1 (1%)
Long-term relationships:			
mp	1	1	-
y [s.e.]	-1.692 [0.147]	-1 [-]	-
mp-y	-	-	1
r3 [s.e.]	0.524 [0.139]	0.442 [0.197]	0.263 [0.716]
d12p [s.e.]	-	-	1.867 [0.643]
t [s.e.]	-0.0006 [.0003]	0.0051 [0.0004]	0 [-]
Loading factors:			
Dmp [s.e.]	-0.107 [0.063]	0 [-]	-
Dy [s.e.]	0.168 [0.039]	0.128 [0.022]	-
D(mp-y) [s.e.]	-	-	-0.027 [0.011]
Dr3 [s.e.]	0 [-]	0 [-]	-0.035 [0.008]
Dd12p [s.e.]	-	-	-0.024 [0.007]
Identification restrictions:			
statistic χ^2 (d.f.)	0.058 (1)	1.467 (3)	0.889 (1)
p-value	0.809	0.690	0.346

Notes: *c* (unrestricted intercept); *t* (restricted trend); *d* (unrestricted impulse dummies); *mp* (real money); *y* (real GDP); *r3* (nominal interest rate); *d12p* (year-on-year inflation rate). "-" means "excluded from the analysis"; "[-]" means "restricted estimate".

In the presence of trending behaviour, we prefer to test for unit roots in univariate models embodying a linear trend (see Table 5): money demand allows for the possibility of a trend in velocity (see Orphanides and Porter, 2000). Furthermore, Carlson, Craig and Schwarz (2000) show that in the US case, equilibrium money velocity is sometimes subject to shifts. Given that structural breaks may reduce the power of DF tests, Table 5 also gives Perron (1997) tests (PB) that allow for a one-off break in both constant and trend slope.

Overall, results suggest that of the three countries, Hungary is the only one in which broad money velocity is trend stationary; thus detrended data can be used to forecast Hungarian inflation. However, since the velocity-gap models entail a number of parameter restrictions (unit money elasticity to output, and the absence of opportunity-cost effects), we are going to test for long-term relationships in more general models by adopting the Johansen cointegrated VAR approach. The main results of this are given in Table 6.

The search for cointegrated money demand is a difficult task; it can only be clearly identified in case of the Czech Republic: long-term money elasticity to income is significantly larger than one, and semi-elasticity to the nominal rate is significantly negative. In the short run, money changes adjust about 10% of previous periods disequilibria, and the nominal short-term interest rate is seen to be weakly exogenous for the analysis of long-term parameters.[26] As far as Hungary is concerned, the identified relationship comes very close to the trend-velocity univariate model: in the long run, elasticity to income point estimate is not significantly different to one, while semi-elasticity is significantly negative. The money gap feeds back to output, rather than to real money balances and interest rates; this fact could prevent money imbalances from being used to explain the inflation path.

The Polish case is the most problematic of the three: in order to find a long-term relationship we have to explain money velocity (by setting the unit elasticity of money to income) in terms of the short-term interest rate and the year-on-year consumer inflation rate. The long-term parameter estimates of the VAR model for Poland resemble those reported in Nijsse and Sterken (1996, p. 16); in particular, the inflation effect (of crucial importance to finding a cointegrated relationship) suggests that high inflation prevents real money holdings.

Despite the said difficulties, the recursive estimates in Figure 8 show fairly stable parameters, and the identification restrictions are not rejected. The corresponding money gaps, a measurement of the difference between real money balances and the corresponding target level, are plotted for the 1993-2000 sample in all three cases (see Figure 9 below).

