MONETARY POLICY TRANSMISSION, INTEREST RATE RULES AND INFLATION TARGETING IN THREE TRANSITION COUNTRIES

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Abstract

In 1991, the rate of inflation in the Czech Republic, Hungary and Poland was between 35% and 70%. At the end of 2001, it is below 8%. We setup a small structural macro model of these economies to explain the process of disinflation. Contrary to a widespread skepticism, which permeated a large part of previous research on these issues, we show that a simple open macroeconomic model, along the lines of Svensson (2000, Journal of International Economics), with forward-looking inflation and exchange rate expectations, can adequately characterize the relationship between the output gap, inflation, the real interest rate and the exchange rate during the course of transition.

We use the estimated models to interpret the main features of monetary policy in each country and identify the channels of policy transmission. We characterize the policy rules and assess the relative importance of the interest rate channel (on aggregate demand) and the exchange rate channel (which affects both aggregate demand and supply) in determining the path of (dis)inflation. In the same context, we also tentatively analyze the consequences of attempting a faster path of disinflation. Finally, we evaluate the appropriateness of the inflation targeting framework which has been adopted recently in all three countries, and discuss to what extent it represents a discontinuity with the past.

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***Keywords*:** Disinflation policy, Interest rate rules, Inflation targeting, Transition economies, Small open-economy macro models.

1. Introduction

During the last decade, several economies in transition have achieved both monetary and real stabilization. In many cases, this has been a remarkable story of success. In this paper we study the process of disinflation in three countries: the Czech Republic, Hungary and Poland (CHP). Table 1 illustrates some essential aspects of the stabilization process.[1] Partly in retrospect and partly looking forward, it seems natural to pose three questions: (i) how did monetary stabilization take place? (ii) could it have been faster or, instead, would a more gradualist approach have reduced the costs of stabilization? (iii) once the process of disinflation has been completed, how should the strategy of monetary policy be adjusted, in particular to take into account the aspiration to EU and EMU membership?

Table 1

	Czech Republic		**Hungary**		**Poland**	
(Annual averages)	1991	2000	1991	2000	1991	2000
Real GDP growth	-11.6	+2.9	-11.9	+5.2	-7.0	+4.0
Unemployment Rate	4.1	8.8	9.8	5.7	11.8	15.0
Inflation (CPI)	56.6	3.9	35.0	9.8	70.3	10.1
Central Govt. Surplus /Deficit	-2.0	-2.4	-4.6	-2.9	-3.8	-2.2
Curr. Units per euro *(Δ% 1999.10-2001.10)°*	-6.8		-0.85		-14.8	

° minus sign = appreciation of local currency
Source: Focus on Transition (2001) and National Central Banks.

On the face of it, the first question is the hardest to answer. Re-constructing and interpreting the history of disinflation in those countries may seem an adventurous task. Reading the official policy reports of some central banks during the past decade, one often encountered

[1] Of course, macro data only tell the surface of the story. Underlying the macro framework, there is a parallel and deeper story of structural changes, towards becoming "functioning market economies". In this respect, the recent "Strategy paper on enlargement" of the EU (http://europa.eu.int/comm/enlargement/report2001/index.htm) includes CHP among the eight "functioning market economies ... [which] should be able to cope with competitive pressure and market forces within the Union in the near term".

statements implying that: (i) monetary aggregates behaved unpredictably, (ii) the relation between money and growth was unpredictable; (iii) price indexes were unreliable and essentially responded to domestic cost pressures ... (iv) in addition to being heavily distorted by changes in administrative prices; (v) changes in interest rates did not significantly affect in a negative way domestic demand ... (vi) but might instead induce undesirable appreciations of the exchange rate ... (vii) thus fuelling into inflation (*sic*) either through the induced currency inflows or (viii) via wealth effects and aggregate demand pressures. What might monetary policy do in such circumstances? Our answer is that things were not necessarily as bleak as they were described, and that throughout the years of transition monetary policy has been a powerful and reliable tool for controlling aggregate demand. Thus, we build an "orthodox" model[2] of monetary policy, focusing in particular on the role of interest and exchange rates in the transmission mechanism, from 1991[3] to 2000. We show that such a model adequately characterizes the path of macro variables in CHP, and that the estimated equations are "well behaved" and reasonably stable across time. By appropriately setting a moderately restrictive path for nominal and hence for real interest rates and thus inducing a controlled appreciation of the real exchange rates, monetary authorities have been able to gradually steer their economies towards single-digit inflation, while containing the output costs of the stabilization. On the basis of the estimated model, we are then also able to answer the two remaining questions, on the choice of the speed of disinflation and how to adjust the policy strategy once disinflation has been completed.

The paper is organized as follows. Section 2 surveys the process of macroeconomic stabilization in the three countries. Section 3 presents a small structural macro model of an open economy, which is the starting point for the empirical part of the paper. The estimated models for the three countries are presented in section 4, and out of sample simulations in section 5. Section 6 concludes.

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[2] We use "orthodox" in reference to models were price-level determination embodies some degree of stickiness and hence inflation is determined by the interplay of aggregate demand and supply, as in the theoretical models proposed e.g. by Buiter (1980) and Svensson (1997; 2000), and the derived empirical models of Rudebusch and Svensson (1999), and Favero and Rovelli (2002).

[3] 1993 for the Czech Republic.

2. Monetary strategy and stabilization policy in the three countries

2.1 Three roads, one ending

For the best part of the last decade, the three countries examined have been actively managing their exchange rates, although in different ways.[4]

The **Czech Republic** (until the end of 1992, Czechoslovakia) introduced a currency peg in 1991, which it formally abandoned at the end of 1997, after a series of speculative attacks. Currently, the exchange rate is in free float, although the euro is informally used as a reference currency. In December 1997, the Czech National Bank (CNB) began to announce inflation targets[5]. Throughout the period, the exchange rate vis-à-vis the DM has been almost constant in nominal terms: it was 18.3 koruna to 1 DM in January 1991, and 17.0 in November 2001, with a maximum revaluation at 16.9 in February 1997, and a minimum at 19.5 in December 1997 (monthly averages). The real effective exchange rate, accordingly, has been under constant appreciation. In the CNB own calculations, it appreciated by 29.5% in CPI terms between January 1993 and November 2001, and 13.9% in PPI terms. The explicit adoption of inflation targets has not apparently had any impact on the behavior of the exchange rate (apart from, possibly, helping to diffuse speculative pressures).

Hungary adopted a crawling band since March 1995, with the rate of depreciation being gradually reduced as disinflation proceeded. The rate of crawl, initially at 1.9% per month, was gradually reduced to 0.3% per month in April 2000, and finally to 0.2% in April 2001. At almost the same time (May 2001) the fluctuation band was widened (±15%), rendering *de facto* irrelevant the rate of crawl. Immediately after, following also the approval of a new Bank Act (July 2001) the National Bank of Hungary (NBH) has begun to announce quantitative targets for inflation.[6]

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[4] Recent surveys and interpretations of this period, with complementary perspectives, are Bofinger and Wollmershaeuser (2000), Kutan and Brada (2000), Orlowski (2000),

[5] Initially, the target was formulated with reference to "net inflation", calculated by the Czech Statistical Office (CSO) as the growth of prices in the unregulated part of the consumer basket adjusted for changes in indirect taxes and for abolition of subsidies. In April 2001, the target was redefined as "headline inflation" (i.e. growth in the total consumer price index). Targets are set for several years ahead, currently until 2005.

[6] The first "Inflation Report" of the NBH was actually published in 1998.

Also **Poland**, similarly to the Czech Republic, started off adopting a currency peg in 1990. The commitment to fix the exchange rate did not last long, and in October 1991 Poland moved to a crawling peg system. The pre-announced monthly rate of crawl was gradually reduced, and then also transformed (May 1995) into a crawling band regime, with a wide (±7%) band. The band was further widened in the next years (up to 15% in March 1999) and finally abandoned, giving way to a free float, in April 2000. While pursuing its exchange rate policy, the National Bank of Poland (NBP) nevertheless defined its official strategy in terms of other variables: monetary targets until 1997; direct inflation targeting since October 1998.

Summing up, currently, the exchange rate regime of the three countries may be described as a free float (although the ±15% band is still valid, though obviously not binding, for Hungary) together with inflation targeting. This raises the question, to what extent the new policy framework signals a *discontinuity* with the past? We turn to this issue in the next sub-section.

2.2 Inflation Targeting: is it a break from the past?

First, we must define what we mean by inflation targeting (IT). Recent literature has analyzed in depth both the pre-requisites and the specific characterization of an IT policy strategy, also in the context of emerging economies. Mishkin (2000) suggests that an IT strategy encompasses five main elements: (i) announcement of targets, (ii) institutional commitment to price stability, (iii) information-inclusive strategy (iv) transparency of strategy (v) accountability of central bank for attaining the stated objectives. Amato and Gerlach (2001) observe that these criteria, while they capture the spirit of IT, "are not very helpful in formally defining this policy framework" and suggest that "IT is best thought of as a range of strategies". This observation is correct, and we think that it should not be interpreted as a criticism of Mishkin. The point is that almost all "strategies" which are not formulated in terms of a stringent rule[7] should best be thought of as defining the policy *framework* chosen by the central bank with respect to other macro agents (the government, financial markets, price and wage setters). For instance, as von Hagen (1999, p.699-700) succinctly states with respect to the monetary-targeting strategy of the Bundesbank: "Money growth targeting served the Bundesbank a number of politico-economic functions: It marked the end of the old regime where the Bank was powerless to control monetary conditions in Germany; it defined the

[7] We can think of only three examples of *stringent* rules: a k-percent rule of money growth; a commitment to fix the exchange rate around a constant value, with a *narrow* band; the adoption of a currency board. All other policy strategies, even if they can be formulated in terms of rules, always

central bank's monetary policy goal and its role in the macro economic policy game; and it served as a focal point in council meetings, strengthening the pursuit of a consistent monetary policy geared at price stability over time". We suggest that IT strategies for emerging countries, and for CHP in particular, have analogous politico-economic functions. In particular, the adoption of an IT strategy by CHP may be interpreted as a clear *signal* that:

1. the central bank is committed to price stability
2. the goal of price stability is in close sight
3. the central bank is independent of the government and other actors on the macroeconomic scene, who might in their turn be engaged in the pursuit of other goals[8]
4. the financial position of the government is sustainable without recourse to monetization[9]
5. the exchange rate policy is not (or not any longer) to be seen as the main anchor (or intermediate target) of the disinflation policy - in particular, since the latter has been almost completed, the pre-announcement of a rate of crawl no longer plays a central role in the strategy for disinflation[10], which in turn is a pre-requisite that ...
6. inflation forecasts have become the main intermediate target, the announcement of which helps to anchor inflation expectations
7. the central bank will adopt an "all-inclusive" information stance, looking at the determinants of inflation within a structured and broad analytical framework, without placing specific emphasis on the role of monetary aggregates[11]

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involve a considerable discretion in the processing of upcoming information, and hence also with respect to the *timing* of policy moves and the quantitative *setting* of policy instruments.

[8] The independence of the three central banks is assessed in Cukierman, Miller and Neyapti (2001).

[9] Hochreiter and Rovelli (2001) discuss this point in greater detail. After examining the relations between government finances and monetary policy in CHP, they conclude: "In all three countries, seigniorage substantially declined in the observation period [1992-99]. It is now negligible in the Czech Republic; around 1% of GDP in Hungary. Although it has increased to almost 4% of GDP in Poland, this is a temporary effect due to the attempt to reduce inflation during 1998 by increasing the level of domestic real rates of interest. Thus, in general, all our data indicate that monetary discipline and independence have been strengthened during the period which we have examined."

[10] For instance, in Hungary "the announcement of the rate of devaluation helped in developing the forward-looking nature of expectations" (NBH, 1999a, p. 17) and "by periodically lowering the devaluation rate the mechanism has been contributing successfully to the gradual reduction in the inflation differential between Hungary and its trading partners..." (NBH, 1999b, p. 13). See also Szapary and Jakab (1998).

[11] An examination of the Inflation Report published by each central bank would give support to this statement (as well as to points 1, 2, and 6 in the above list).

