Exchange rate pass-through and real exchange rate in EU candidate countries

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Abstract

This paper studies a particular aspect of the choice of exchange rate regime by EU candidate countries in the run-up to membership of European Economic and Monetary Union (EMU). The fact that these countries have adopted various exchange rate systems reflects a divergence of opinion on the appropriate way to handle macroeconomics and, in particular, to curb inflation. This paper studies the connection between changes in the exchange rate and inflation as perceived in various exchange rate systems in order to draw conclusions with regard to the possible role of exchange rate management in achieving and maintaining low inflation in these countries. To this end we model price and exchange rate changes simultaneously and incorporate adjustment towards the equilibrium real exchange rate. We use a state-space model and the Kalman filter to infer unobserved variables and time-varying parameters.

Zusammenfassung

Dieses Papier untersucht einen besonderen Aspekt bei der Wahl eines Wechselkursregimes von EU Beitrittskandidaten auf dem Weg zu einer Mitgliedschaft in der Europäischen Wirtschafts- und Währungsunion. Die Tatsache, dass diese Länder verschiedene Wechselkursregime gewählt haben, deutet auf divergierende Ansichten darüber hin, wie die Wirtschaftspolitik geführt werden soll und insbesondere, wie die Inflation gesenkt werden soll. In diesem Aufsatz werden die Beziehungen zwischen Änderungen in den Wechselkursen und den Preisen in verschiedenen Wechselkursregimen untersucht. Dabei soll herausgefunden werden, welche Rolle Wechselkurse spielen, wenn es darum geht, niedrige Preissteigerungsraten zu realisieren und diese dann auch beizubehalten. Preis- und Wechselkursänderungen werden simultan modelliert unter gleichzeitiger Berücksichtigung der Anpassung an einen Gleichgewichtwechselkurs. Raum-Zustands-Modelle und Kalmanfilter werden verwendet, um auf unbeobachtete Variablen und zeitvariable Parameter schließen zu können.

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Exchange rate pass-through and real exchange rate in EU candidate countries*

1 Introduction

The existence of a wide range of different exchange rate regimes is an important characteristic of EU candidate countries: virtually every possible type can be found, including the two extreme systems, currency boards and freely floating exchange rates. All of these countries place strong emphasis on the objective of low inflation – both because of its inherent benefits and because of their aim to fulfil the Maastricht criteria so that they can participate in European Economic and Monetary Union (EMU). The common goal of low inflation and the diversity of exchange rate regimes is a clear indication that policymakers in these countries lack a consensus view with regard to the best monetary regime to be adopted in the pre-EMU period.

The issues associated with the choice of an exchange rate regime can be divided into two main categories: (1) "fundamental issues", such as the transmission of monetary policy, the role of asymmetric shocks, labour mobility, and wage and price flexibility; and (2) "sustainability issues", such as the vulnerability of rigid regimes to speculative attacks and the possible role of different regimes in reinforcing the destabilising effect of capital flows. With regard to both issues, the literature clearly considers more flexible regimes to be superior (e.g. Mishkin, 1998; Stockman, 1999). However, there are also papers that challenge this view (e.g. McKinnon–Pill, 1999; Darvas–Szapáry, 2000) and suggest that the choice of a regime should be based on fundamental issues. A key issue is the role of the exchange rate in controlling inflation, either by direct exchange rate targeting or by the indirect influence of other (e.g. inflation targeting) regimes.

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Hundreds of theoretical and empirical papers analysing the pass-through issue have been published in developed countries but far fewer in developing countries. Given the wealth of theoretical research and the shortage of empirical research for EU candidate countries, this paper treats the possible role of using the exchange rate to control the inflation process in candidate countries as an empirical issue. However, application of the methods developed in the literature is rendered more complex by data availability and by some particular characteristics of these countries that set them apart from current EMU participants. Two of these characteristics also have crucial implications for the pass-through issue, as can be seen in Table 1, which shows some basic indicators for the countries to be analysed[1] in comparison with three EU member states, Greece, Portugal and Spain.[2] First, their income per capita is far lower than the EU average and most of them are growing faster than the EU average, suggesting that they are on a catch-up growth path. Second, their price levels, especially the price levels of non-tradables, are also much lower than the EU average, although these levels will probably increase gradually as these countries make progress in catching up. Consequently, the real exchange rates of these countries will probably continue to appreciate. If EU and EMU membership were considered imminent, this might enhance expectations of a real exchange rate appreciation.

These points have important implications for the pass-through issue: domestic prices are likely to change without there being any movement in exchange rates. Should an exchange rate movement be evident, the price changes should be carefully decomposed into pass-through and the underlying price convergence process.

The paper analyses four EU candidate countries (Czech Republic, Hungary, Poland, Slovenia) in a simultaneous time-varying error correction framework. The sample consists of quarterly data in the period from 1993 to 2000. It might be argued that this is a short sample, especially for time-varying parameter estimation. However, this seems to be the only solution if econometrics are to be performed for this issue. First, we will give a number of reasons why we would expect time variation of coefficients even in industrial countries, not to mention candidate countries. Second, the empirical section will show that fixed parameter pass-through estimates are invalidated by significant Chow breakpoint tests. Third, techniques based on vector autoregression (VAR) are not acceptable options. Even in a fixed parameter set-up they require the estimation of many parameters, rendering the degrees of freedom very

[1] There are 12 countries currently involved in accession negotiations with the European Union (EU). The "Luxembourg group" (Cyprus, Czech Republic, Estonia, Hungary, Poland and Slovenia) started accession negotiations in March 1998, while the "Helsinki group" (Bulgaria, Latvia, Lithuania, Malta, Romania and the Slovak Republic) started talks in February 2000. Turkey has been granted the status of an accession candidate, but the starting date for accession negotiations has not yet been settled.

[2] Greece joined the EU in 1981, Portugal and Spain in 1986.

low. Sensitivity of a VAR-based inference to the sample period, the lag length and the correlation between unadjusted innovations is demonstrated in the empirical section.

The rest of the paper is organised as follows. Section 2 briefly reviews the literature on pass-through and highlights those aspects of previous research that might be relevant for the empirical study of candidate countries. The model is presented in Section 3. Section 4 briefly describes exchange rate regimes and inflation developments. Section 5 presents the empirical results. Section 6 concludes. The annex gives the reasoning for our choice of the equilibrium real exchange rate model and presents some background calculations.