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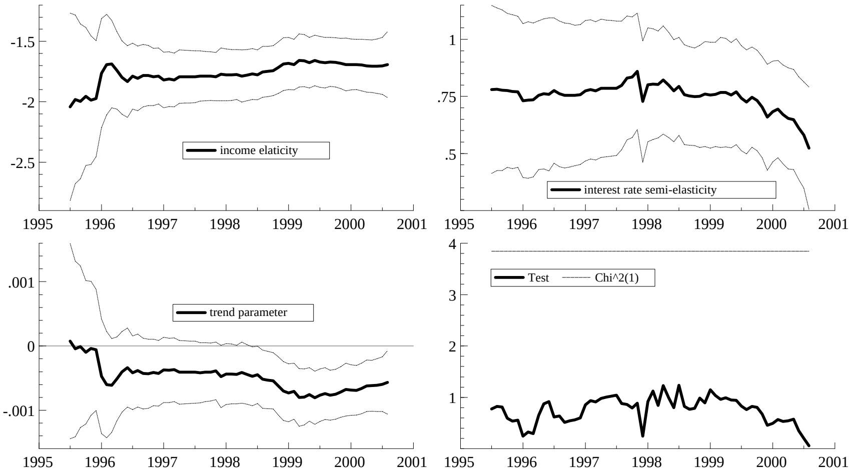
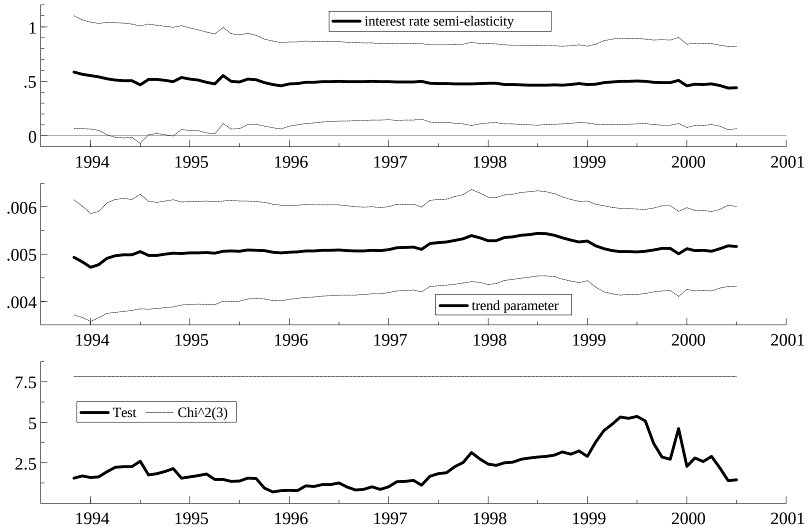
[26] The only drawback is the use of three impulse dummy variables in May, June, and July 1997, when the CNB defended the Czech crown from a speculative attack. Because of dummy variable additions, we must be particularly careful in assessing the cointegration rank.

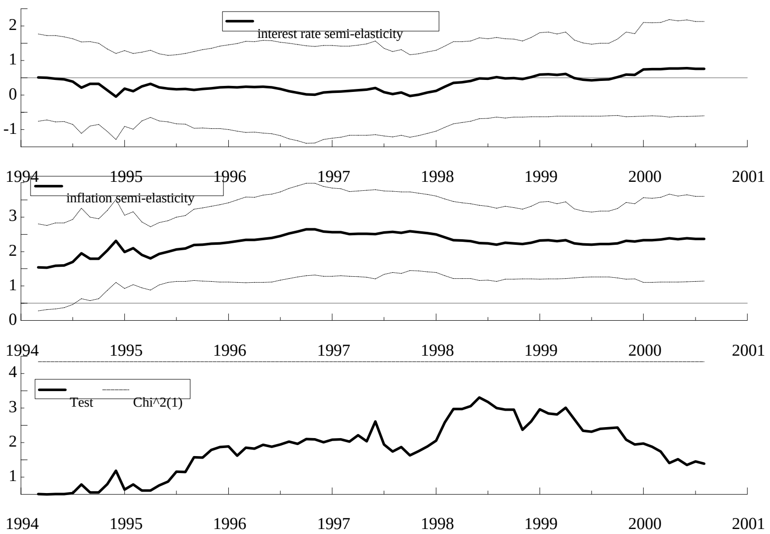
Fig. 8 – A recursive, long-term analysis of the broad money model