We think that these seven "signals" convey the role and purpose of IT in most emerging countries, and also with specific reference to CHP.[12] However, for those signals to be meaningful, some prerequisites must be satisfied. Amato and Gerlach (2001) list four such preconditions: (i) central bank independence; (ii) sound fiscal policy; (iii) resiliency of the economy to changes in interest and exchange rates; (iv) need for econometric models of the inflation process and the transmission mechanism. They also point out, however, that these preconditions "apply equally well to *any* other monetary policy framework" (id., p.5). This observation should imply that, to the extent that monetary policy has been successful in bringing about disinflation, the above prerequisites must have been satisfied also in the past, at least to some extent. In fact, this seems to have been the case: as a proof *by contradiction*, for instance, several authors have pointed out that temporary setbacks in the process of stabilization have occurred at times when fiscal policy has behaved inconsistently, and central bank independence has been side-stepped.[13]

But what about the econometric evidence on the inflation process and the transmission mechanism? On the one hand, it is easy to point out that the size and sophistication of money and financial markets, and the composition and quality of bank portfolios, have improved considerably in the course of transition. Thus the channels of transmission must have evolved somewhat. Also, the productive sector has been subject to considerable restructuring, so that in principle one might expect to find some evidence of instability in the equations that reflect the pricing-setting behavior of producers or their sensitivity to the cost of capital. On the other hand, it is also true that in each of the three countries the basic structure of money markets was already in place at the beginning of the 1990s, and interest rates have been liberalized towards the beginning of the decade. Hence, the extent to which the macroeconomic features of the transmission mechanism have remained stable, in a broad sense, throughout this period is largely an empirical question. In a previous paper on Hungary (Golinelli and Rovelli, 2001), we conducted several tests of parameter stability, based upon a small structural model, similar to the one presented in the next section. Our conclusion was that most coefficients do not change significantly when estimated over different sub-samples within 1991-99, and in particular that "the crucial parameters of the exchange-rate and interest

[12] Kutan and Brada (2000) express a different opinion on this issue. However their paper examines data and facts only up to the first half of 1999, hence it is possible that some of their reservations would not apply any longer today.

[13] For instance, see Begg (1998) for the fiscal policy reasons behind the Czech crisis of 1997.

rate equations are clearly stable across the two sub-samples".[14] In the rest of the paper, we shall propose and evaluate three parallel small macro models of the transmission mechanism, which will allow us to evaluate the main aspects of the transmission mechanism in CHP. This will help us to fulfil three tasks:

- to characterize the process of disinflation which has taken place in the last decade
- to understand to what extent the adoption of IT represents a discontinuity with the previous strategies of monetary policy
- to assess whether all the preconditions for a successful strategy of IT are in place in CHP.

3. Model specification

Our approach to modeling the transmission mechanism of monetary policy may be synthesized as follows. By controlling the evolution of nominal interest rates (for given foreign rates, and taking into account the required risk premium) the central bank also influences the path of the nominal exchange rate and may induce a pressure towards a desired degree of real appreciation.[15] Both directly (via the chosen level of the real interest rate) and indirectly (through the induced real exchange-rate appreciation), the central bank may thus influence the evolution of aggregate demand and of the inflation rate. As a result, taking into account the lag structure with which these effects take place in practice, a policy of persistent high real interest rates will lead to a process of gradual disinflation. This view of the transmission mechanism places domestic interest rates at the center of the analysis, and is compatible in principle with different strategies concerning the exchange rate policy. This raises the question of how to describe the policy rule(s) for setting interest rates, which the monetary authorities have followed in each country during the last decade. In particular we suggest that the choice of one or the other policy (or regime) with respect to the exchange rate,

[14] In less structured modeling frameworks, Siklos and Abel (2001) for Hungary and Christoffersen, Sloek and Wescott (2001) for Poland provide evidence in favor of a significant relation between inflation and monetary instruments. Also, Orlowski (2001) provides some preliminary evidence on the effectiveness of the adoption of IT strategies in changing the dynamics of the inflation process in the Czech Republic and Poland.

[15] Normally, one would expect that the rate of real appreciation is maximum under policies that peg the nominal exchange rate, and null under policies that result in a nominal depreciation equal to the inflation differential. Occasionally, both boundaries may be overshot. This happens either because (despite the inflation differential) the *nominal* exchange rate *appreciates* under speculative pressures, or when it *depreciates faster* than inflation, as during a currency run. See also the discussion in Bofinger and Wollmershaeuser (2000).

will influence the relative importance of different channels of transmission, but will not alter in a fundamental way the structure of the underlying macroeconomic relations.[16]

According to this view, we follow the modeling strategy suggested by Svensson (2000) for a small open economy. Our model includes four structural equations:

- aggregate supply, normalized in the rate of inflation (AS)
- aggregate demand, normalized in the output gap (AD)
- uncovered interest parity, augmented by a risk premium, normalized in the current exchange rate (ER)[17]
- a policy rule for setting the domestic interest rate (IR).

External shocks to the model originate for foreign prices, foreign demand and foreign interest rates. Equations AS, AD, and IR are specified with general lag structures. All four equations are affected by different white noise shocks ε_j (j = 1, 2, 3, 4). In the estimation, lags have been selected by imposing data admissible restrictions (see section 4).

In the AS equation (1), domestic inflation (Δp) is driven by the long-run purchasing power parity condition $[e + p^* - p]$, where e is the log-level of the nominal exchange rate, p^* and p are the foreign and the domestic log-levels of prices. In the short-run, domestic inflation is also affected by the output gap (y) and the foreign inflation rate:

$$\Delta p_t = \alpha_0 + \alpha_1(L)\,\Delta p_{t-1} + \alpha_2(L)\,\Delta p^*_t + \alpha_3(L)\,y_{t-1} + \alpha_4(L)\,[e_{t-1} + p^*_{t-1} - p_{t-1}] + \varepsilon_{1t} \tag{1}$$

[16] This point is well expressed also in a recent quotation from the NBH, which also stands in sharp contrast to the skepticism which emanated from many previous statements from the same source:

> "The objective of monetary policy is to achieve and maintain price stability. The central bank's primary instrument in attaining this goal is to change its benchmark rates. In this way, the Bank can influence inflation partly through the exchange rate's direct disciplinary power over price increases and partly through the effect of the real exchange rate and real interest rates on aggregate demand. In small countries such as Hungary, the exchange rate channel is the central bank's most powerful and fastest means of influencing domestic prices. Therefore, the exchange rate will retain its prominent role even though the Bank has much smaller power over the path of the exchange rate than under the previous narrow-band regime. In the future, if the National Bank wishes to influence the exchange rate in support of the inflation target, it will do so primarily by changing interest rates". (*National Bank of Hungary, Quarterly Report on Inflation, August 2001, pp.35-36).*

[17] Strictly speaking, this specification may apply also to the case of an exchange rate peg with a given fluctuation band.

In the AD equation (2), the output gap is linked to the expected real interest rate $[i - E\ \Delta p]$, the real exchange rate, the growth of world demand (Δw), and a measure of the fiscal impulse (bg):

$$y_t = \beta_0 + \beta_1(L)\, y_{t-1} + \beta_2(L)\, [i_{t-1} - E_{t-1}\, \Delta p_t] + \beta_3(L)\, [e_{t-1} + p^*_{t-1} - p_{t-1}] + \beta_4(L)\, \Delta w_t + \beta_5(L)\, bg_{t-1} + \varepsilon_{2\,t} \qquad (2)$$

The ER equation (3) is based on the UIP hypothesis, augmented with a stationary risk premium, φ_t, modeled as an AR1 process:

$$e_t = E_t\, e_{t+1} - (i_t - i^*_t) + \varphi_t\,, \quad \text{were: } \varphi_t = \gamma_0 + \gamma\, \varphi_{t-1} + \varepsilon_{3\,t} \qquad (3)$$

The IR equation (4) defines the nominal interest rate differential $(i - i^*)$ in terms of a forward-looking dynamic[18] Taylor rule ; ε_4 are random shocks to the interest rate differential.

$$i_t - i^*_t = \delta_0 + \delta_1(L)\, (i_{t-1} - i^*_{t-1}) + \delta_2\, E_t\, (\Delta p_{t+1} - \Delta p^*_{t+1}) + \delta_3\, E_t\, y_{t+1} + \varepsilon_{4t} \qquad (4)$$

All equations are jointly estimated and simulated with model-consistent expectations. If the variables in each equation are stationary or cointegrated, then we may obtain statistical estimates of the steady state solution of the model, conditional on the steady state values of the foreign variables.

In the out-of-sample simulations, we study the path to nominal convergence. To ensure convergence, we impose the following long-run restrictions of the rate of domestic inflation, on the nominal exchange and interest rates and on the risk premium:

$$\Delta p = \Delta p^* \qquad \Delta e = 0 \qquad i = i^* \qquad \gamma_0 = \delta_0 = 0$$

Notice that this also implies a long run condition on the *constancy* of the real exchange rate.[19]

4. Estimation method and results

The data set for each country includes seven variables, measured quarterly from 1991.1 to 2001.1 (1993.1 to 2001.1 for the Czech Republic). All variables except interest rates and the

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[18] Goodfriend (1987) and Walsh (1998, ch.10) rationalize the inclusion of autoregressive parameters in interest rate rules with reference to the authorities' desire to smooth the path of interest rates.

[19] In addition, we shall also need to specify a terminal condition on the *value* of the exchange rate, in order to be able to solve forward the ER equation. More on this in the next section.

ratio of capacity utilization are in logs. We use the CPI for domestic prices, *p;* and the capacity utilization ratio to measure the output gap, *y.*[20] Inflation (Δp_t , Δp^*_t) and exchange rate changes are measured as annualized quarterly changes; for output we use quarterly changes. The domestic interest rate, *i,* is the three-month Treasury Bill rate. The nominal exchange rate, *e*, is the weighted average of the exchange rate of each country against the DM (or ECU) and the USD, with variable weights chosen according to the objectives or targets indicated by each national central bank. Foreign prices, p^*, are the weighted average of the German and US CPI, with weights chosen according to the same criterion as for exchange rates; similarly, the foreign interest rate, i^*, is the weighted average of German and US short-term rates (respectively, the call money rate for Germany and the rate on 3-mos. CD for the US). World demand, *w,* is proxied by the index of world trade in manufactures. Data sources and more detailed definitions of all variables are reported in Appendix 1.[21]

The model is estimated by three stages least squares (3SLS).[22] Hsiao (1997) shows that the structural equation approach for estimation and testing is still valid, even when some regressors are non stationary, provided they are cointegrated[23]. In Appendix 2 we show that these requirements are satisfied for all the non-stationary variables in the set of regressors.

The estimated equations track the data quite well, and their statistical performance is satisfying, despite the short sample of observations. At a general level, a strong check on the validity of the model is the fact that very similar specifications emerge independently for the three countries in our data set. In the Appendix 3 we report results of the mis-specification and stability tests, and plots of the estimated equations.

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[20] The time series properties of capacity utilization are quite similar to those of an HP filter on GDP. In particular, the two variables indicate a very similar evolution of the business cycle. These comparisons are available on request from the authors.

[21] Notice that Δ indicates the difference between two consecutive quarters and Δ_4 indicates the difference between the same quarter in two consecutive years. In all cases, inflation rates are defined on an annual scale, i.e. $\Delta p_{t-k+1} \equiv 4\ (p_{t-k+1} - p_{t-k})$. E_{t-k} measures expectations formed at time t-k.

[22] The alternative GMM estimator brings efficiency gains in presence of heteroskedasticity and autocorrelation but asymptotically it becomes equivalent to 3SLS if the disturbances are white noise (see the diagnostic tests on the residuals, in Appendix 3). In addition, if the disturbances are also normally distributed, 3SLS has the same asymptotic distribution as the (asymptotically efficient) full information maximum likelihood estimator. Finally, GMM methods are more data demanding than 3SLS; in our case they are often infeasible because of small samples.

[23] Hsiao (1997, p. 395) concludes that “in a structural approach one still needs to worry about the issues of identification and simultaneity bias, but one need not worry about the issues of non-stationarity and cointegration.”

For ease of comparison across countries, we discuss the results of estimation for each equation in turn. In Table 2 we show the estimates of the aggregate supply equation. For each country, the pressure on domestic inflation from aggregate demand and the real exchange rate are well identified.