Table 1. Population, development, growth, openness, foreign direct investment

	(A) Population (millions) 1998	(B) GDP price level (% EU average) 1999	(C) GDP/capita in PPS♥ (% EU average) 1999	(D) Average GDP growth (%) 1997-99♣	(E) Imports• / GDP (%) 1999	(F) FDI inflows (% of GDP) average of 1994-98
Bulgaria	8.3	30	22	-0.4	46.3*	2.4
Croatia	4.5**			3.0	48.1	
Cyprus	0.8**	75	81	4.0	49.2	0.7
Czech Republic	10.3	38	59	-1.1	65.2	3.4
Estonia	1.5**	43	36	4.7	89.4*	7.1
Hungary	10.1	42	51	4.7	56.7	5.3
Latvia	2.5	41	27	4.2	57.6	6.4
Lithuania	3.7	44	29	2.8	50.1	3.2
Poland	38.7	48	37	5.2	30.0**	3.1
Romania	22.5	25	27	-4.9	34.3	2.3
Slovak Republic	5.4	34	47	4.3	74.8*	1.8
Slovenia	2.0	63	71	4.4	58.1*	1.1
Turkey	63.5	46	28	1.9	30.4**	0.5
Greece	10.5**	79	67	3.5	25.3	
Portugal	9.8**	65	76	3.3	40.0*	
Spain	39.3**	82	82	3.8	28.2*	

Source: Columns B, C, D, F: Eurostat, *Statistics in Focus: The GDP of the candidate countries* 27/2000, columns A and E: International Monetary Fund, *International Financial Statistics*, except columns D and E for Croatia: National Bank of Croatia, column E for Hungary: National Bank of Hungary.

♥: Purchasing power standard.

♣: The average GDP growth of the EU-15 countries was 2.5 in 1997-99.

•: Imports of goods and services according to national accounts statistics.

* 1998

** 1997

2 Exchange rate pass-through

2.1 A brief review of the literature

Hundreds of papers which analyse different aspects of exchange rates and product prices have been published.[3] The expression "exchange rate pass-through" is generally used to refer to the effects of exchange rate changes on one of the following: (1) import and export prices, (2) consumer prices, (3) investments and (4) trade volumes. Of these four topics, the primary focus is on the effects of exchange rate changes on import and export prices because, on the one hand, this is the natural ground for studying the pricing practices of firms and, on the other, a response by import-export prices to exchange rate change is usually needed before there is any consequence for consumer prices, investment and trade volumes.

Several studies are surveyed in Menon (1995) and Glodberg–Knetter (1997). Glodberg–Knetter report a fairly well-supported consensus for a 60%, hence incomplete, pass-through into US import prices, while also indicating important differences depending on the characteristics of the product under analysis. Menon also emphasises the general result of incompleteness, and stresses the differences among countries and even differences between studies for a given country. There is no agreement as to whether the pass-through relationship remained stable in past decades nor as to whether the pass-through symmetrically follows the direction of the exchange rate change or not. Menon attributes these divergences to heterogeneity of methodologies, model specification and variable construction. Of the literature that we have examined, the papers published since these two surveys clarify many details but still fail to reach a consensus other than with regard to the incompleteness of pass-through in most cases.[4]

A vast number of theoretical models have been established to explain the incomplete nature of exchange rate pass-through. As far back as the 1970s Magee (1975) emphasised that "there is no single coherent theory of devaluation (or revaluation), but rather an amalgam of reasons why prices may not respond fully to exchange rate changes".[5]

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[3] For example, Goldberg–Knetter (1997) reported in a survey paper that EconLit gives approximately 700 entries published dated from the 1970s for the key terms "law of one price", "purchasing power parity", "exchange rate pass-through" and "pricing to market".

[4] A recent example of disagreement with regard to the evaluation of the extent of pass-through is found in Gordon (1999). He turns conventional wisdom completely on its head: by analysing all "leavers" and "stayers" in the aftermath of the 1992 ERM crisis, he shows that most of the devaluation (combined with heavy fiscal tightening) in the "leaver" countries passed into higher inflation and only a minor part into real GDP growth.

[5] The quotation is taken from Menon (1995).

Recent theoretical literature builds almost exclusively on the concept of market segmentation, although incomplete pass-through is not necessarily evidence of a lack of market integration. In fact, even if there were perfect competition and product homogeneity, the pass-through may be different from one due to non-zero price elasticity of demand and the supply side effects of exchange rate changes (Menon, 1995; Hens, 1997). Nonetheless, the literature is set (with rare exceptions) in an oligopolistic framework with imperfect competition and third-degree price discrimination.

There are some obvious reasons why national (or regional) markets might be segmented, with scope, therefore, for price discrimination: (i) transportation costs, (ii) customs duties, (iii) non-tariff barriers, (iv) physical differences in product characteristics (e.g. 110 volt versus 220 volt electrical appliances) and (v) home or brand loyalty of consumers. Given segmented markets, the seminal papers of Krugman (1987) and Dornbusch (1987) initiated hundreds of models studying the variations in mark-ups in response to exchange rate changes in oligopolistic settings.

Another line taken in the literature has its roots in Baldwin (1988), Dixit (1989) and Baldwin–Krugman (1989); it emphasises hysteretic effects arising from the sunk costs of entering a market that firms cannot recoup when they leave the market. Baldwin and Baldwin–Krugman show, in an oligopolistic setting, that "large" exchange rate movements can alter a country's market structure and that a change in market structure permanently shifts import prices and trade volumes. Dixit demonstrates in a perfectly competitive setting with price-taking firms that there is almost perfect pass-through into domestic prices when foreign firms enter or exit the market and near zero pass-through otherwise.

A third selection from the literature concentrates on institutional settings such as the effects of non-tariff barriers or the role of multinational corporations and intra-firm trade, as surveyed and emphasised in Menon (1995, 1996). Multinational corporations play a very significant role in candidate countries due to privatisation and foreign direct investment inflows (Table 1).

In a recent paper, Hens et al. (1999) follow the tradition of oligopolistic competition without sunk costs and eliminate several restrictive assumptions made in the earlier literature. They underscore the fact that the direction and magnitude of pass-through depends on the combination of both demand and costs functions; specifically, it depends on the extent of economies of scope and the strategic impact of competitors' sales on their profits. Their theoretical model shows that prices may move in the same direction in two countries, and can even increase in the country whose currency is appreciating and decrease in the country whose currency is depreciating. As they stress, their result "is in the spirit of 'no structure' results elsewhere in the economic literature" (p. 624).