(a) Czech Republic

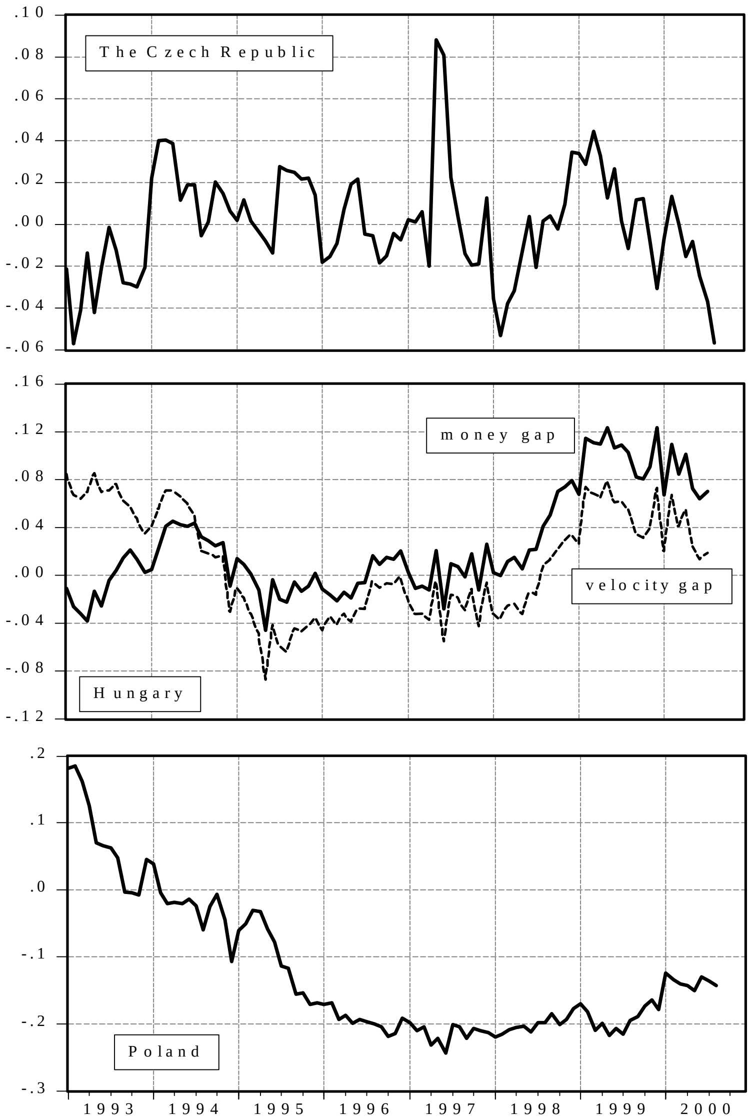


(b) Hungary



(c) Poland

Fig. 9 – Estimated money gaps

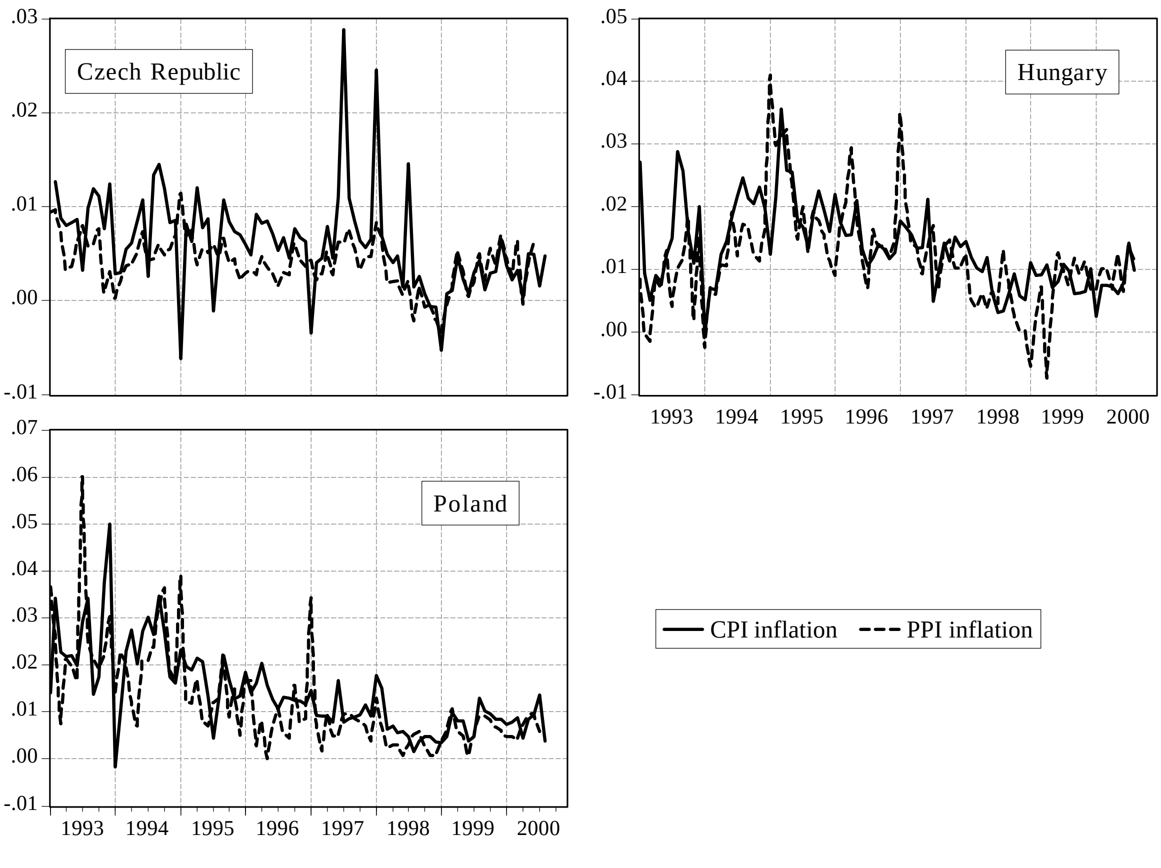


In the Hungarian case two plots are shown, because we identified two alternative, albeit rather similar, equilibrium relationships: the money demand imbalance from Table 6 (money gap), and the inverted-sign money velocity differences from the trend line resulting from unit root tests in Table 5 (velocity gap). Money gap paths behave quite differently from one country to another: in the Czech case, the money gap fluctuates quite regularly after an initial discrepancy, and peaks during the crisis of the Czech crown; in Hungary the two gap measurements behave quite similarly, displaying an excess of money in the last part of the sample; finally, a continuous tendency to shrink money can be seen in Poland.

4. Short-term inflation models

Before evaluating the main causes of inflation during the period of transition to a market economy, we need to acknowledge the role played by relative price adjustments. In fact, as Wozniak (1999) points out, in Poland relative domestic prices have undergone a dramatic realignment in the wake of price liberalisation, of the removal (or sharp reduction) of subsidies, and of exchange rate movements. Given that we believe these realignments to have mainly affected (although probably in different ways) the inflation rates of individual groups of commodities, including both consumer (CPI) and producer (PPI) price indexes, in this section we are going to model and compare outcomes for both inflation rates, measured by first differences of CPI and PPI log-levels (the historical paths are illustrated in Figure 10).

Fig. 10 – Monthly inflation rates



The plots in Figure 10 suggest that inflation variables are stationary (as confirmed by a look at correlograms and unit root test results), and are affected by some outliers. Nevertheless, CPI and PPI inflation patterns are quite different, and the PPI inflation rate is often lower than the CPI inflation rate, although there is a common tendency to converge towards lower rates by 1998. From 1999 onwards, inflation has recovered mainly due to oil and raw materials price increases and to appreciation of the dollar.