A homogeneity condition has been imposed in each equation, to ensure that domestic inflation converges to the foreign rate in the long run. This restriction is easily not rejected for the Czech R. (χ^2 = 0.431, P-value = 0.512), and Hungary (χ^2 = 0.579, P-value = 0.447). For Poland, the restriction is rejected at 5% but not rejected at 1% (χ^2 = 5.04, P-value = 0.025). For the Czech R. we may also identify the *level* of long-run capacity utilization at which inflation converges to the foreign rate. For the other two countries, the condition for nominal convergence is simply $\Delta p_t - \Delta p^*_t = \Delta y_t = 0$, which may be satisfied at *any* constant rate of capacity utilization.

Similarly, note that for the Czech R. and Hungary, the *level* of the real exchange rate affects domestic inflation. For Poland, we find a direct link between inflation and changes in the nominal exchange rate. This implies that, in the long run solution for the Czech R. and Hungary, the inflation differential attains zero when the real exchange rate attains its long run equilibrium value, whereas for Poland the convergence condition is simply $\Delta p_t - \Delta p^*_t = \Delta e_t = 0$, which may be verified at *any* level of the real exchange rate.

We observe that the effect of the real exchange rate appreciation is much more direct (with an impact coefficient of 0.695 on inflation) in the case of the Czech Republic, which corresponds well to the fact that the CNB consistently placed a greater emphasis, relative to the other two central banks, on the role of real exchange rate appreciations as a means to disinflation.

In the AD equations (Table 3), we notice that shocks have a high persistence: the sum of the autoregressive coefficients of the lagged dependent variables is in each case around 0.85. The effects of the real interest rate and of changes in the real exchange rate and in world demand are in each case well identified and significant. However, real interest rates affect AD in the Czech Republic only with a very long lag (first impact is after 5 quarters) whereas the competitiveness effect takes place with a shorter lag (3 quarters, same as in the other two countries) and is much stronger than in the other countries. Again, this corresponds well to the different exchange rate strategy adopted by the CNB. Finally, in the case of the Czech Republic we also found a small, significant effect of the government budget variable on AD.[24]

Notice that, on the basis of the estimated AD equations, we may also compute the level of capacity utilization, which prevails in the long run, as a function of the real interest rate and of the rate of growth of world demand. This will be required for the simulations of the model.

In Table 4 the current exchange rate is shown to depend on the current interest rate differential and on the expected future exchange rate, augmented by a risk premium. The AR1 coefficient on the risk premium varies between 0.45 in Hungary and 0.75 in the Czech Republic. Note that there is no constant component in the risk-premium for the Czech Republic. This may plausibly be linked to the fact that for a large part of the sample the CNB was successfully targeting the nominal exchange rate.[25]

In the estimation, we imposed the restriction $\gamma_0/(1-\gamma) = \delta_0/[1-\delta_1(1)]$ between equations (3) and (4). This implies equality in the long run between the risk premia in the UIP and interest rate equations. The restrictions are never rejected.[26]

[24] *bg* measures the government balance (a positive value equal surplus) on GDP. Comparable data are not available on a quarterly basis for the other countries.

[25] For this reason, we also estimated an alternative version of the model, setting the exchange rate as an exogenous variable. Estimates of all the parameters in the other three equations are virtually identical. Results are available on request. Notice however that exchange rate targeting by the CNB has been suspended at the end of 1997.

[26] For the Czech Republic: $\chi^2 = 2.087$ (p-value = 0.149). The additional restriction of both coefficient being equal to zero is also not rejected: $\chi^2 = 4.62$, p-value = 0.099. For Hungary: $\chi^2 = 0.033$ (p-value = 0.855). For Poland: $\chi^2 = 1.706$ (p-value = 0.191). For Hungary and Poland we reject the zero coefficient restrictions; hence, in these cases we have a constant component of the risk premium, which is between 0.5% and 0.9% per quarter. These results are embodied in the estimates in Table 5.

Table 2. Aggregate supply (Equation 1)

Czech R.

$$\Delta p_t = -1.227 + \Delta p^*_t + 0.657\, y_{t-1} + 0.997\, y_{t-3} + 0.695\, (e_{t-1} + p^*_{t-1} - p_{t-1})$$

(0.318) (0.339) (0.317) (0.112)

Hungary

$$\Delta p_t = 0.055 + 0.872\, \Delta p_{t-1} - 0.596\, \Delta p_{t-4} + 0.392\, \Delta p_{t-5} + 0.332\, \Delta p^*_{t-2} + 0.600\, (y_{t-1} - y_{t-3}) + 0.277\, (e_{t-1} + p^*_{t-1} - p_{t-1})$$

(0.021) (0.091) (0.131) (0.113) (-) (0.18) (0.104)

Poland

$$\Delta p_t = 0.650\, \Delta p_{t-1} + 0.199\, \Delta p_{t-4} - 0.067\, \Delta p_{t-5} + 0.217\, \Delta p^*_{t-1} + 0.457\, (\Delta y_{t-1} - \Delta y_{t-3}) + 0.168\, (\Delta e_{t-1} + \Delta e_{t-3})/2$$

(0.042) (0.011) (0.013) (-) (0.158) (0.040)

Table 3. Aggregate demand (Equation 2)

Czech R.

$$y_t = 0.073 + 0.536\, y_{t-1} + 0.339\, y_{t-3} - 0.122\, [(i_{t-5} - E_{t-5}\, \Delta p_{t-4}) + (i_{t-6} - E_{t-6}\, \Delta p_{t-5})]/2 \; +$$

(0.061) (0.091) (0.086) (0.040)

$$0.211\, \Delta(e_{t-3} + p^*_{t-3} - p_{t-3}) + 0.223\, \Delta w_t + 0.238\, \Delta w_{t-4} - 0.163\, bg_{t-2}$$

(0.081) (0.045) (0.050) (0.110)

Hungary

$$y_t = 0.082 + 0.516\, y_{t-1} + 0.369\, y_{t-2} - 0.122\, [i_{t-1} - E_{t-1}\, \Delta_4 p_t] + 0.115\, \Delta_4(e_{t-1} + p^*_{t-1} - p_{t-1}) + 0.205\, \Delta_4 w_{t-1}$$

(0.036) (0.137) (0.128) (0.052) (0.046) (0.068)

Poland

$$y_t = 0.099 + 0.843\, y_{t-1} - 0.047\, (i_{t-1} - E_{t-1}\, \Delta p_t) + 0.025\, \Delta(e_{t-3} + p^*_{t-3} - p_{t-3}) + 0.133\, \Delta w_{t-2}$$

(0.034) (0.053) (0.020) (0.009) (0.041)

Table 4. Exchange rate (Equation 3)

Czech R.

$$e_t = E_t e_{t+1} - [i_t - i^*_t]/4 + \varphi_t, \quad \text{where:} \quad \varphi_t = \underset{(0.179)}{0.751}\, \varphi_{t-1}$$

Hungary

$$e_t = E_t e_{t+1} - [i_t - i^*_t]/4 + \varphi_t, \quad \text{where} \quad \varphi_t = \underset{(-)}{0.0091} + \underset{(0.152)}{0.445}\, \varphi_{t-1}$$

Poland

$$e_t = E_t e_{t+1} - [i_t - i^*_t]/4 + \varphi_t, \quad \text{where:} \quad \varphi_t = \underset{(-)}{0.0054} + \underset{(0.139)}{0.660}\, \varphi_{t-1}$$

Table 5. Interest rates (Equation 4)

Czech R.

$$\Delta(i_t - i^*_t) = \underset{(0.065)}{-0.188}\, [i_{t-1} - i^*_{t-1}] + \underset{(0.079)}{0.208}\, E_t\, [\Delta p_{t+1} - \Delta p^*_{t+1}]$$

Hungary

$$\Delta(i_t - i^*_t) = \underset{(0.066)}{-0.318}\, [i_{t-1} - i^*_{t-1} - \underset{(0.016)}{0.065}] + \underset{(0.116)}{0.528}\, \Delta[i_{t-1} - i^*_{t-1}] + \underset{(0.125)}{0.335}\, \Delta[i_{t-3} - i^*_{t-3}] + \underset{(0.065)}{0.228}\, E_t\, (\Delta p_{t+1} - \Delta p^*_{t+1})$$

Poland

$$\Delta(i_t - i^*_t) = \underset{(0.033)}{-0.303}\, [i_{t-1} - i^*_{t-1} - \underset{(0.014)}{0.063}] + \underset{(0.047)}{0.309}\, \Delta[i_{t-1} - i^*_{t-1}] - \underset{(0.001)}{0.006}\, \Delta[i_{t-3} - i^*_{t-3}] + \underset{(0.041)}{0.213}\, [\Delta p_t - \Delta p^*_t] - \underset{(0.028)}{0.224}\, \Delta[\Delta p_{t-2} - \Delta p^*_{t-2}]$$

In Table 5 the differential between domestic and foreign interest rates is shown to depend on an autoregressive component and on the expected inflation differential. For Poland, in particular, we found an equilibrium error-correction model, with the target interest rate differential linked to the expected future inflation differential, on the basis of the following relation: $i_t - i^*_t = 0.065 + 0.7\, E_t\, (\Delta p_{t+1} - \Delta p^*_{t+1})$.

In contrast to interest rate rules estimated for advanced industrial countries, we found that the rate of domestic capacity utilization does not help to explain the interest rate differential in any one of the three countries.[27]

As regards the measurement of the anti-inflationary stance, we notice that there are in general two possibilities for a non-accommodating policy. A traditional Taylor rule requires that interest rates react to inflation expectations with a coefficient higher than unity (at least in the long run). However, this does not apply to the case where a *persistent* policy of disinflation is required. For instance, if interest rates are set initially at a high enough level, relative to the rate of inflation, then the *less they react* to changes in the rate of inflation, the *more aggressive* the policy stance becomes. This observation clearly applies to all the three countries, as the estimates in Table 5 show.[28]

Detailed diagnostic tests for all equations, including tests of parameter constancy, are presented in Appendix 3. Here we briefly present (Table 6) only on the results of the stability tests.[29] The tests are reported for each equation and also for the full model. In general, evidence of instability is limited to the initial years in the sample. For the Czech Republic,

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[27] The fact that interest rates do not respond to the rate of capacity utilization (in reference to equation 4, $\delta_3 = 0$) does not imply that central banks neglect the real costs of stabilization policy. Concern for these costs is embodied in the choice of the speed of disinflation. That capacity utilization is not significant in the interest rate setting equations only implies that this policy is unchanged across the business cycle.

[28] For Hungary and Poland, this point may easily be seen in reference to the estimated constant component of the risk premium. If we arbitrarily constrain this parameter to zero in the estimation of equations 3 and 4, then the resulting estimates of the (long-run) effect of expected inflation on interest rates becomes greater than unity (not reported, available on request). Although as we mention in the text this restriction is formally rejected, its imposition would not appreciably change the remaining parameter estimates.

[29] Tests have been conducted adding to each equation one dummy variable for each quarter tested for stability. Testing for predictive failure is then equivalent to testing for the joint significance of zero-restrictions on the coefficients of the dummy variables. In this way we can test for predictive failure both in the first and last part of the sample. We present the p-values of the F-version of the chi-square Wald test, as the F statistic is expected to have better small sample properties than the (asymptotic) chi-square statistic. Notice that in any case the results of these tests have to be interpreted with caution, as Candelon and Lutkepohl (2001) show that, in the context of multiple time series systems, Chow-type tests have distorted size (i.e. they reject far too often the null hypothesis of stability).

instability originates from eqs. 1 (AS) and 4 (interest rate rule). For Hungary, from eqs. 2 (AD) and 4. For Poland, from eq 1. Only for Poland there is some evidence of instability also at the end of the sample, in particular for the year 1999, in eqs. 2 (AS) and 3 (ER).