2.2 Aggregation, simultaneity and time variation

Both theoretical and empirical papers reviewed up to now suggest that, following an exchange rate movement, a wide range of outcomes can occur at the industry level. Different industries aggregate up to the whole economy, but the aggregation raises difficulties in this case, too. Among others, Parsley (1995) heavily criticises studies using aggregate price indices for the pass-through issue because of (i) measurement errors, (ii) changes in commodity composition, (iii) third-country effects and (iv) simultaneity. However, as our primary goal is to study the behaviour of aggregate inflation in response to exchange rate changes, we will use indices at some levels of aggregation but address simultaneity and bear in mind that time variation may also occur due to aggregation. The no structure result coupled with the potential time variation of each possible structure extends the set of feasible outcomes.[6]

In addition to aggregation, there are other reasons to expect time variation in parameters in the pass-through relationship. Most of the countries being studied are transition countries that moved from a planned to a market economy in the 1990s. Their behavioural relationships probably changed during the 1990s, which is the sample period analysed in this paper. Moreover, in both industrial and EU candidate countries time variation might also emerge from the changing inflationary environment. For example, J. B. Taylor (2000) argues that in the low inflation episode of the 1990s the persistence of inflation declined in the USA; therefore, firms would expect a change in costs or prices to be less persistent and fewer exchange rate changes to be passed through to prices. Cheung–Lai (2000) show a negative correlation between the inflation rate and the persistence in purchasing power parity (PPP) deviations. In most candidate countries the high inflation environment of the early 1990s gradually changed into a single-digit inflation rate episode. Consequently, we might expect this development to influence the pass-through relationship.

As for simultaneous modelling of inflation and exchange rate changes, its desirability is obvious for countries with floating exchange rates. However, the issue is fairly complex for countries with less flexible regimes. The case of those countries that operate currency boards or fixed exchange rates is clear, as there is no adjustment to the nominal exchange rate

[6] Of the literature referred to in this paper, only Kim (1990) and Parsley (1995) have performed time-varying parameter estimates, both in a single equation framework using the Kalman filter and assuming that the parameters follow random walks.

(Cyprus, Estonia, Latvia, Lithuania).[7] However, with the exception of Slovenia the countries we analyse have changed their exchange rate systems. For example, the Czech Republic had a fixed exchange rate system for several years but made its system more flexible during the period under review. This case can be modelled with time-varying parameters by setting the starting values of the parameters of the exchange rate equation to zero and observing how these parameters evolve over time. We will evaluate the adequacy of the model for countries with mixed regimes by looking at the time path of the parameters and contrasting them with our knowledge of the exchange rate regime.

2.3 Time horizon of pass-through and the real exchange rate

A crucial issue is the dynamics and time horizon of exchange rate pass-through. Most research is concentrated on the effect on prices over a relatively short horizon. Two points need to be made here.

First, the pass-through of exchange rate changes depends on whether the change is perceived to be transitory or permanent (see, for example, the pioneering works of Krugman (1987) and Froot–Klemperer (1989)). In a recent paper J. B. Taylor (2000) stresses the "expectations theory of pass-through". According to this consideration, a transitory exchange rate shock will have no or very little effect on prices. The main problem with an empirical incorporation of this effect is the measurement of transitory versus permanent shocks to the exchange rate. Froot–Klemperer use survey data for an empirical estimation. The applicability of survey data is limited even in industrial countries and data are not available consistently over a sufficiently long horizon for candidate countries. J. B. Taylor adopts a forward-looking model with pre-set parameters to simulate the effects of different types of exchange rate shocks and beliefs. Naturally, this is not the model we would choose to follow since this paper has empirical goals. Feenstra–Kendall (1997) argue that the forward exchange rate might be important due to hedging. However, in most candidate countries forward markets are either non-existent or they are used solely for speculation and not for hedging trade flows (Darvas–Szapary, 2000). Parsley (1995) also emphasises that expected future exchange rates are crucial determinants of exchange rate pass-through and gives a simple model in which the pass-through coefficient varies over time, i.e. it is a function of the expected future exchange rate. This consideration gives further support to our time-varying parameter approach.

[7] Even in a country with a fixed exchange rate linked to the currency of a major trading partner, the exchange rates vis-à-vis (some of the) other trading partners fluctuate. Therefore, it is not completely irrelevant to study pass-through in these countries.

Second, the issue of whether the real exchange rate is stationary is of major importance here. Feenstra–Kendall (1997) and Betts–Devereux (2000) argue that deviations from PPP can be explained partly by imperfect pass-through and pricing to market. Conversely, a perfect pass-through into consumer prices would imply a constant real exchange rate (defined, of course, as relative to consumer prices in a common currency). However, exchange rates can change from one second to the next whereas product prices change much less frequently, which means that the real exchange rate cannot be constant. Instead, stationarity of the real exchange rate might imply full pass-through. There is a vast body of literature analysing whether or not real exchange rates are stationary but it is not our intention to review it here.[8] After all, if the real exchange rate were stationary, the issue of the time horizon would come into play: if the half life of a shock, say, is four years, can we speak about full pass-through?

By analysing the real exchange rates of 94 countries, Cheung–Lai (2000) found that although there is substantial heterogeneity in the persistence of deviations from real exchange rate parity, parity reversion is more likely in developing than in industrial countries.[9] Bergin–Feenstra (1999) show that that the greater degree of openness of an economy limits the degree of persistent deviations from PPP. This result is consistent with the empirical findings of McCarthy (1999), who showed that higher pass-through to domestic inflation is associated with higher import shares. If these conclusions were true for candidate countries, which are not studied in Cheung–Lai and are highly open economies (Table 1), they would demonstrate the importance of incorporating reversion to the equilibrium real exchange rate in any model of pass-through.

The real exchange rate issue has to be carefully included in an analysis of candidate countries, since there are both empirical observations and theoretical arguments suggesting that the real exchange rate of these countries will appreciate in the upcoming years.[10] Naturally, expected real appreciation has decisive implications for the relationship between nominal exchange rates and product prices.

Among others, A. M. Taylor (2000), M. P. Taylor–Peel (2000) and Obstfeld–A. M. Taylor (1997) emphasise that the process of adjustment towards the law of one price or to the equilibrium real exchange rate is likely to be non-linear, as already highlighted by Heckscher in the 1910s. Time variation of the effects of various shocks is a natural consequence of non-linearity, since the effect depends on the magnitude of the shock. These considerations lend further support to time-varying analysis.

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[8] For a recent investigation of the problems associated with previous tests for purchasing power parity (PPP) and the law of one price, see A. M. Taylor (2000).

[9] Parity is proxied by stationary autoregressions, fractional autoregressions or breaking trends.

[10] See Begg–Halpern–Wyplosz (1999) for a recent assessment.