The starting point for short-term (second step) multivariate modelling is a (general) VEqCM where previous consumer price and producer price inflation rates are included, together with all other variables of interest in differences (listed in the lower section of Tables 1 and 6).[27] VEqCM specifications present one-time lagged values of: the wage gap, PPP gaps and the money gap (as defined in sections 3.1, 3.2 and 3.3). These gaps are what Hendry (2000) calls "feedback effects"; the corresponding parameters are the VEqCM loading factors that measure the speed of adjustment of the variables towards the identified long-term cointegrated relationships. In the adopted specification we have added three more variables in order to take account of the foreign determinants of domestic inflation: oil price inflation (*Dpoil*), inflation in the prices of other raw materials (*Dpraw*), and production growth in the Euro area (*Dyeuro*), respectively.[28] Finally, the demand-pull determinants of inflation in our VAR models also include lags in the rate of capacity utilisation (*qr*, see data appendix), since this is the typical demand pressure variable.[29]

At first it would appear that a huge number of parameters need to be estimated, and the resulting problem of efficiency is further exacerbated by the considerable number of lags (often 9-11) needed to provide non-autocorrelated, homoskedastic residuals. We paid special attention to the dynamic specification of our models, since the degree of inertia in the inflation rate is an important policy constraint, and reliable enough to lead us to believe that economic agents primarily base their inflation expectations on past results (given the high inflation experienced during the first phase of transition). In addition, in setting the lag length, long lags may appear significant not because the result of dynamic structure, but because such coefficients try to capture regime shifts, as emphasised by Hall, Mizon and Welfe (2000, p. 352).