Table 6. Stability Tests

Country	*Period*	*eq. 1*	*eq. 2*	*eq. 3*	*eq. 4*	*System*
Czech Republic	1993	0.0061	0.7548	0.9777	0.0080	0.0243
	2000	0.9920	0.6872	0.7635	0.9892	0.9845
Hungary	1991	0.3696	0.0024	0.0147	0.5503	0.0074
	1991-1992	0.3859	0.0002	0.1003	0.0370	0.0039
	1999-2000	0.9101	0.6205	0.9875	0.7247	0.9581
	2000	0.9431	0.9374	0.8462	0.4980	0.9646
Poland	1991	0.4642	0.5642	0.1834	0.0757	0.1848
	1991-1992	0.0108	0.7566	0.2801	0.2006	0.0756
	1999-2000	0.7383	0.0030	0.0470	0.9728	0.0652
	2000	0.9597	0.6085	0.1696	0.9142	0.7976

Notes. Reported values are *P*-values of the F version of the second Chow test, for the relevant period. Test for year 2000 include five quarters, up to 2001.1.
See Tables A.6.1, A.6.2, A.6.3 for other details.

Overall, while it is not surprising that we should find instability in the first years of the transition process it is instead quite comforting[30] to find that the estimated equations are in general remarkably stable in the most recent years.

[30] Also keeping in mind the results of Candelon and Lutkepohl (2001), mentioned in the previous footnote.

5. Simulation results

In order to examine the long run implications of the estimated models, we have simulated them outside the sample of estimation. Any solution to the model must take into account the forward-looking nature of the exchange rate in equations (3). This requires in each case the choice of an appropriate terminal condition. We have assumed that the real exchange rate converges, at the end of the solution period, to the long run value, computed on the basis of the reduced form, steady state solution. This allows us to find a consistent expectations solution to the model. In particular, the steady state solution is formulated in terms of the *level* of the real exchange rate for the Czech Republic and Hungary. For Poland instead the terminal condition only imposes that the real exchange rate becomes *constant*[31]. For the three countries, in addition, convergence of the iterative procedure is ensured by the existence of a steady state solution for the inflation rate, capacity utilization and domestic interest rates.

To simulate the model outside the period of estimation, we also need to forecast the three exogenous (foreign) variables. To this purpose, we assumed the following long run values: $\Delta p^* = 0.024$; $\Delta w = 0.072$; $i^* = 0.043$.[32]

Finally, for Hungary and Poland we also need to forecast the evolution of the constant component of the risk premium, γ_0 , outside the estimation sample. In the simulations reported below, we assumed a decay parameter equal to 0.6.[33]

We simulated the model from 2001.2 to 2020.4. [34] In Fig. 1.1 we plot the point forecasts from a deterministic simulations for the Czech Republic, in Fig.1.2 for Hungary and Fig. 1.3 for Poland. In each case we also add a forecasting interval (± one forecasting error).[35]

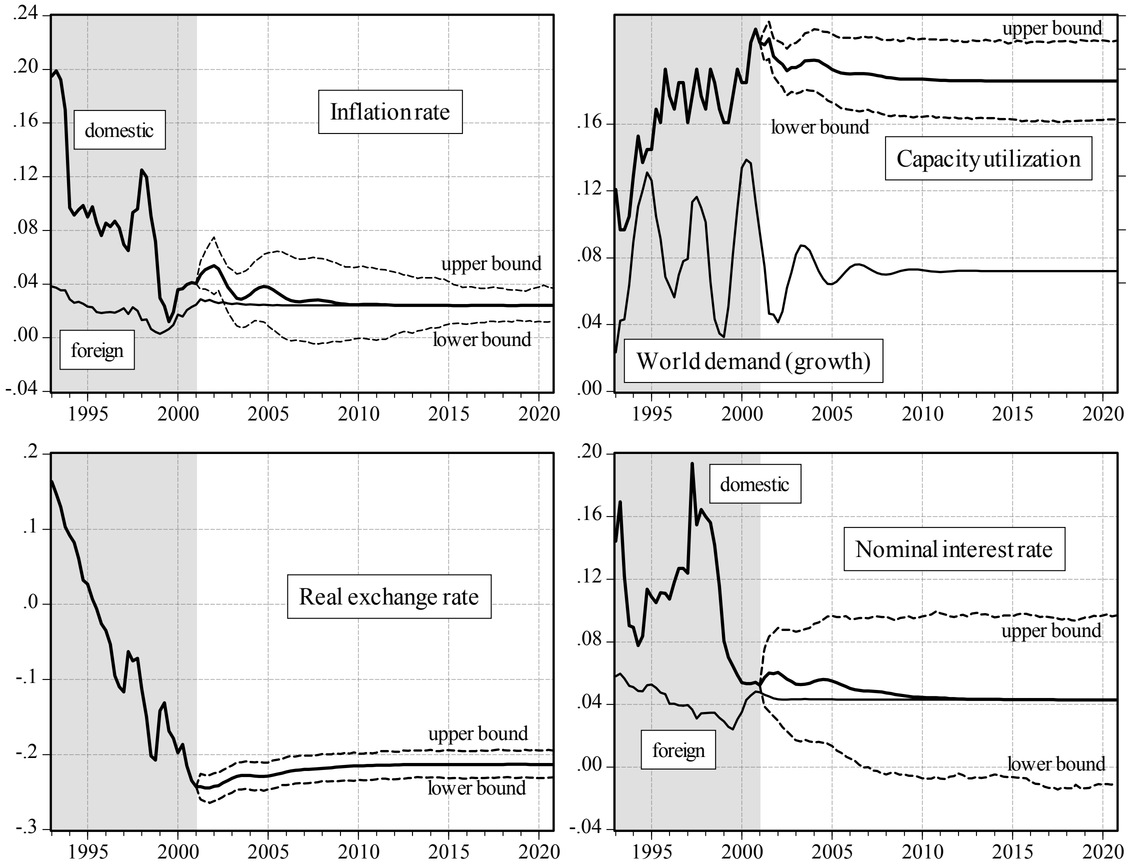
[31] The level is estimated at –0.213 for the Czech Republic, and –0.200 for Hungary. Note that in Golinelli and Rovelli (2001) the corresponding estimate for Hungary was –4.635, but this was due to a different normalization of the equation. The effective difference between the two terminal conditions for Hungary is about 7%. For Poland, the value at which the real exchange rate becomes constant is –0.340.

[32] This is the same as in Golinelli and Rovelli (2001). In addition, for the Czech model, we set the fiscal impulse, *bg* ,equal to zero in the long run.

[33] We did two checks on this assumption. First we assumed a decay parameter of γ_0 equal to 0.8. Second, we re-estimated the model imposing from the outset the restriction $\gamma_0 = 0$, and simulated the model on the basis of the new estimated parameters. In both cased the simulated path of the endogenous variables was very insignificantly different from the first simulations (results available on request).

[34] We performed the simulations with a Fair and Taylor (1983) scheme, without the automatic extension of the time horizon, using Eviews 4. The simulation results are unaffected when moving the end point of the simulation either to 2015 or forward.

Fig. 1.1 – History (shaded) and forecast, the Czech Republic



The three sets of the simulations attract similar comments. In general, we accept that our models do not yield particular insights on the evolution of output or of the real exchange rates in the long run. As we explained in Section 3, the steady state values of these variables are computed from the estimated parameters of the equations for aggregate demand (eq.1) and from the cointegrating vector for the real exchange rate (Tables A.5.1, -2, -3 in Appendix 2). This implies that we do not take into account the evolution of potential output over time, nor the effects of productivity on the real exchange rate. These limitations cannot be avoided in the context of our model.

We now move on to more specific comments. As regards the continuation of disinflation, for the Czech Republic, as the process has been almost completed, there is not much to be said,

[35] Intervals have been obtained performing a stochastic simulation of the model (1000 repetitions). The unknown equation errors were measured by pseudo-random values, extracted from a distribution with a covariance matrix from the estimated equations.

except that there might still be a few bumps at the end of the story. For Hungary and Poland the graphs obviously point out that disinflation is slightly slower, but the point forecast indicate that inflation should reach 4% by 2003-04 in each country, although the interval of the forecasts is wider for Poland.

We were instead surprised, in a negative way, that the path of nominal interest rates is in each case so imprecisely defined. For instance, the entire range of values between 0 and 9% is encompassed in the forecasting interval for the Czech Republic. Similar or slightly wider intervals apply to the other countries. As the interest rate rules were estimated with sufficient precision (Tables A.6.1, -2, -3 in Appendix 3), this behavior must probably be attributed to the high persistence of the autoregressive risk premium (See Table 4).

Fig. 1.2 – History (shaded) and forecast, Hungary

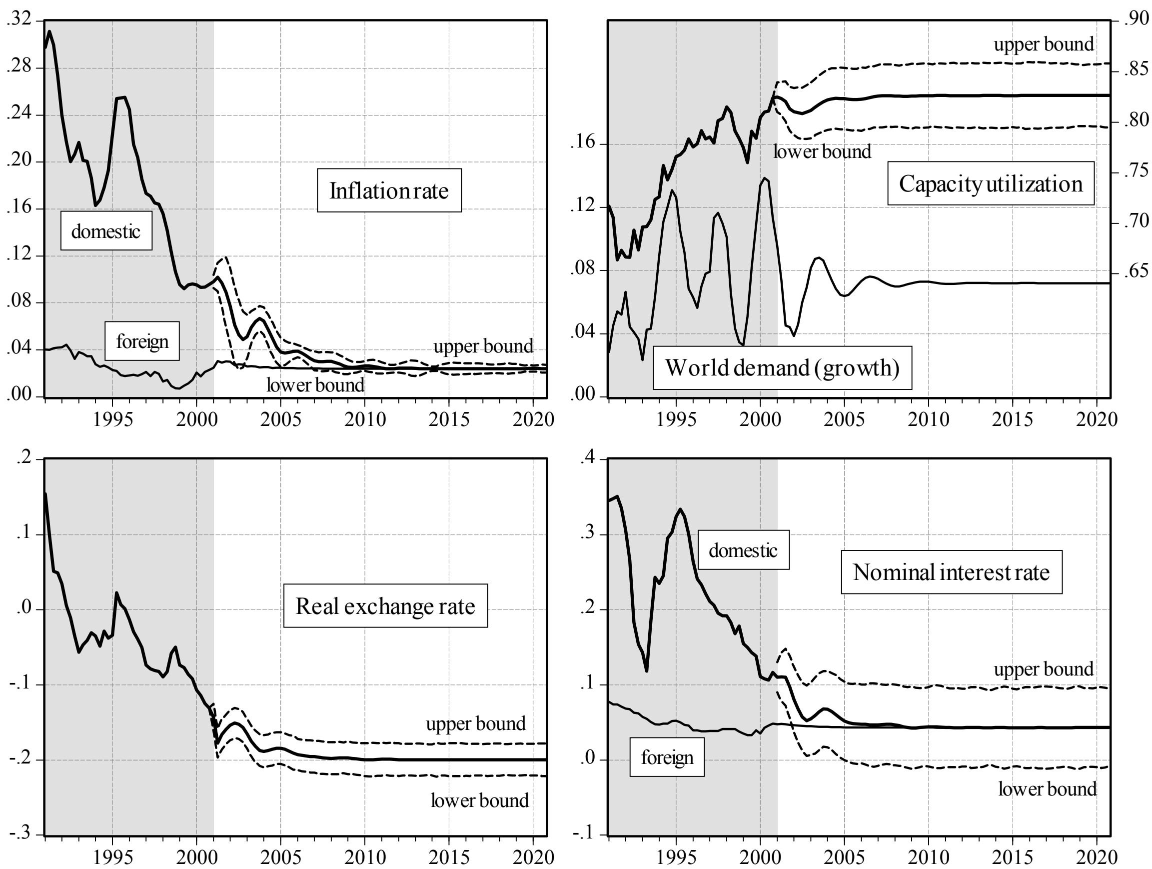
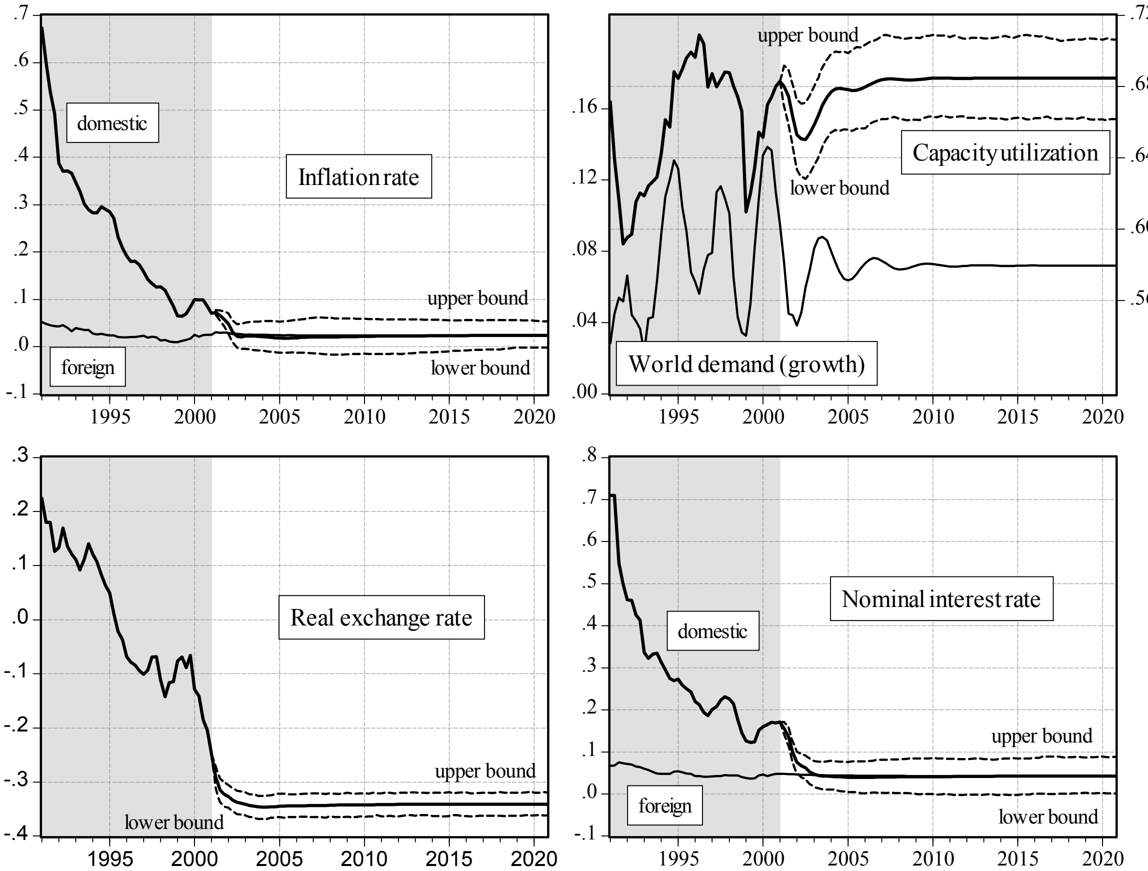