2.4 Pass-through into consumer prices

Our central focus will be the study of exchange rate pass-through into consumer prices. The reasons for this stem partly from the underlying objective of the paper (the analysis of the possible role of exchange rate management in controlling inflation) and partly from data problems. Even for industrial countries there are severe data problems in implementing a usual mark-up over cost model. As surveyed and criticised by both Menon (1995) and Goldberg–Knetter (1997), empirical studies tend, by and large, to use different cost indices, such as wholesale prices, which "may be reasonable measures of average costs incurred domestically, but are unlikely to be good measures of marginal costs, the concept relevant for pricing behaviour of a profit maximising firm. Furthermore, cost indices may introduce measurement error into equation (3) that is correlated with exchange rates in a way that biases the coefficients toward finding incomplete pass-through and excess mark-up adjustment" (Goldberg–Knetter, p. 1251).[11] This problem might be even more acute in candidate countries since they have considerable imports from less developed countries, for which finding suitable cost indices is very problematic.

As we have already indicated, most of the literature analyses pass-through into import and export prices and there are only a few papers available which deal with the analysis of pass-through into consumer prices. McCarthy (1999) studied nine industrial countries with a VAR of six variables: oil price, output gap, exchange rate, import price, producer price and consumer price. Estimated impulse response functions indicated substantial pass-through into import prices, but much less into consumer prices.[12] However, historical decompositions showed that external factors (exchange rate and import prices) played a significant role in the unexpected disinflation of the late 1990s. McCarthy also shows that higher pass-through to domestic inflation is associated with higher import shares, less volatile exchange rates, less volatile GDP and less competitiveness. Smets–Wouters (1999) also adopt a VAR framework. They studied the monetary transmission mechanism in a large open economy (Germany) and found the direct effect of an exchange rate appreciation on import prices to be rather strong and the indirect effect on consumer prices to also be of significance.

The issue of invoicing might also be important for empirical pass-through research. Goldberg–Knetter (1997) highlight the literature that emphasises the complications arising from the choice of the invoicing currency when observing pricing to market and exchange rate pass-through. In the countries examined in our study, invoicing is carried out almost

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[11] The equation number in the quotation refers to equation (3) in Goldberg–Knetter (1997).

[12] For example, the point estimate of the impulse response for two years ahead even had a wrong sign in the case of France and Switzerland, was virtually zero in Japan, the United Kingdom and Sweden, was around 15% in the USA and Germany, and was around 30% in Belgium and the Netherlands.

exclusively in a foreign currency due to the perceived uncertainties of the national currencies from the point of view of foreign partners, 80% of which, on average, are located in industrial countries. *Ceteris paribus* this would indicate complete pass-through into local import prices but not necessarily into consumer prices.

3 Pass-through in EU Candidate Countries

3.1 The model

Although extensive research into the pass-through issue has been carried out in industrial countries, there is far less empirical research for EU candidate countries.[13] Some applications for these and other developing countries can be found in Hamecz et al. (1998), Kamas (1995), Leiderman–Bufman (1996) and Lee (1997).

Based on the considerations discussed in the previous section, the model of this paper will have three important features:

(1) simultaneous analysis of price and exchange rate changes;

(2) time-varying parameters in both the pass-through and the nominal exchange rate adjustment relationship;

(3) error correction formulation incorporating the disequilibrium of the real exchange rate; the equilibrium real exchange rate model is assumed to have time-invariant parameters.

In order to ensure a reasonable measure of prices, we study the behaviour of non-food, non-energy and non-administered prices and refer to this aggregate as *fundamental prices*. Administered prices are excluded because they are determined by law and not in direct response to exchange rate changes. Food and energy prices are excluded because they are highly volatile and the former shows a marked seasonal pattern. Although several items of consumer prices were excluded, some of them, especially energy and administered prices, might affect our aggregate. In the empirical implementation, administered prices proved to be important for some countries but we were unable to find a statistically significant effect of energy prices.

One of the most difficult problems in international economics is the modelling of short-run changes of the exchange rates. We included the previous period disequilibrium of the real

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[13] For example, searching the EconLit database 06/2000 for “exchange rate pass-through” reveals no empirical studies for EU candidate countries.

exchange rate in the short-run exchange rate equation since it is central to our model. However, in the empirical section we were unable to find any other adequate explanatory variables,[14] so our exchange rate equation differs from the random walk by the time-varying drift and the error correction term.

Therefore, the core equations of the model are:

(1a) $$\Delta p_t = \beta_{0,t} + \beta_{1,t}\Delta s_t + \beta_{2,t}\Delta p_t^* + \beta_{3,t}\left(p_t^{(A)} - p_t\right) + \beta_{4,t}\left(q_{t-1} - q_{t-1}^{EQ}\right) + u_{1,t}$$

(1b) $$\Delta s_t = \gamma_{0,t} + \gamma_{1,t}\left(q_{t-1} - q_{t-1}^{EQ}\right) + u_{2,t}$$

where Δ is the first difference operator, p_t is the domestic and p_t^* is the foreign fundamental price level, s_t is the nominal exchange rate, $p_t^{(A)}$ is the price of goods controlled or "strongly" influenced by administrative measures, $q_t \equiv p_t - p_t^* - s_t$ is the real exchange rate, q_t^{EQ} is the equilibrium real exchange rate, $\beta_{i,t}$ and $\gamma_{i,t}$ are time-varying parameters, and $u_{i,t}$ are the residuals.

The reason for including administered prices in relative terms is that they are not additional to the exchange rate changes, foreign inflation and the effect arising from error correction. We expect an "additional" inflationary effect only if their increase is faster than that of fundamental prices.

In the model there are no instantaneous price effects on the exchange rate. However, price changes still influence the exchange rate via the error correction term.

In our interpretation the time-varying intercepts in the equations represent autonomous changes not explained by other factors.

Estimation of the equilibrium real exchange rate might be done by various ways (see Annex 8.2). We adopt the so-called reduced form approach to the equilibrium real exchange rate, which consists of estimating a static, single-equation model for the real exchange rate:

(2) $$q_t = \alpha' x_t + v_t$$

where x_t denotes the ($l\times1$) vector of determinants of the real exchange rate with coefficient vector α, and v_t is the residual. The fundamentals used to associate with the equilibrium rate are productivity developments, terms of trade, world real interest rate, net foreign assets,

[14] We tried using the change in the interest rate differential and the change in stock prices; the later was intended to proxy short-run shocks. We also tried contemporaneous change in prices, but its parameter varied substantially across sub-samples, had a huge standard error and the point estimate was frequently outside the [0.1] interval.

government policy measures (consumption and controls over capital flows and trade) and the investment ratio. Justification for using this approach is given in the annex.

Parameters of the equilibrium real exchange rate model are assumed to be time-invariant. From a theoretical perspective, it might be said that any equilibrium concept should be a long-run phenomenon with fixed parameters. However, it could also be claimed that the parameters might have changed as a result of the major structural changes experienced by these countries even after 1993. Second, practical considerations warn us not to allow every parameter to vary in time in an eight-year sample since we would thus lose a considerable degree of freedom. Fortunately, Chow tests presented later do not indicate structural breaks in the real exchange rate equation in all cases.