In order to get a more parsimonious specification for our models, we adopted the following procedure: if a variable does not Granger-cause any other variable in the VeqCM, and at the same time is not Granger-caused by all other variables, then its equation is

---

[27] There are two exceptions to this rule: (1) the *Dd12p* variable is dropped (for obvious reasons) from the Polish VAR model; (2) the *Dwpc* (or *Dwpp*) and *Dprod* variables are restricted to *D(wpc-prod)* and *D(wpp-prod)*, as in the long run, in order to preserve degrees of freedom.

[28] Tables A1 and A2 in the appendix list all relevant data sources and the transformation of these variables.

dropped, and a new VEqCM with fewer equations (and fewer parameters to be estimated) is used as the general model. This is how we obtained the three VEqCMs for the Czech Republic, Hungary and Poland, the main features and diagnostic tests of which are summarised in Table 7 below.

The three models are quite similar from the point of view of their endogenous and predetermined-exogenous variables. The problem of the degree of freedom is clear if we consider that the number of parameters to be estimated is about 50% of the number of observations. There is no evidence of autocorrelation, heteroskedasticity (when the test is computable), and skewness; however, the rejection of the absence of the kurtosis hypothesis mainly depends on the outliers detected by previous inspections.

[29] "The typical market mechanisms need to be [further] introduced. Hence, prices will follow not only the cost-push, but their changes will also reflect the demand pressures" (W. Welfe, 1995, p. 497).

Tab. 7 – The main features of short term VEqCM

	Czech Republic	Hungary	Poland
endogenous:			
Dpc_t	x	x	x
Dpp_t	x	x	x
$D(wpc\text{-}prod)_t$	x	x	x
$Dr3_t$	x	x	x
Dmp_t	x		x
Dy_t		x	
lags	9	9	11
deterministic terms	*c*, *t*	*c*, *t*	*c*, *t*
predetermined and exogenous:			
wage gap$_{t-1}$	x	x	x
PPP CPI gap$_{t-1}$	x	x	x
PPP PPI gap$_{t-1}$	x	x	x
money gap$_{t-1}$	x	x	x
qr_{t-1}		x	
qr_{t-4}			x
$Dqr_{\text{from } t-1 \text{ to } t-4}$	x		
$Dpoil_{t-1}$	x	x	x
$Dpraw_{t-1}$	x	x	x
$Dyeuro_t$		x	x
$Dyeuro_{t-1}$	x		
period	1993.1-2000.6	1991.1-2000.7	1991.2-2000.7
number of observations	90×5	115×5	114×5
number of parameters	58×5	55×5	65×5
residual tests:			
- autocorrelation	no	no	no
- heteroskedasticity	-	no	-
- skewness	no	no	no
- kurtosis	yes	yes	yes
number of restrictions tested:			
Dpc_t equation [p-value]	26 [0.987]	19 [0.998]	29 [0.664]
Dpp_t equation [p-value]	33 [0.302]	24 [0.977]	27 [0.985]
overall [p-value]	59 [0.883]	43 [0.999]	56 [0.971]

Notes: D is the first difference operator. The labels of the variables are explained in Table A2.

The results, as a whole, appear quite satisfactory: furthermore, a number of not significant short-term parameters can be restricted to zero. As these restrictions imply the use of different sets of regressors in each equation within the multivariate system, higher efficiency can be achieved by using the SURE estimators, the zero restrictions being assessed by the Wald test. Test results, reported in the lower part of Table 7, suggest that all restrictions are not rejected at either the single equation level or in the two equations together.[30]

The last part of this section is going to look at the measurement and analysis of the feedback effects in the two restricted inflation equations, as they may be interpreted by economic theory and given that they refer to alternative inflation determinants. Remember that, even though we will report a subset of the results (those referring to inflation equations), we modelled inflation simultaneously with other macroeconomic variables, as shown in the first rows of Table 7. Restricted model SURE estimates are reported in Table 8.