Fig. 1.3 – History (shaded) and forecast, Poland



5.1 A more aggressive policy stance

How would the path of the endogenous variables changed if the three central banks had conducted a more aggressive policy against inflation since 1993? To answer this question, we have compared the baseline simulation of the model from 1993 to 2020, with an alternative, where the response of nominal interest rates to the expected inflation differential has been increased relative to the estimated value. In particular, we set the long run elasticity of the nominal interest rate differential to the inflation differential to approximately 1.6 for all three countries.[36] The effects of this more aggressive policy stance on the variables of the model are plotted in Figures 2.1-2.3[37].

[36] In particular, we multiplied by 1.5 the impact parameter in the equation 4 for the Czech Republic and Hungary (which imply long-run effects respectively equal to 1.662 and 1.624), and by 1.8 the parameter for Poland (1.627 in the long-run).
[37] The right columns plots report the differences between the simulated baseline policy and the simulated *more aggressive* policy. Interval estimates of one standard error where obtained by running 1000 stochastic simulations.

Since the estimated model for Poland does not yield a direct solution for the level of the long-run real exchange rate, and the terminal condition is simply set at that level where the real exchange rate becomes constant, one consequence of simulating an alternative policy for Poland is that the terminal value of the real exchange rate will be changed accordingly. For this reason, the large increase in the "proactiveness" of the NBP towards the inflation differential implies a real appreciation of the exchange rate with respect to the baseline solution.[38]

Briefly, the other main results of these simulations are the following. For each country, we find that a more aggressive stance would deliver a benefit in terms of faster disinflation. For the Czech Republic, inflation would have been about 4% lower in 1995-96, but the subsequent resurgence of inflation in 1997 would not have been avoided. Thus, the induced lower levels of output (-1% for about five years) would appear as a net cost, with no permanent gain.

For Hungary, the gains on the inflation front are more pronounced and permanent (as inflation drops by about 12% in 1994-96), but the output costs are also large (-4% in 1994-95, and then gradually fading away, but not disappearing until 2001). Interestingly, most of the effects would seem to come through the appreciation of the real exchange rate, which approaches much faster its steady-state value[39].

For Poland, the behavior of inflation is similar to that of Hungary, but the costs to output are more short-lived, since in any case the economy ends up in the recession of 1998-99, which is almost identical in either policy scenario. In the more aggressive scenario, the path of nominal interest rates would have been almost unchanged (apart from the obvious initial increase in 1993-94). The real exchange rate would end up permanently more appreciated by about 1.5%.

Overall, these simulations point out that faster disinflation would have been possible, but that in each case it would have imposed additional output costs. We must stress that these results are the outcome of somewhat mechanical counterfactual simulations, which certainly are not immune to the Lucas' (1976) critique, and thus must be interpreted with caution. In particular, it is quite possible that some of the estimated parameters are not *deep,* i.e. they might not be

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[38] For Hungary and Poland, in these alternative simulations we used parameters estimated subject to the restriction that the constant risk premium is zero (See the discussion of this point in footnote 28). This is consistent with the fact that an increased proactiveness of the central bank would most likely increase credibility and hence result in reduced risk premia. Moreover, this avoids the need to set an arbitrary path of decay for the exogenous component of the risk premium.

[39] This effect is similar to that obtained in Golinelli and Rovelli (2001).

independent of the policy rule, which the monetary authority is expected to follow. For instance, it is possible that, if the central bank *credibly* announces a more resolute path of disinflation, the degree of persistence of inflation (the $\alpha_1(L)$ polynomial in eq.1) and of the risk premium (the autoregressive parameter, γ, in eq. 4) might have been lower. If this were the case, then it is also possible (but by no means unambiguously necessary) that the results of the simulations indicate an upper bound for the output loss resulting from a faster disinflation.

Fig. 2.1 – Simulation of a more aggressive policy stance, Czech Republic.