Finally, we should emphasise that the system in (1) is not an error correction representation of some variables.

3.2 Statistical representation

The model consists of two major blocks (equilibrium real exchange rate and pass-through) which will be separated in estimation. For the statistical representation of the pass-through relationships we use a state-space formulation and the Kalman filter for modelling time-varying parameters. We could have estimated the parameters of the long-run real exchange rate equation within the state-space models, but it proved to be sensitive to starting values of the maximisation. Therefore, we estimated the parameters of the equilibrium real exchange rate as a single equation and imported the parameters into the state-space pass-through set-up.

The general structure of state-space models is the following. The so-called state or transition equations model the unobserved variables and time-varying parameters:

(3) $$\xi_t = T_t \xi_{t-1} + c_t + R_t \eta_t$$

where ξ_t is an $(m\times 1)$ vector of unobserved variables, T_t is an $(m\times m)$ matrix, c_t is an $(m\times 1)$ vector, R_t is an $(m\times g)$ matrix, and η_t is an $(g\times 1)$ vector of white noise series. In our model $m = g = 7$ and $\xi_t \equiv \begin{pmatrix}\beta_{0,t} & \beta_{1,t} & \beta_{2,t} & \beta_{3,t} & \beta_{4,t} & \gamma_{0,t} & \gamma_{1,t}\end{pmatrix}'$. The space or measurement equations are:

(4) $$y_t = Z_t \xi_t + d_t + \varepsilon_t \quad ,$$

where y_t is an $(n\times 1)$ vector of observed variables, Z_t is an $(n\times m)$ matrix, d_t is an $(n\times 1)$ vector, and ε_t is an $(n\times 1)$ vector of white noise series. In our model $n = 2$, $y_t \equiv \begin{pmatrix}\Delta p_t & \Delta s_t\end{pmatrix}'$, $Z_t = \begin{bmatrix} 1 & \Delta s_t & \Delta p_t^* & \Delta\left(p_t^{(A)} - p_t\right) & ecm_t & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & ecm_t \end{bmatrix}$, where $ecm_t = \left(p_{t-1} - s_{t-1} - p_{t-1}^*\right) - \alpha' x_{t-1}$

and d_t = 0. The white noise processes in (3) and (4) are uncorrelated and have variance-covariance matrices: var(η_t)=Q_t, var(ε_t)=H_t.

We assume that the time-varying parameters (ξ_t) follow random walks, in which case $T_t = R_t = I_m$, where I_m is the ($m\times m$) identity matrix, and c_t = 0.

Given the parameters of the model, Kalman filtering comprises a series of recursive steps in producing forecasts of ξ_t. The first step is the forecast of ξ_1 given initial conditions, which is denoted by $\hat{\xi}_{1|0}$, the second step is the calculation of $\hat{\xi}_{2|1}$ given $\hat{\xi}_{1|0}$ and data at time t=1, and so on until $\hat{\xi}_{T|T-1}$. Each of these forecasts is associated with an ($m\times m$) mean squared error matrix $P_{t|t-1} = E\left[\left(\xi_t - \hat{\xi}_{t|t-1}\right)\left(\xi_t - \hat{\xi}_{t|t-1}\right)'\right]$, which is also calculated recursively.

When parameters are unknown, the Kalman filter can be used to evaluate the likelihood function. The likelihood function is written in terms of the one period ahead conditional forecast of y_t, denoted as $\hat{y}_{t|t-1}$. The assumption that $\{\eta_t, \varepsilon_t\}_{t=1}^{T}$ is multivariate Gaussian allows a convenient representation of the sample log-likelihood function. However, even if the innovations are non-Gaussian, maximisation of the likelihood function (referred to as the quasi maximum likelihood estimation) yields consistent and asymptotically normal estimates of the parameters.[15]

In our model the parameters to be estimated are the coefficients of the long-run equation (α_i) and the elements of Q_t and H_t. We assume that $Q_t = Q$ and $H_t = H$ are time-invariant, but even in this case Q has $g(g+1)/2$ parameters, so it is not possible to estimate unconstrained parameters using our sample (which consists of about 30 observations). There are two possible solutions to this problem:

(1) Estimate the coefficient as fixed by OLS, calculate the estimated variance-covariance matrix of parameters $(\hat{\Omega})$ and assume that $Q = \mu\hat{\Omega}$, where μ is a parameter to be estimated.[16]

(2) Assume that Q is diagonal and estimate only its diagonal elements.

[15] See Chapter 13 of Hamilton (1994) for an excellent discussion of Kalman filtering.

[16] This approach is adopted in Kim (1990). The model in Parsley (1995) has only one time-varying coefficient.

We tried both options and decided for the second due to unstable OLS estimates. We also had to assume that H is a diagonal (2×2) matrix for a practical reason: the maximisation never converged when we allowed the estimation of $\mathrm{cov}\left(\varepsilon_t^{(\Delta p_t)}\varepsilon_t^{(\Delta s_t)}\right)$.

Initial conditions should be determined for ξ_0 and P_0, where P_0 expresses confidence in ξ_0. In most cases we have set ξ_0 to the parameter values obtained by a fixed parameter estimate for the first three years of the sample and set P_0 to be the equal to the variance-covariance matrix of this estimate.[17]

4 Exchange Rate Regimes, Inflation and Real Exchange Rate

4.1 Exchange Rate Regimes

Until May 1997 the Czech Republic had a fixed exchange rate regime linked to a basket of Deutsche Mark (65%) and US dollar (35%). The width of the band was essentially zero until February 1996, when it was widened to ±7.5%. In May 1997 the koruna was allowed to float in response to a speculative attack and inflation targeting was announced. As a consequence of exchange rate pegging until February 1996, we do not expect to see exchange rate pass-through before this date and the time-varying coefficients in the exchange rate equation are also expected to be zero until then.[18]

Hungary had an adjustable peg until the first half of the 1990s. Inflation burst in the early 1990s was handled by some large step devaluation during that time and several smaller adjustments were made in the period 1992-94 (i.e. five to seven cases a year with magnitudes of 1-3%, with the exception of August 1994 when there was an 8% devaluation). Difficulties raised by the adjustable peg and a huge current account deficit were major factors for introducing a pre-announced crawling band as part of the stabilisation program in March 1995, following a 9% devaluation. The monetary regime of Hungary can be characterised as an exchange rate anchored regime where, taking into account the faster rate of productivity growth in Hungary compared with its trading partners, the rate of crawl is set somewhat below the expected rate of inflation differential so as to help disinflation while also maintaining competitiveness.