Tab. 8 – Feedback effects of the restricted VEqCM in the PPI and CPI equations

	Czech Republic		Hungary		Poland	
	PPI equation	CPI equation	PPI equation	CPI equation	PPI equation	CPI equation
wage gap$_{t-1}$	0.005	0.017	0.071**	0.011*	0.017	-0.012
PPP CPI gap$_{t-1}$	[-]	0.099**	[-]	0.026**	[-]	0.127**
PPP PPI gap$_{t-1}$	0.017**	[-]	0.013	[-]	0.067**	[-]
money gap$_{t-1}$	0.009	0.058*	-0.149**	[-]	-0.016**	0.025**
qr_{t-1}	-	-	0.183**	0.052**	-	-
qr_{t-4}	-	-	-	-	-0.033	0.086**
Dqr_{t-3}	0.040	0.209**	-	-	-	
$Dpoil_{t-1}$	0.008**	-0.012	0.017*	0.011*	-0.002	0.002
$Dpraw_{t-1}$	0.018**	0.041*	[-]	[-]	-0.002	0.020
$Dyeuro_{t}$	-	-	[-]	0.169*	0.284**	0.167*
$Dyeuro_{t-1}$	0.063*	[-]	-	-	-	-

Notes: ** and * mean 1% and 5% rejection of the zero null. The labels of the variables are explained in Table A2. *D* is the first difference operator. "-" means "excluded from the model"; "[-]" means "zero restricted estimate".

The country-by-country analysis can be summarised as follows. In the Czech Republic, the foreign sector plays an important role in influencing the level of domestic inflation, especially in terms of producer prices: the rate of inflation of oil and raw material prices, foreign demand and the imbalances resulting from long-term PPP, all exert

[30] The reduction in the parameter dimension is quite relevant, as the restrictions imply a drop, ranging from 35% to 57%, in the total number of parameter to be estimated in each unrestricted equation.

considerable pressure, often 1% significant. On the other hand, wage costs, although positive, are not of significance in directing the short-term path of inflation. Demand-pull determinants are measured using money and output gaps; as far as CPI inflation is concerned, there is strong evidence of the effects of such factors, whereas PPI inflation does not appear to be significantly affected by them.

Foreign determinants of inflation also play an important role in Hungary, while demand-pull effects are relevant if measured by the output gap but not by the money gap (when significant, the money gap effect has the wrong sign). On the contrary to what happens in the Czech Republic, in Hungary the cost-push explanation for inflation (represented by the wage gap) is important. Overall, Hungarian inflation needs to be explained in terms of almost all the alternative (foreign, cost push, demand pull) determinants.

In Poland, cost-push explanations are not significant (whether labour costs or oil and raw materials), while foreign and domestic (output gap) demand pressures do constitute inflation determinants, at least in the case of the CPI equation. Furthermore, the money gap is significant, but has the right sign only in the case of the CPI inflation rate. As in the Czech Republic and Hungary, PPP gaps are significant.

An analysis of inflation determinants (comparing one country with another) shows that the effect of labour costs is only of importance (albeit limited) in the Hungarian case. This would support what we found in the weak exogeneity test results for the Czech and Polish cointegrated wage models (see Table 1 in section 3.1).

PPP gaps are always significant (with the exception of the Hungarian PPI equation); thus the real exchange rate is seen to be a key determinant of inflation in all three countries. The same is true of the capacity utilisation effect, in particular for the CPI equations.

The money gap only seems to be a more appropriate explanation of inflation when CPI measurements are of interest, whereas in other cases it is not significant or bears the wrong sign; the difficulties experienced in section 3.3 in identifying some stable form of relationship involving money, together with these results, would seem to that we ought to be particularly cautious about interpreting inflation using monetary developments.[31]

[31] Money instability is the reason why Golinelli and Rovelli (2001) prefer to pin down the Hungarian transmission mechanism without taking into account the evolution of monetary aggregates (see also Nemenyi,1997).

5. Conclusions

The present paper sets out to analyse the determinants of the inflation rate in the Czech Republic, Hungary and Poland. The methodological approach adopted in order to do so consists of two stages. During the first stage, we test for the existence of any long-term relationships by using the Johansen's cointegrated VAR method on subsets of variables of interest, and we try to solve the non-stationarity problem in terms of a number of stationary feedback effects (the equilibrium correction terms) that we call "gaps". During the second stage, a multivariate VEqCM approach is introduced in order to study inflation rate dynamics together with other macroeconomic variables, given the information nesting within both the short-term co-movements of the variables of interest (all variables, not only inflation) and the previously-identified gaps.