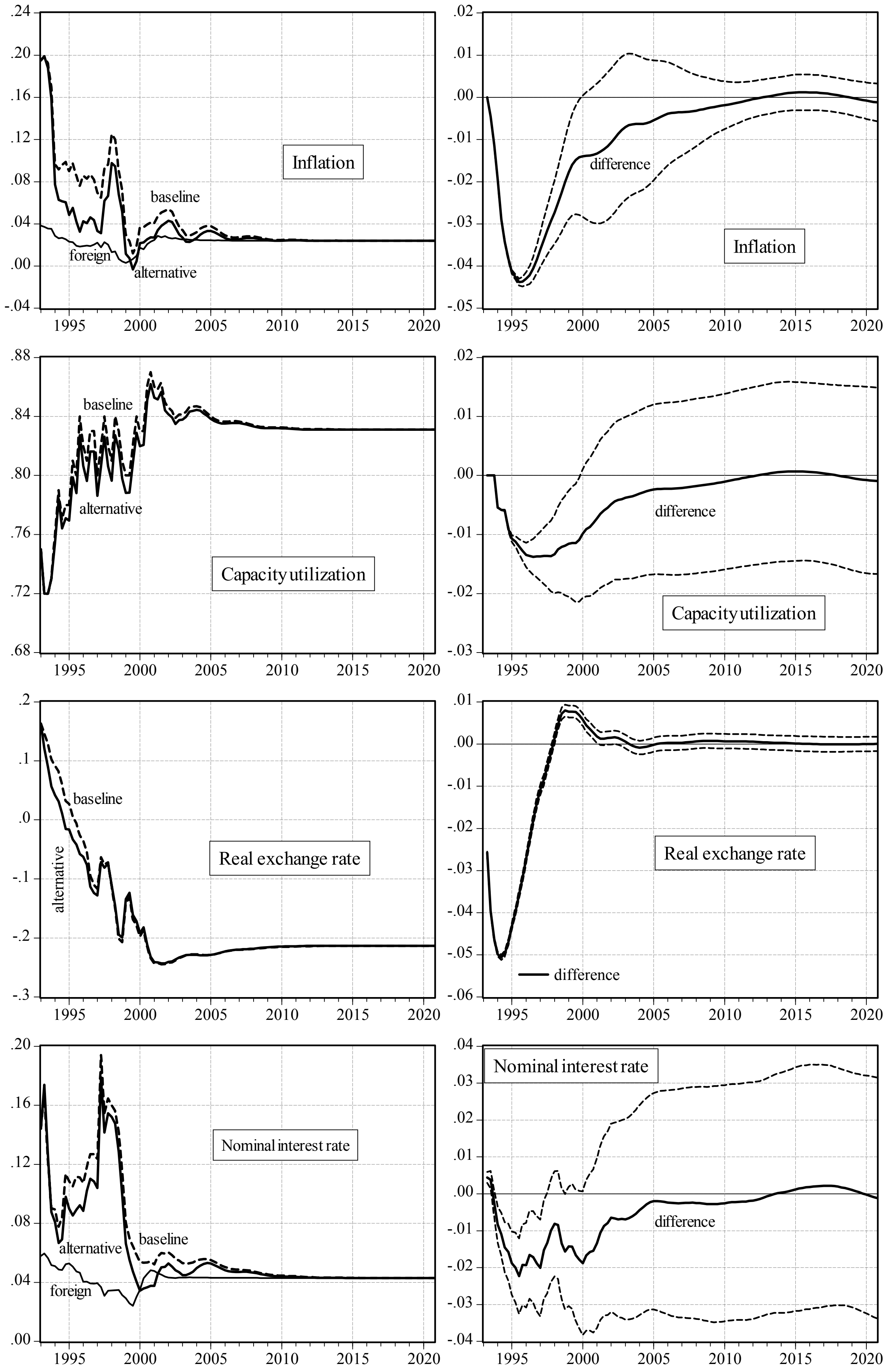


Fig.2.2 – Simulation of a more aggressive policy stance, Hungary

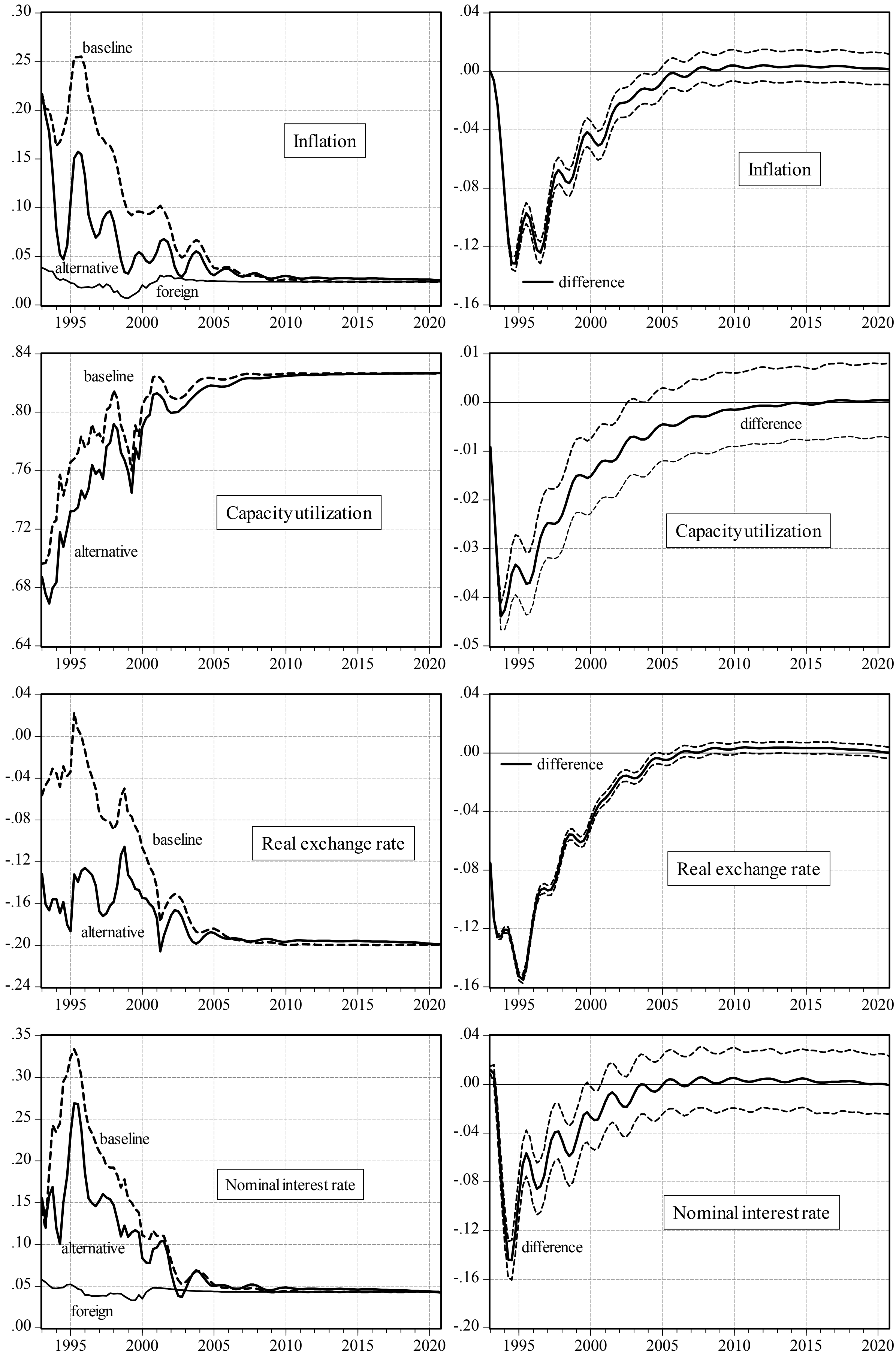
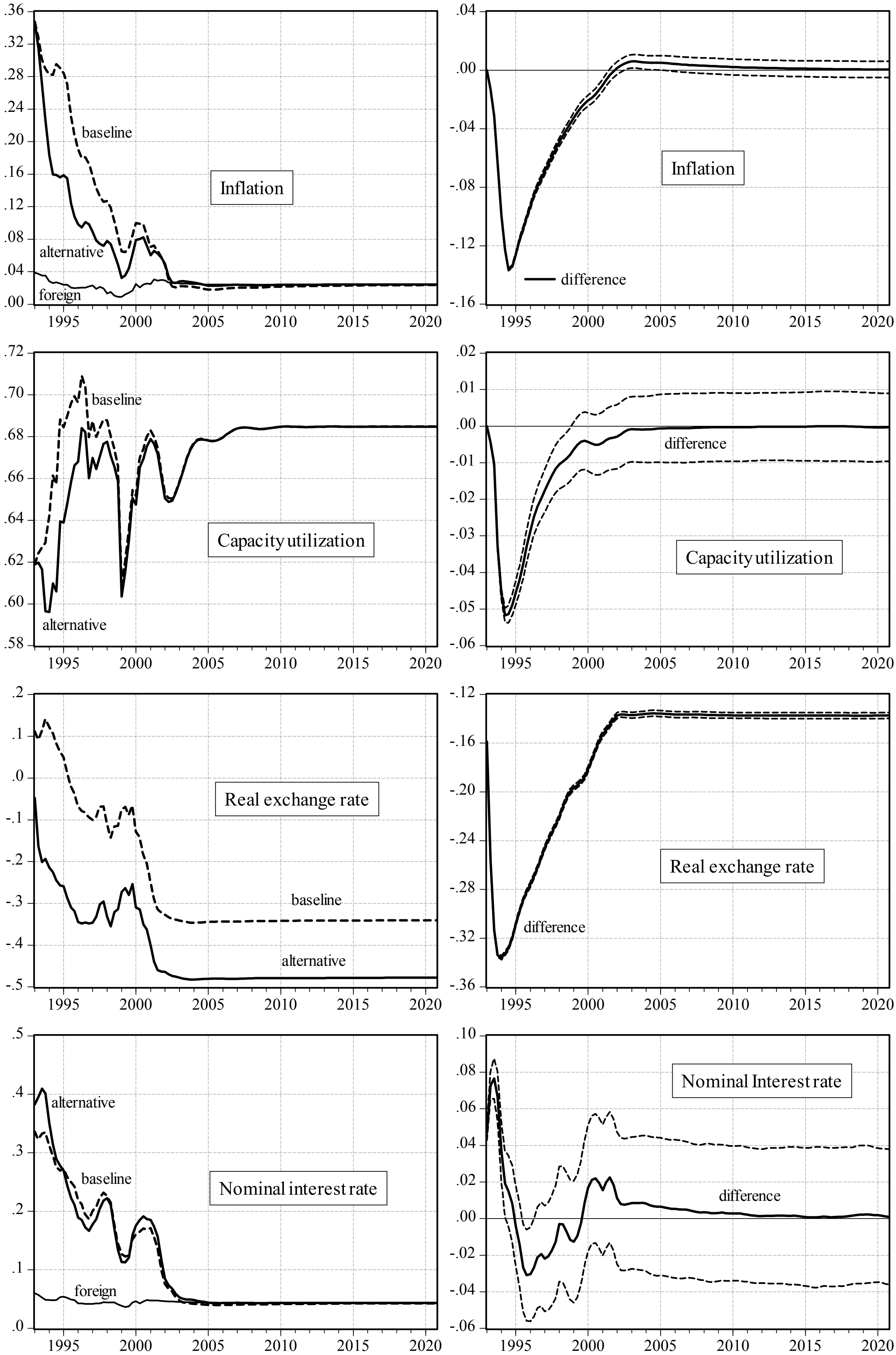


Fig. 2.3 – Simulation of a more aggressive policy stance, Poland



6. Conclusions

In this paper we have shown that a simple open macroeconomic model, with forward-looking inflation and exchange rate expectations, can adequately characterize the relationships between the output gap, inflation, the real interest rate and the exchange rate during the course of transition. These findings go against a widespread skepticism, which permeated a large part of previous research on these issues.

We have specified and estimated three small, similar models for the Czech Republic, Hungary and Poland, for the period 1991-2000. In our view of the transmission mechanism, domestic interest rates are at the center of the analysis, and their role is compatible in principle with different strategies concerning the exchange rate policy. Accordingly, we have characterized empirically the policy rules prevailing in each country, and assessed the relative importance of the interest rate channel (on aggregate demand) and the exchange rate channel (which affects both aggregate demand and supply) in determining the path of (dis)inflation. From this point of view, it is interesting to observe that the effect of the real exchange rate on the evolution of inflation is much more direct in the case of the Czech Republic, relative to the other countries. This corresponds well to the fact that the CNB consistently placed a greater emphasis, relative to the other two central banks, on the role of real exchange appreciations as a means to disinflation. Similarly, although the effects of the real interest rate and of changes in the real exchange rate on aggregate demand are in each case well identified and significant, we also found that in the Czech Republic real interest rates affect demand only with a very long lag, whereas the competitiveness effect takes place with a shorter lag and is much stronger than in the other countries: again, this corresponds well to the different exchange rate strategy adopted by the CNB.

Apart from these quantitative differences, the models are broadly similar from the qualitative point of view. This provides an important crosscheck on the validity of our modeling framework, and suggests that “orthodox”[40] relations between macroeconomic variables were already effective soon after the beginning of transition. The diagnostic tests on the estimated equations also suggest some evidence of instability, but this is limited to the first few years after the start of transition.

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[40] As defined in footnote 2.

This econometric evidence has two main implications for the current debate on monetary policy in transition. First, since we show that a well specified mechanism of policy transmission and of the inflation process is at work and can be kept under control by appropriately setting the level of domestic interest rates, this provides strong support for the *feasibility* of an inflation targeting (IT) strategy (and indeed for other structured strategies of monetary policy).

Second, since we show that this mechanism has been operating for quite some time and under different characterizations of the policy regime, we believe that this underlines the basic *continuity* between IT and other strategies of monetary policy. We support the idea that IT is best thought of as a policy framework, compatible with a *range* of policy strategies. In this context, IT is designed to serve a number of politico-economic functions, signaling the stance of the central bank vis-à-vis other actors on the macroeconomic scene and suggesting an appropriate nominal anchor for inflation expectations. This becomes valuable especially once the pre-announcement of a rate of crawl of the exchange rate can no longer play a central role in the strategy for disinflation, as the latter has been almost completed. From this point of view, IT emerges as the natural follower to a strategy of crawling band, in particular during a process of accession to the EU and EMU.

Finally, we have also used the estimated model to simulate the continuation of disinflation in the coming years, and to evaluate the effects of a more aggressive stance of monetary policy, if it had been adopted in the past. Unsurprisingly, we found that in the Czech Republic, which is the first country to complete disinflation in our sample, a more aggressive policy would have mostly induced negative output effects. On the other hand the counterfactual evidence is more dubious for the other countries. Overall, we do not place too much confidence in the results of counterfactual simulations, as we cannot be sure how “deep” our estimated parameters are, in the sense of Lucas (1976). However, we also believe that, if a more aggressive policy stance goes together with more credibility and hence less persistence, then plausibly its costs would be correspondingly reduced.

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APPENDIX 1: Data description and sources

Table A1 - Basic data

Label	Country	Period, from-to	Description	Source
P	Czech Republic	1991.1 -	Consumer Price Index All items, 1995 = 1	MEI
	Hungary	1989.1 -	Consumer Price Index All items, 1995 = 1	MEI
	Poland	1989.1 - 1994.4	Consumer prices (interpolator series)	IFS
		1995.1 -	Consumer Price Index All items, 1995 = 1	MEI
P_{GE}	Germany	1989.1 -	Consumer Price Index All items, 1995 = 1	MEI
P_{US}	United States	1989.1 -	Consumer Price Index All items, 1995 = 1	MEI
Y	Czech Republic	1991.2 -	Capacity Utilisation (Business tendency surveys), %	MEI
	Hungary	1991.1 -		MEI
	Poland	1989.1 -		MEI
\$	Czech Republic	1991.1 -	CK/\$ exchange rate, quarterly average	MEI
	Hungary	1989.1 - 1990.4	Forint official rate (interpolator series)	IFS
		1991.1 -	F/\$ exchange rate, quarterly average	MEI
	Poland	1989.1 - 1990.4	Zloty official rate (interpolator series)	IFS
		1991.1 -	Z/\$ exchange rate, quarterly average	MEI
D	Germany	1989.1 -	D-Mark/\$ exchange rate, quarterly average	MEI
R_{3m}	Czech Republic	1991.1 - 1992.4	Avg. interest rate, 12-mth dep., (interpolator series)	LAM
		1993.1 -	3-month PRIBOR, %	MEI
	Hungary	1989.1 - 1990.4	Treasury bill rate, (interpolator series)	IFS
		1991.1 -	90 day Treasury bill yield, %	MEI
	Poland	1989.1 - 1992.2	Treasury bill rate, (interpolator series)	IFS
		1992.3 -	3 month Treasury bill rate, %	MEI
WT	World	1989.1 -	World Trade in manufactures (BN 1995 \$)	MEI
BG	Czech Republic	1993.1 -	Public deficit (-) or surplus (BN CK)	IFS
YU	Czech Republic	1993.1 -	GDP at current prices(BN CK)	MEI
R_{GE}	Germany	1989.1 -	Call Money Rate, % p.a.	MEI
R_{US}	United States	1989.1 -	3-month Certificates of Deposit, % p.a.	MEI

MEI = OECD, Main Economic Indicators; IFS = IMF, International Financial Statistics; LAM = Univ. of Gdansk data.
Significant seasonal patterns have been removed using the Tramo/Seats procedure, see Gomez and Maravall (1997).
To avoid shortening the sample, some missing data have been interpolated using an “interpolator” series.