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[17] In the case of the Czech Republic, which had a fixed exchange rate regime at the start of our sample period, we set the initial value of the parameter of the exchange rate in the price equation and the two parameters of the exchange rate equation to zero. Since we are of this choice, we set the standard error to 0.01. Since we set the initial value of the parameter of the exchange rate to zero in the price equation instead of its OLS estimate, we have to modify the intercept as well. This was done by rearranging the price equation for the intercept and inserting the initial values of the parameters and the first observations of the variables.

[18] Figures 20, 21, 22 and 23 in the annex show the level of exchange rates at daily frequency that is effectively illustrative of the exchange rate regimes adopted by these countries.

Poland had a pre-announced crawling peg from October 1991 – first with no band, then with a band which was widened in steps to ±15%. The introduction of the crawling band was not preceded by a devaluation, but two discrete devaluations of 10.7% and 7.4% and a revaluation of 6% were effected subsequently. There were marked central bank interventions to prevent an appreciation until early 1998, when the authorities adopted an inflation-targeting regime and allowed the exchange rate to fluctuate across the full band. In April 2000 the band was abandoned and the zloty was allowed to float.

Slovenia has a highly managed floating exchange rate regime.

4.2 Nominal exchange rate and inflation developments

Figures 1, 2, 3 and 4 plot consumer price inflation, fundamental inflation and exchange rate changes. The exchange rate measure shown in the figures are calculated against a weighted average of Deutsche mark and US dollar. The weights used are representative for the official currency basket adopted by these countries.[19] As expected, the floating exchange rate regime of the Czech Republic and Poland was characterised by marked exchange rate fluctuations compared to Hungary and Slovenia.

The Czech Republic could achieve the lowest inflation rate among the four countries.

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[19] See Table 2 for the weights used. In the case of the Czech Republic, prior to Q2 1993 the composition of the official currency basket was not 65% DEM-35% USD (the weights used), which explains why our exchange rate measure shows appreciation in Q1 1993.