Though not optimal, this two-steps approach is inevitable, given the high number of parameters to be estimated and the relatively short time span available for analysis. Furthermore, we have to point out that any explanation of inflation patterns needs to take account not only of macroeconomic indicators, but also of the gradual administered liberalisation of prices and the removal of state subsidies. Given that these facts go beyond the focus of the present analysis, we decided to model both CPI and PPI inflation rates, in order to compare outcomes of alternative measures of inflation that are probably affected in different ways by such policy decisions.

Although previous drawbacks revealed the need for caution when drawing conclusions, our main findings can be laid out in the following six points.

Firstly, real wages, labour productivity and the vacancy-unemployment ratio form a stable, long-term relationship that can be interpreted as a target real wage. The gaps between the actual and the target real wage feed back to productivity (efficiency wage effect), while there does not appear to be any clear influence of gaps on short-term wages and inflation.

Secondly, long-term nominal exchange rates against the Euro follow a PPP (both in terms of CPI and PPI) whose trending path is the result of the Balassa-Samuelson effect. This means that, in the presence of productivity catch-up, the domestic CPI inflation rate in the three countries in question is expected to be larger (from 2% to 5%) than the Euro area inflation rate.

Thirdly, a stable long-term relationship between money, prices, output and interest rates is quite hard to find. The absence of money demand stability is understandable given the presence of substantial financial innovation and the mobility of capital flows. Our subsystem inferences suggest that money gaps explain more in terms of short-term real activity developments than in terms of domestic inflation rates.

Fourthly, the inflation rate may be forecast using econometric models employing economic theory-based indicators, although noisy monthly data and the considerable size of the parameter make it difficult to measure the effects of alternative inflation determinants using narrow intervals.

Fifthly, the exchange rate is the main long-term factor influencing domestic prices, and can be seen to be the common inflation-adjusting mechanism utilised in all three countries.

Sixthly, subsequent to progressive market liberalisation, demand-side pressure, measured in terms of the output gap, is seen to be a significant cause (mainly CPI) of inflation.

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Data appendix

Details of basic data sources are reported in Table A1. If “SA”does not appear in the description, then the time series are not seasonally adjusted; all raw data are seasonally adjusted by using Tramo/Seats procedure, see Gomez and Maravall (1997). Missing data in sub-periods were interpolated using an “interpolator” series, in order not to excessively shorten the sample.

Table A1 - Basic data

Label	Country	Period, from-to	Description	Source
W	Czech Republic	1991.5 -	Avg. monthly earnings, manufacturing, index 1995=1	MEI
	Hungary	1990.1 -	Earnings in industry, monthly avg., index 1995=1	Gd
	Poland	1990.1 -	Wages, avg. earnings, index 1995=1	IFS
Pc	Czech Republic	1991.1 -	Consumer Price Index (CPI), all items, 1995=1	MEI
	Hungary	1990.1 -	CPI all items, index 1995=1	MEI
	Poland	1990.1-1994.12;	Consumer prices (interpolator series)	IFS
		1995.1 -	CPI all items, index 1995=1	MEI
Pp	Czech Republic	1991.1	Producer Price Index (PPI), industrial products, 95=1	MEI
	Hungary	1990.1	PPI, industrial products, index 1995=1	MEI
	Poland	1990.1-1995.12;	Producer prices, industry (interpolator series)	IFS
		1996.1-	PPI, industrial products, index 1995=1	MEI
Y	Czech Republic	1990.1 -	Production, total industry (IIP), SA, index 1995=1	MEI
	Hungary	1990.1 -	IIP, SA, index 1995=1	MEI
	Poland	1990.1 -	IIP, SA, index 1995=1	MEI
N	Czech Republic	1990.1 -	Employees, manufacturing, thousands	MEI
	Hungary	1990.1-1992.12	Employment in industry (interpolator series)	GO
		1993.1 -	Employees, industry, thousands	MEI
	Poland	1990.1 -	Employees, industry, thousands	MEI
V	Czech Republic	1990.9 -	Unfilled vacancies, thousands	MEI
	Hungary	1992.1 -	Unfilled job vacancies, thousands	MEI
	Poland	1990.1 -	Unfilled vacancies, thousands	MEI
U	Czech Republic	1990.1 -	Registered unemployment, thousands	MEI
	Hungary	1990.1 -	Unemployment registered, thousands	MEI
	Poland	1990.1 -	Unemployment registered, thousands	MEI
$euro	United Sates	1990.1 -	Euro exchange rate, monthly averages	MEI
Pc^*	Euro area	1990.1-1994.1	CPI all items, 1995=1 (interpolator series)	GP
		1995.1 -	HICP, index 1995 =1	MB
Pp^*	Euro area	1990.1 -	Industrial producer prices, excl. construction, 95=1	MB
Y^*	Euro area	1990.1 -	Industrial production , excl. construction, 95=1, *SA*	ECB
N^*	Euro area	1990.1 -	Employment in industry, excl. construction (monthly interpolation of quarterly data)	ECB
Ex$	Czech Republic	1991.1 -	US dollar exchange rate, monthly avg., CK/$	MEI
	Hungary	1990.1-1990.12	US dollar exchange rate, monthly avg., F/$	IFS
		1991.1 -		MEI
	Poland	1990.1 -	US dollar exchange rate, monthly avg., Zt/$	MEI