Table A2 - The weights of the foreign variables

country	periods	$w_{US,t}$	$w_{GE,t}$
Czech Republic	before 1993, May	0.49	0.51
	from 1993, May to 1997, May	0.35	0.65
	after 1997, May	0.00	1.00
Hungary	before 1994, May	0.50	0.50
	from 1994, June to 1999 December	0.30	0.70
	after 1999, December	0.00	1.00
Poland	before 1991, May	1.00	0.00
	from 1991, June to 2000, March	0.45	0.55
	after 2000, March	0.00	1.00

Table A3 - Model variables

Label	Definition	Description
$\boldsymbol{p_t}$	$\log(P_t)$	***domestic price logs** (1995 = 0)*
$\boldsymbol{\Delta p_t}$	$4\ (p_t - p_{t-1})$	***domestic inflation rate***
$Pdot_{GE,t}$	$P_{GE,t}/P_{GE,t-1} - 1$	Germany inflation rate
$Pdot_{US,t}$	$P_{US,t}/P_{US,t-1} - 1$	US inflation rate
$\boldsymbol{p^*_t}$	$\log\{\Pi_{j=1,t}\ [1 + (w_{GE,j}\ Pdot_{GE,j} + w_{US,j}\ Pdot_{US,j})]\}$	***foreign price logs** (1995 = 0)*
$\boldsymbol{\Delta p^*_t}$	$4\ (p^*_t - p^*_{t-1})$	***foreign inflation rate***
$\$dot_t$	$\$_t/\$_{t-1} - 1$	Dollar exchange rate % change
$Ddot_t$	$(\$_t/D_t)/(\$_{t-1}/D_{t-1}) - 1$	D-Mark exchange rate % change
$\boldsymbol{e_t}$	$\log\{\Pi_{j=1,t}\ [1 + (w_{GE,j}\ Ddot_j + w_{US,j}\ \$dot_j)]\}$	***log nominal exchange rate,**1995=0*
	$e_t + p^*_t - p_t$	***log real exchange rate,**1995 = 0*
$\boldsymbol{y_t}$	$Y/100$	***capacity utilization ratio***
$\boldsymbol{i_t}$	$R_{3m}/100$	***domestic interest rate***
$\boldsymbol{\Delta w_t}$	$4\ \log(WT_t/WT_{t-1})$	***world demand***
$\boldsymbol{bg_t}$	BG_t/YU_t	***government balance on GDP***
$\boldsymbol{i^*_t}$	$w_{GE,t}\ R_{GE,t} + w_{US,t}\ R_{US,t}$	***foreign interest rate***

Variables used in the models are in bold and italic. Variable definitions are the same for all countries
The data set used for estimation is available at: http://www.spbo.unibo.it/pais/golinelli/macro.html

APPENDIX 2: Unit roots and cointegration properties

Unit root analysis. Table A4 reports Dickey and Fuller (1979) unit root test statistics (column *ADF*). The order of the test (column *k*) was selected by following the general (max $k = 5$) to specific ("testing down") procedure advocated by Ng and Perron (1995). Testing models are set as indicated in the column *model*. Phillips and Perron (1988) test results (with West truncation lag = 3) are reported in the column *PP*.

The sample period is 1993.1-2001.1 for the Czech Republic, 1991.1 2001.1 for Hungary and Poland. Overall, the variables examined are at most I(1).

Table A4 also reports tests for simple combinations between I(1) variables: if the unit root null hypothesis is rejected, the corresponding I(1) variables are cointegrated. As real ex post short-term interest rates are stationary, both foreign and domestic nominal interest rates are (1, -1) cointegrated with inflation. On the contrary, the real exchange rate is I(1) in all countries.

Table A4 – Unit root tests

		Czech Republic			Hungary			Poland		
(*a*)	*mod*	*ADF* (*b*)	*k*	*PP* (*b*)	*ADF* (*b*)	*k*	*PP* (*b*)	*ADF* (*b*)	*k*	*PP* (*b*)
p_t	c, t	-2.603	0	-2.572	0.592	4	-0.088	-1.909	5	-4.455^{**}
Δp_t	c	-3.876^{**}	0	-3.746^{**}	-1.348	4	-1.890	-1.624	4	-2.063
p^*_t	c, t	-3.234	3	-3.160	-2.218	3	-2.902	-2.109	3	-3.016
Δp^*_t	c	-1.921	1	-3.409^{*}	-1.922	2	-3.828^{**}	-2.934^{*}	2	-4.599^{**}
$\Delta p_t - \Delta p^*_t$	c	-3.987^{**}	0	-3.869^{**}	-1.587	4	-1.744	-5.863^{**}	3	-1.874
e_t	c, t	-3.141	1	-2.209	-0.492	3	-0.395	0.573	0	0.500
Δe_t	c	-4.937^{**}	2	-4.726^{**}	-5.505^{**}	0	-5.685^{**}	-2.027	1	-3.341^{*}
$e_t + p^*_t - p_t$	c, t	-2.934	3	-3.422	-1.830	3	-2.260	-2.635	4	-4.876^{**}
$\Delta(e_t + p^*_t - p_t)$	c	-4.367^{**}	2	-4.459^{**}	-4.700^{**}	2	-8.013^{**}	-6.204^{**}	3	-4.542^{**}
y_t	c	-1.460	0	-1.189	-0.908	4	-0.607	-2.274	2	-1.835
Δy_t	c	-5.709^{**}	1	-6.768^{**}	-2.691	3	-7.277^{**}	-5.130^{**}	0	-5.177^{**}
Δw_t	c	-4.499^{**}	4	-3.136^{*}	-4.155^{*}	3	-3.503^{*}	-4.155^{**}	3	-3.503^{*}
bg_t	c	-4.800^{**}	0	-4.826^{**}	-	-	-	-	-	-
i_t	c	-1.272	0	-1.381	-2.769	3	-1.482	-2.287	5	-1.359
Δi_t	c	-5.850^{**}	0	-5.850^{**}	-3.462^{*}	3	-3.964^{**}	-4.933^{**}	4	-4.458^{**}
$i_t - \Delta p_{t+1}$	c	-7.476^{**}	0	-6.732^{**}	-5.618^{**}	3	-3.174^{*}	-3.629^{**}	3	-6.592^{**}
i^*_t	c	-2.045	1	-2.001	-2.136	5	-3.003^{*}	-1.495	4	-2.700
Δi^*_t	c	-3.860^{**}	0	-3.923^{**}	-3.163^{*}	4	-5.807^{**}	-5.482^{**}	0	-5.446^{**}
$i^*_t - \Delta p^*_{t+1}$	c	-4.289^{**}	0	-4.453^{**}	-5.036^{**}	0	-5.030^{**}	-5.527^{**}	0	-5.547^{**}

([a]) Data Labels are in Table A3.
([b]) MacKinnon critical values for 1% and 5% rejection of the unit root null:
-3.60 and -2.93, for models with constant and without trend (*c*);
-4.20 and -3.52, for models with constant and trend: (*c, t*);
* and ** respectively indicate 5% and 1% rejection of the unit root null.

Cointegration analysis. In order to reach a long-run-balanced data representation for each equation, integrated variables must be cointegrated. Since cointegration is invariant if we add redundant variables to an irreducible cointegrating relation (see Davidson, 1998), we have applied the Johansen (1995) cointegration approach to subsets of variables to save on the number of parameters to be estimated (a very similar approach is followed by Juselius, 1992). In the last two equations of the structural models (the UIP and the interest rate rule), the balanced representation of the relationships among the exchange rate, interest rates and inflation differentials requires that the real exchange rate is at most I(1), and that the real interest rates are stationary. Results in Table A4 satisfy these conditions as a whole, while the cointegration results in the aggregate supply and demand equations will be discussed separately for each country.

For the **Czech Republic**, the AS equation implies a balanced representation between domestic and foreign inflation rates, and between the real interest rate and the rate of capacity utilization. Table A4 suggests that the inflation differential $\Delta p_t - \Delta p^*_t$ is stationary. In addition, the Johansen trace test in Table A5.1 detects rank = 2 in a VAR(4) model for the four at most I(1) variables: (Δp_t ; $e_t+p^*_t-p_t$; y_t ; Δp^*_t). The identification restrictions on the two long run relationships are not rejected.
For the AD equation, Table A5.1 reports cointegration results between the only two I(1) variables: (y_t ; i_t) in a VAR(5) model conditional to past values of world demand, inflation rate, public deficit on GDP, and real exchange rate. Given the relevant number of lags necessary to obtain white noise residuals, the long run relationship is not efficiently estimated. The weak exogeneity of the nominal interest rate is largely not rejected; this supports the irrelevance of capacity utilization in explaining the cyclical path of the nominal interest rate in the interest rate rule equation.

Table A5.1 – Czech Republic. Cointegration between the I(1) variables

AS equation

Cointegration rank test between (Δp_t ; $e_t+p^*_t-p_t$; y_t ; Δp^*_t)

null	*eigenvalue*	*Trace*	*5% c.v.*	*1% c.v.*
None *	0.776517	92.66400	53.12	60.16
At most 1 *	0.589486	43.21614	34.91	41.07
At most 2	0.273146	13.83477	19.96	24.60
At most 3	0.095349	3.306804	9.24	12.97

* denotes rejection of the hypothesis at the 1% level
Trace test indicates 2 cointegrating equations at 1% level

Long run relationships (s.e. in brackets):

$$\Delta p = \underset{(0.018)}{0.067} + \Delta p^* \qquad e + p^* - p = \underset{(0.313)}{1.0154} - \underset{(0.384)}{1.326}\, y$$

LR test for binding restrictions (rank = 2):

Chi-square(3) 1.979
P-value 0.577

AD equation

Cointegration rank test between (y_t ; i_t) variables:

null	*eigenvalue*	*Trace*	*5% c.v.*	*1% c.v.*
None *	0.557283	27.20282	19.96	24.60
At most 1	0.009457	0.313581	9.24	12.97

* denotes rejection of the hypothesis at the 1% level
Trace test indicates 1 cointegrating equation at 1% level

Long run relationships (s.e. in brackets):

$$y = \underset{(0.058)}{0.827} - \underset{(0.278)}{0.266}\, i$$

LR test for i_t weak exogeneity restriction (rank = 1):

Chi-square(1) 0.065
P-value 0.799

Hungary. In the AS equation, we found a cointegration relationship between the I(1) triplet (Δp_t ; Δp^*_t ; $e_t+p^*_t-p_t$). Results in Table A5.2 suggest that the long run inflation differential is explained by the level of the real exchange rate (the corresponding restrictions are not rejected).

In the AD equation, we detect rank = 2 in a VAR(5) model for (y_t ; i_t ; Δp_t), conditional to past values of world demand growth and past variations of the real exchange rate. The results, reported in Table A5.2, can also be interpreted (and identified) as a long run demand curve, where the level of the capacity utilization depends on the level of the real interest rates. As in the Czech case, the weak exogeneity of the interest rate is not rejected.

Table A5.2 – Hungary. Cointegration between the I(1) variables

AS equation

Cointegration rank test between (Δp_t ; Δp^*_t ; $e_t+p^*_t-p_t$) variables:

null	*eigenvalue*	*Trace*	*5% c.v.*	*1% c.v.*
None *	0.392819	37.53525	34.91	41.07
At most 1	0.259873	17.57812	19.96	24.60
At most 2	0.129353	5.540764	9.24	12.97

* denotes rejection of the hypothesis at the 5% level
Trace test indicates 1 cointegrating equation at 5% level

Long run relationships (s.e. in brackets):

$$\Delta p - \Delta p^* = \underset{(0.024)}{0.125} + \underset{(0.303)}{1.385}\ (e+p^*-p)$$

LR test for binding restrictions (rank = 1):

Chi-square(1)	0.034
P-value	0.854

AD equation

Cointegration rank test between (y_t ; i_t ; Δp_t) variables:

null	*eigenvalue*	*Trace*	*5% c.v.*	*1% c.v.*
None **	0.532740	53.40690	34.91	41.07
At most 1 *	0.355440	22.21129	19.96	24.60
At most 2	0.097469	4.204626	9.24	12.97

** (*) denotes rejections of the hypothesis at the 1% (5%) levels
Trace test indicates 1 cointegrating equation at 1% level

Long run relationships (s.e. in brackets):

$$y = \underset{(0.027)}{0.682} - \underset{(0.511)}{0.769}\ (i - \Delta p)$$

LR test for binding restriction (rank = 1):

Chi-square(1)	3.075
P-value	0.380

Poland. The specification of the AS equation implies a static long run solution, where the inflation rate and changes in the nominal exchange rate, both I(1), are cointegrated. The stationarity of the real exchange rate in differences (reported in Table A4) supports the cointegration of the supply equation. This result is confirmed by the cointegration experiments in Table A5.3, where a relationship in differences between the inflation differential and the nominal exchange is detected, and the corresponding identification restriction is not rejected.[41] As expected, in the demand equation there are two cointegrated relationships between (y_t ; i_t ; Δp_t), all I(1). In addition, as shown in Table A5.3, the long run relationship between capacity utilization and the real ex post interest rate (the latter being weakly exogenous for the estimation of the long run demand equation) is identified.