Figure 1. Czech Republic: Inflation and exchange rate changes, Q1 1993 – Q2 2000

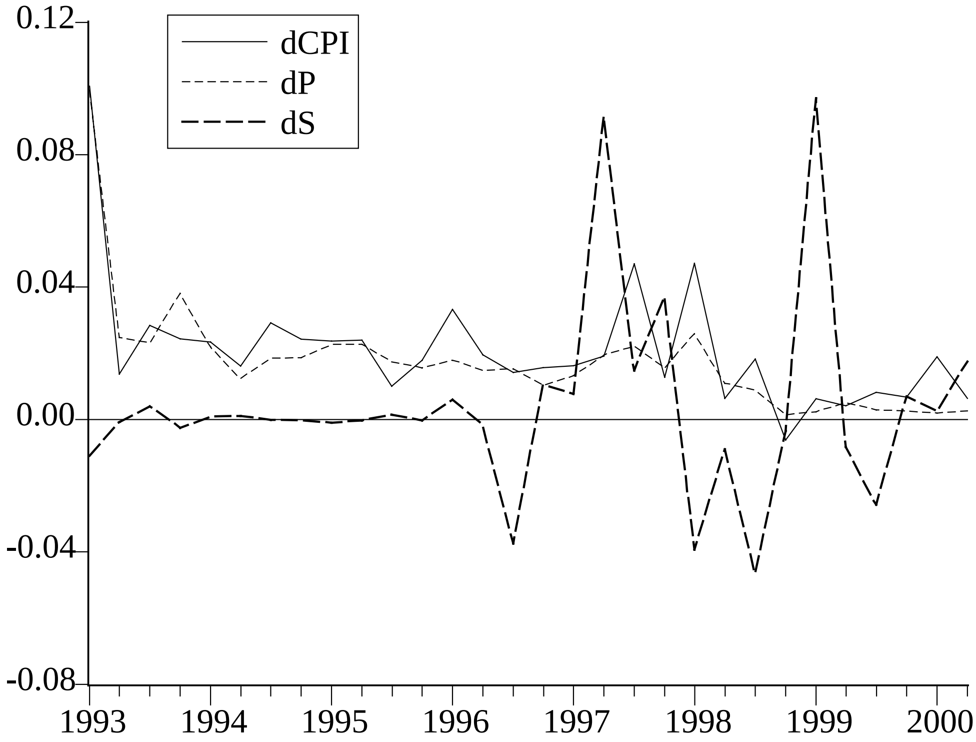


Figure 2. Hungary: Inflation and exchange rate changes, Q1 1992 – Q2 2000

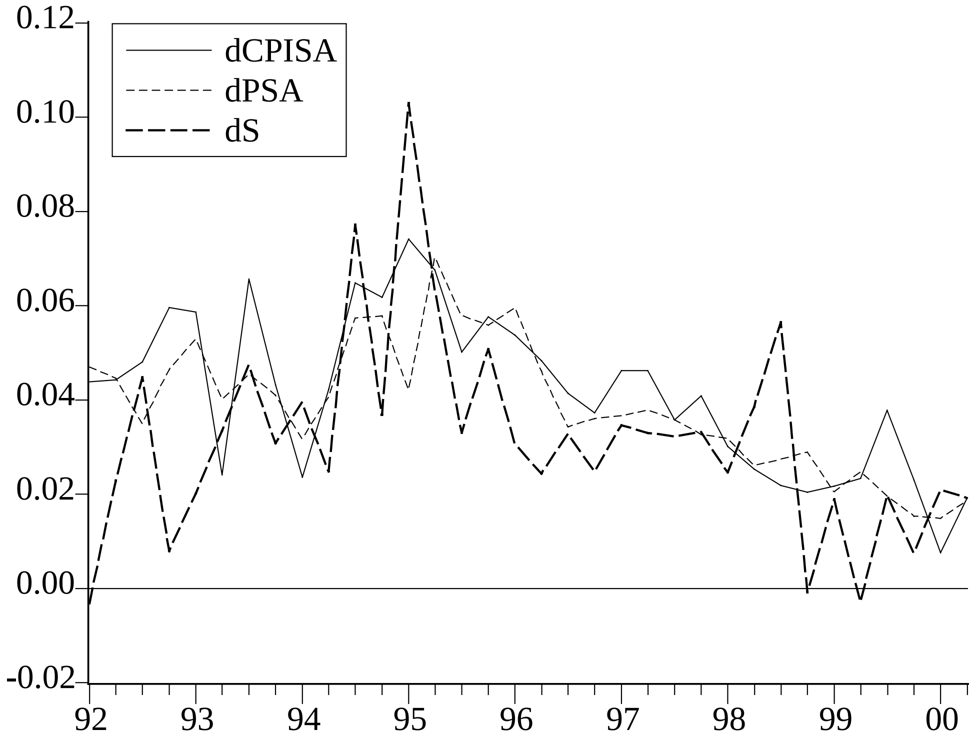


Figure 3. Poland: Inflation and exchange rate changes, Q1 1992 – Q2 2000

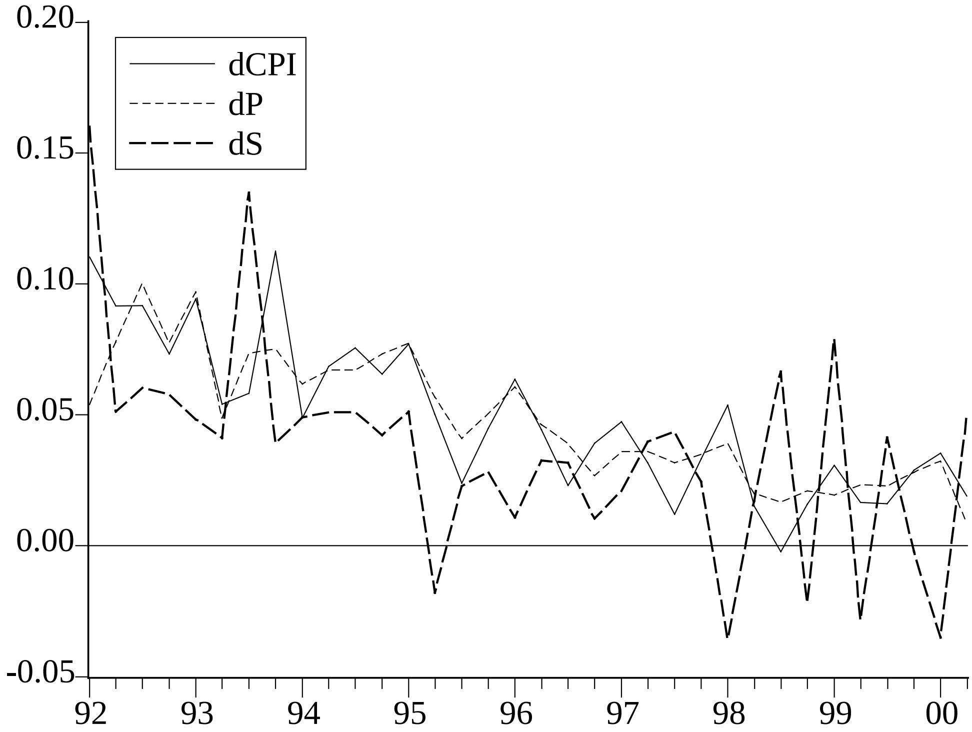
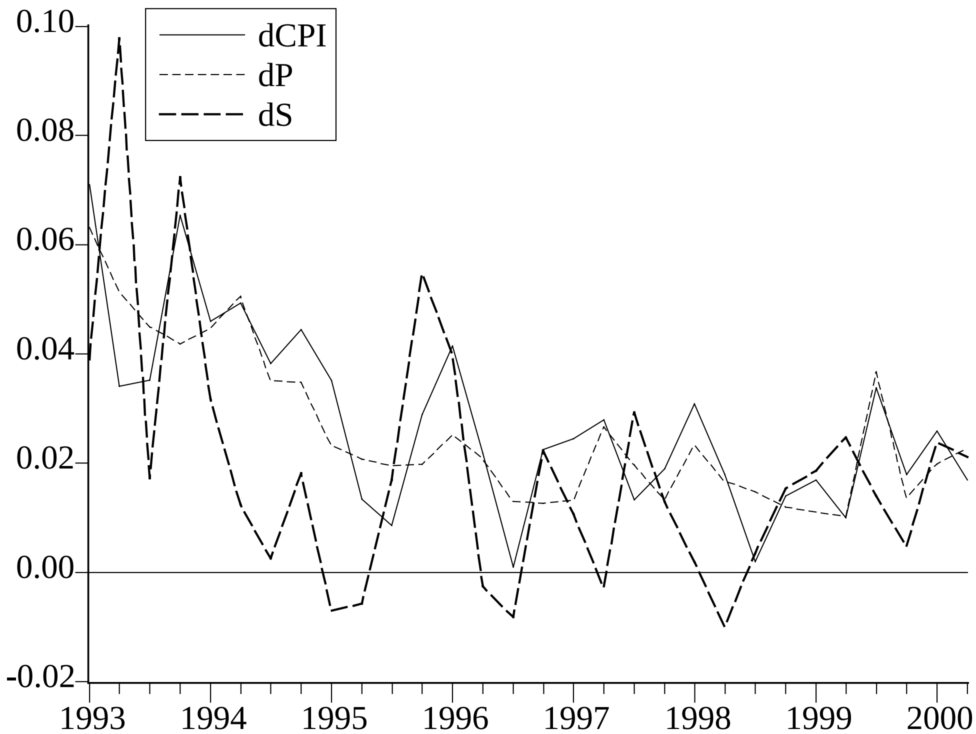


Figure 4. Slovenia: Inflation and exchange rate changes, Q1 1993 – Q2 2000



4.3 Real exchange rate and the Balassa-Samuelson effect

Figure 5 shows real exchange rates with respect of the Deutsche mark calculated on the basis of fundamental prices. There are several possible explanations for the real exchange rate appreciation in the early years of transition (Halpern–Wyplosz, 1997), of which the Balassa-Samuelson (BS) effect has a principal importance. The BS effect says that when productivity in the tradable sector relative to the nontradable sector increases faster than abroad, the real exchange rate appreciates.[20] This effect is, of course, not a special characteristic of transition countries, and should take effect in any country in any period provided its assumptions are fulfilled.

Figure 6 shows the level of labour productivity differential between the two sectors in case of the four candidate countries, Germany, and US. It is striking how much the rate of increase in relative productivity in Hungary exceeds the rates of the rest of countries. However, we know that this feature is in contrast to the paths of real exchange rates where the rate of real appreciation of the Hungarian forint lags behind the other transition countries, in the period Q1 1992 – Q3 2000. There are several possible explanations for this apparent anomaly:

- The other three countries started the transition period with a substantially undervalued currency while Hungary did not. In light of this starting conditions the process of the strong appreciation of their currencies might be the result of a gradual restructuring of relative (tradable/nontradable) prices towards the long-run equilibrium. In Hungary, due to its heritage of a system that was closer to a market economy this effect might have been much smaller.
- Wages in the Hungarian tradable sector might probably be lower than their marginal product. This suspicion cannot be confirmed statistically but it is supported by the fact that unit labor costs in the tradable sector decreased markedly in the past ten years. Consequently, profits in this sector increased sharply and are probably higher than in the nontradable sector.
- A considerable part, maybe more than half of the tradable sector is owned by foreigners. The persistent divergence between profits of the tradable versus the non-tradable sector might be explained by some monopolistic power of these foreign investors, arising from their know-how, their financial strength that makes them easier to pay the high entry costs

[20] The crucial assumptions are that (1) labour is paid by its marginal product in the tradable sector, (2) wages equalise between the tradable and nontradable sectors, (3) both capital and labour is mobile between the two sectors, (4) the prices of tradable goods are equal at home and abroad.

in the industry and the informational or risk barriers that prevent profits from equalizing internationally.