(to be continued)

(continued)

M	Czech Republic	1992.1 -	Monetary aggregate M2, Billion CK	MEI
	Hungary	1990.1-1990.11	Quasi-Money monthly interpolation	IFS
		1990.12 -	Monetary aggregate M3, Billion F	MEI
	Poland	1990.1 -	Money supply M2, Million Zt	MEI
GDP	Czech Republic	1990 -	Gross Domestic Product (monthly interpolation with Y), Billion of 1995 CK	IFS
	Hungary	1990 -	Gross Domestic Product (monthly interpolation with Y) , Billion of 1995 F	MEI
	Poland	1990 -	Gross Domestic Product (monthly interpolation with Y) , Billion of 1995 Zt	MEI
R_{3m}	Czech Republic	1992.1 -	3-month PRIBOR	MEI
	Hungary	1990.1-1990.12	Treasury bill rate	IFS
		1991.1 -	90 day Treasury bill yield	MEI
	Poland	1990.1-1992.6	Treasury bill rate	IFS
		1992.7 -	3 month Treasury bill rate	MEI
Qr	Czech Republic	1991.6 -	Rate of capacity utilisation (%); linear mach last interpolation of quarterly data	MEI
	Hungary	1990.1 -	Rate of capacity utilisation (%); linear mach last interpolation of quarterly data	MEI
	Poland	1990.1 -	Rate of capacity utilisation (%); linear mach last interpolation of quarterly data	MEI
P_{oil}	World	1990.1 -	Crude oil, spot price index in $, 1995 = 1	IFS
P_{raw}	World	1990.1 -	Raw materials, all exports price index in $, 1995 = 1	IFS

MEI = OECD, Main Economic Indicators; IFS = IMF, International Financial Statistics; Gd = Macroeconomic and Financial Data Centre at the University of Gdansk; GO = Golinelli-Orsi (2000) monthly interpolated; MB = ECB, Monthly Bulletin; GP = Golinelli-Pastorello (2001);

Table A2 - Models variables

Label	Definition	Description
wpc	log(W/Pc)	real consumer wage
wpp	log(W/Pp)	real producer wage
prod	log(Y/N)	labour productivity (index)
vur	log(V/U)	vacancy-employment ratio
euro	log(Ex$ $euro)	nominal exchange rate (national currency against the Euro)
ppp_pc	*euro*+log(Pc*/Pc)	real exchange rate (in terms of consumer prices)
ppp_pp	*euro*+log(Pc*/Pc)	real exchange rate (in terms of producer prices)
prod_diff	*prod*-log(Y*/N*)	labour productivity relative to euro area (index)
mp	log(M/Pc)	real money balances
y	log(GDP)	real GDP at constant (1995) prices
r3	$R_{3m}/100$	nominal short-term interest rate
d12p	$\log(Pc/Pc_{-12})$	year-on-year inflation rate
qr	Qr/100	rate of capacity utilisation
Dpoil	$\log[(Ex\$\ P_{oil})/(Ex\$\ P_{oil})_{-1}]$	oil price monthly inflation
Dpraw	$\log[(Ex\$\ P_{raw})/(Ex\$\ P_{raw})_{-1}]$	raw material price monthly inflation
Dyeuro	$\log(Y^*/Y^*_{-1})$	monthly growth of Euro area industrial production