Table A5.3 – Poland. Cointegration between the I(1) variables

AS equation

Cointegration rank test between (Δp_t ; Δe_t ; Δp^*_t) variables:

null	*eigenvalue*	*Trace*	*5% c.v.*	*1% c.v.*
None *	0.428637	36.09319	34.91	41.07
At most 1	0.185024	13.14424	19.96	24.60
At most 2	0.109520	4.755763	9.24	12.97

* denotes rejection of the hypothesis at the 5% level
Trace test indicates 1 cointegrating equation at 5% level

Long run relationships (s.e. in brackets):

$$\Delta p - \Delta p^* = \underset{(0.047)}{0.023} + \underset{(0.278)}{0.622}\ \Delta e$$

LR test for binding restrictions (rank = 1):

Chi-square(1)	0.329
P-value	0.566

AD equation

Cointegration rank test between (y_t ; i_t -Δp_t) variables:

null	*eigenvalue*	*Trace*	*5% c.v.*	*1% c.v.*
None *	0.276075	22.57671	19.96	24.60
At most 1	0.193033	9.007850	9.24	12.97

* denotes rejection of the hypothesis at the 5% level
Trace test indicates 1 cointegrating equation at 5% levels

Long run relationships (s.e. in brackets):

$$y = \underset{(0.027)}{0.729} - \underset{(0.315)}{0.851}\ (i - \Delta p)$$

LR test for (i_t -$\Delta p_{t)}$ weak exogeneity restriction (rank = 1):

Chi-square(1)	0.518
P-value	0.472

[41] Note also that (as implied by the first-order integrated real exchange rate in Table A4) the long run elasticity of the inflation differential to the nominal exchange rate is not significantly different from unity.

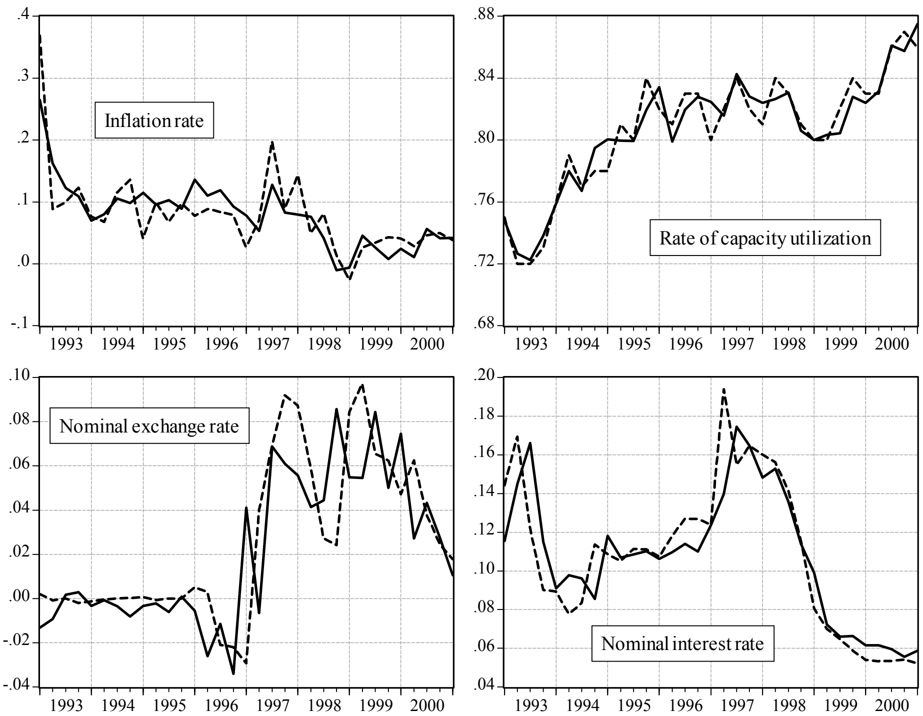
APPENDIX 3: Diagnostic tests and model performance

Tab. A6.1 – Diagnostic tests, Czech Republic

	equation 1	equation 2	equation 3	equation 4
S.E. of regression	0.0408	0.0122	0.0253	0.0181
Durbin-Watson statistic	2.055	2.163	2.458	1.913
Autocorrelation, LM(1) *([a])*	0.4395	0.5217	0.1340	0.9973
Autocorrelation, LM(4) *([a])*	0.8661	0.3929	0.0300	0.8629
Heteroskedasticity *([b])*	0.0115	0.2024	0.0266	0.0343
ARCH(1) *([c])*	0.2615	0.9506	0.4282	0.4917
ARCH(4) *([c])*	0.1835	0.8056	0.9363	0.9977
Normality *([d])*	0.6562	0.6338	0.0037	0.0422
Parameter constancy, 1993 *([e])*	0.0061	0.7548	0.9777	0.0080
Parameter constancy, 2000 *([f])*	0.9920	0.6872	0.7635	0.9892

([a]) *P*-values of the F version of Lagrange multiplier tests of residual serial correlation
([b]) *P*-values of the F version, based on the regression of squared residuals on squared fitted values
([c]) *P*-values of the F version of autoregressive conditional heteroskedasticity tests, order in brackets
([d]) P-values of the chi-square statistic based on a test of skewness and kurtosis of residuals
([e]) *P*-values of the F version of the second Chow test from 1993.1 to 1993.4; system *P*_value = 0.0243
([f]) *P*-values of the F version of the second Chow test from 2000.1 to 2001.1; system *P*_value = 0.9845

Fig. A1.1 – Actual (dotted lines) and fitted values, Czech R.

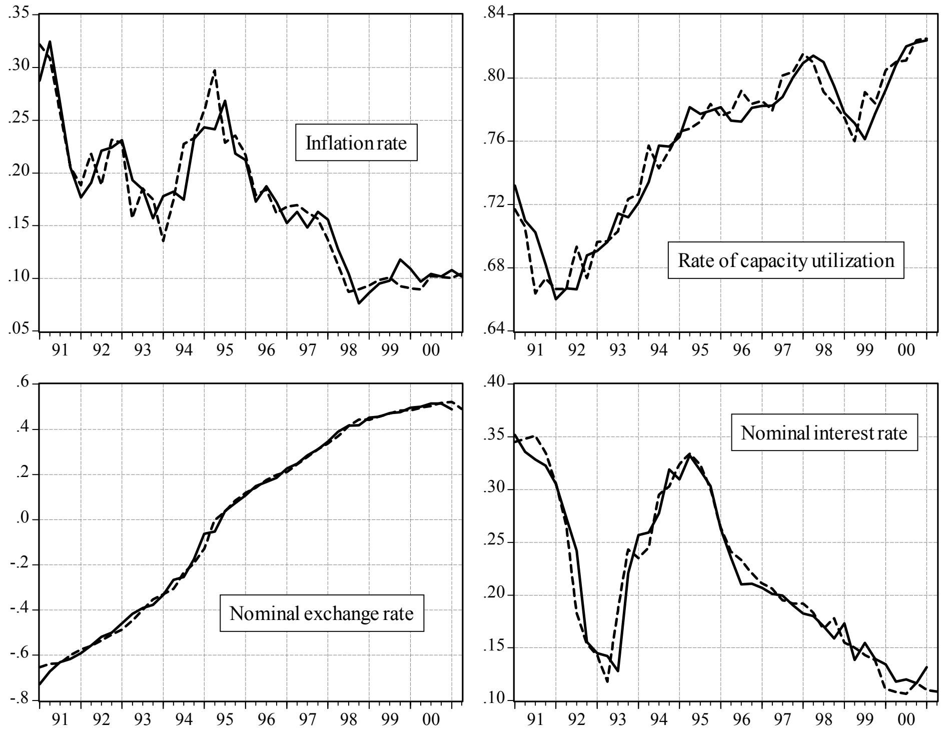


Tab. A6.2 – Diagnostic tests, Hungary

	equation 1	equation 2	equation 3	equation 4
S.E. of regression	0.0224	0.0137	0.0241	0.0195
Durbin-Watson statistic	2.245	1.924	1.673	1.943
Autocorrelation, LM(1) (a)	0.3260	0.8967	0.9030	0.9530
Autocorrelation, LM(4) (a)	0.1852	0.0752	0.4848	0.7301
Heteroskedasticity (b)	0.1815	0.1000	0.7265	0.4693
ARCH(1) (c)	0.6514	0.5128	0.0551	0.8961
ARCH(4) (c)	0.4008	0.0410	0.1364	0.0323
Normality (d)	0.3494	0.3003	0.0130	0.0009
Parameter constancy, 1991 (e)	0.3696	0.0024	0.0147	0.5503
Parameter constancy, 1991-92 (e)	0.3859	0.0002	0.1003	0.0370
Parameter constancy, 1999-00 (e)	0.9101	0.6205	0.9875	0.7247
Parameter constancy, 2000 (e)	0.9431	0.9374	0.8462	0.4980

(a) *P*-values of the F version of Lagrange multiplier tests of residual serial correlation
(b) *P*-values of the F version, based on the regression of squared residuals on squared fitted values
(c) *P*-values of the F version of autoregressive conditional heteroskedasticity tests, order in brackets
(d) P-values of the chi-square statistic based on a test of skewness and kurtosis of residuals
(e) *P*-values of the F version of the second Chow test for all the quarters in the year(s); system *P*_values: 0.0074 (1991), 0.0039 (1991-1992)
(f) *P*-values of the F version of the second Chow test from 1999.1 (2000.1) to 2001.1; system *P*_values: 0.9581 (1999-00), 0.9646 (2000)

Fig. A1.2 – Actual (dotted lines) and fitted values, Hungary



Tab. A6.3 – Diagnostic tests, Poland

	equation 1	equation 2	equation 3	equation 4
S.E. of regression	0.0207	0.0117	0.0309	0.0143
Durbin-Watson statistic	1.695	2.393	2.391	2.013
Autocorrelation, LM(1) *([a])*	0.8589	0.1983	0.0780	0.9011
Autocorrelation, LM(4) *([a])*	0.6318	0.5742	0.2025	0.7959
Heteroskedasticity *([b])*	0.8066	0.8989	0.3630	0.0343
ARCH(1) *([c])*	0.3060	0.6531	0.2312	0.2592
ARCH(4) *([c])*	0.6691	0.9527	0.3678	0.4122
Normality *([d])*	0.0673	0.0000	0.8527	0.9754
Parameter constancy, 1991 *([e])*	0.4642	0.5642	0.1834	0.0757
Parameter constancy, 1991-92 *([e])*	0.0108	0.7566	0.2801	0.2006
Parameter constancy, 1999-00 *(e)*	0.7383	0.0030	0.0470	0.9728
Parameter constancy, 2000 *(e)*	0.9597	0.6085	0.1696	0.9142

([a]) *P*-values of the F version of Lagrange multiplier tests of residual serial correlation
([b]) *P*-values of the F version, based on the regression of squared residuals on squared fitted values
([c]) *P*-values of the F version of autoregressive conditional heteroskedasticity tests, order in brackets
([d]) P-values of the chi-square statistic based on a test of skewness and kurtosis of residuals
([e]) *P*-values of the F version of the second Chow test for all the quarters in the year(s); system *P*_values: 0.1848 (1991), 0.0756 (1991-1992)
([f]) *P*-values of the F version of the second Chow test from 1999.1 (2000.1) to 2001.1; system *P*_values: 0.0652 (1999-00), 0.7976 (2000)

Fig. A1.2 – Actual (dotted lines) and fitted values, Poland