- Hungarian workers are weakly organised and labor is still abundant. This might probably prevent labor from achieving better wage bargains.

The nine-year period shown on the figures is probably too short and too much burdened with structural changes for allowing the market to force out an elimination of the above mentioned imperfections.

It is interesting, that while relative productivities changed dramatically in Hungary, this phenomenon was not a by-product of an outstanding general productivity increase among the countries in the region. As it is shown in Figure 7, in terms of the level of labour productivity measured as GDP per employee Hungary is only in the midfield among the candidate countries, implying that productivity increases in the nontradable sector were smaller than in the other countries.

Figure 5. Real exchange rate with respect to Germany, Q1 1991 – Q4 2000

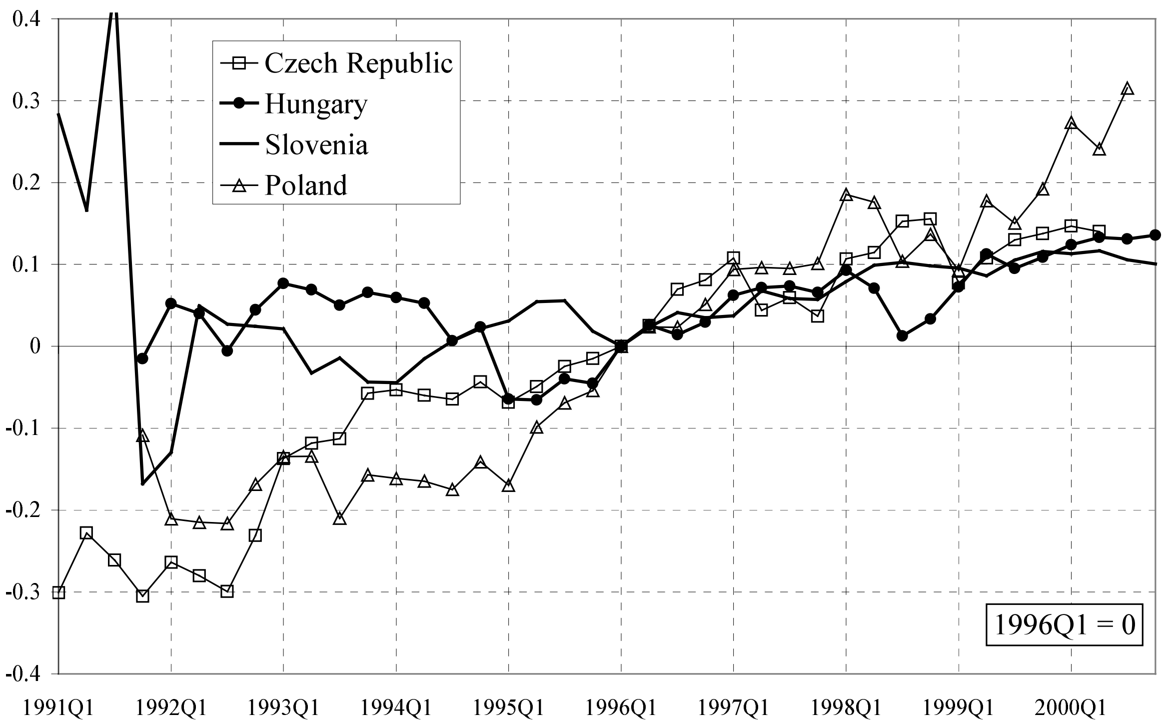
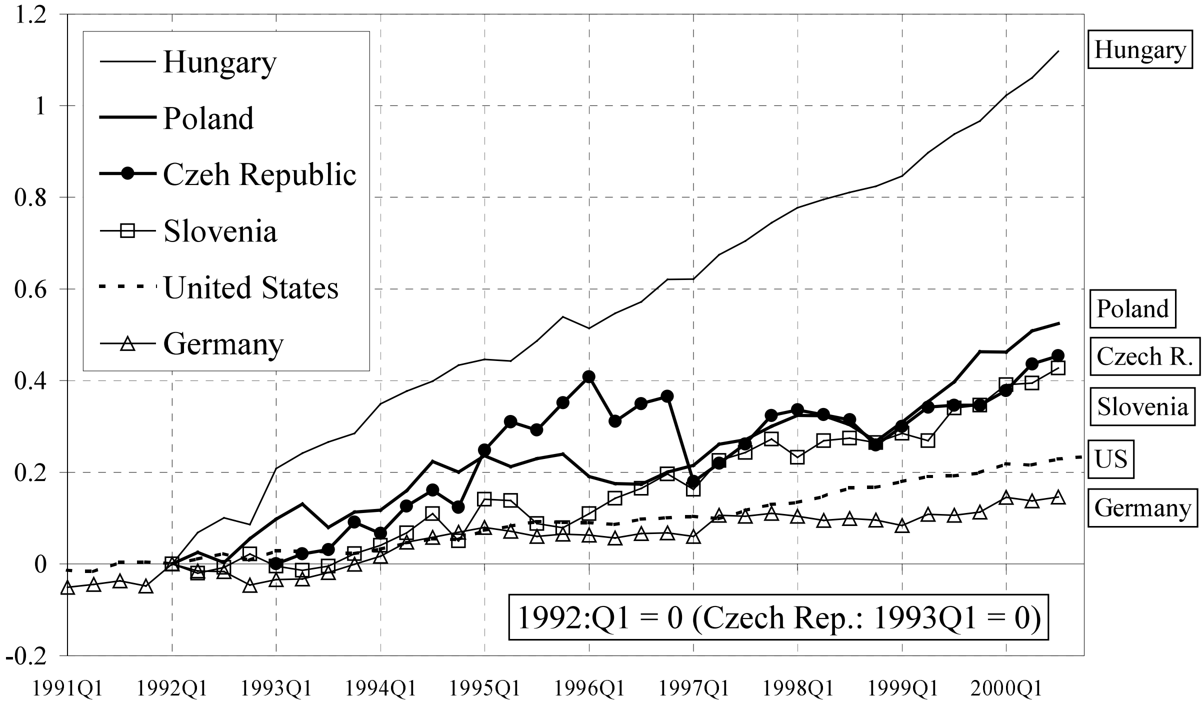
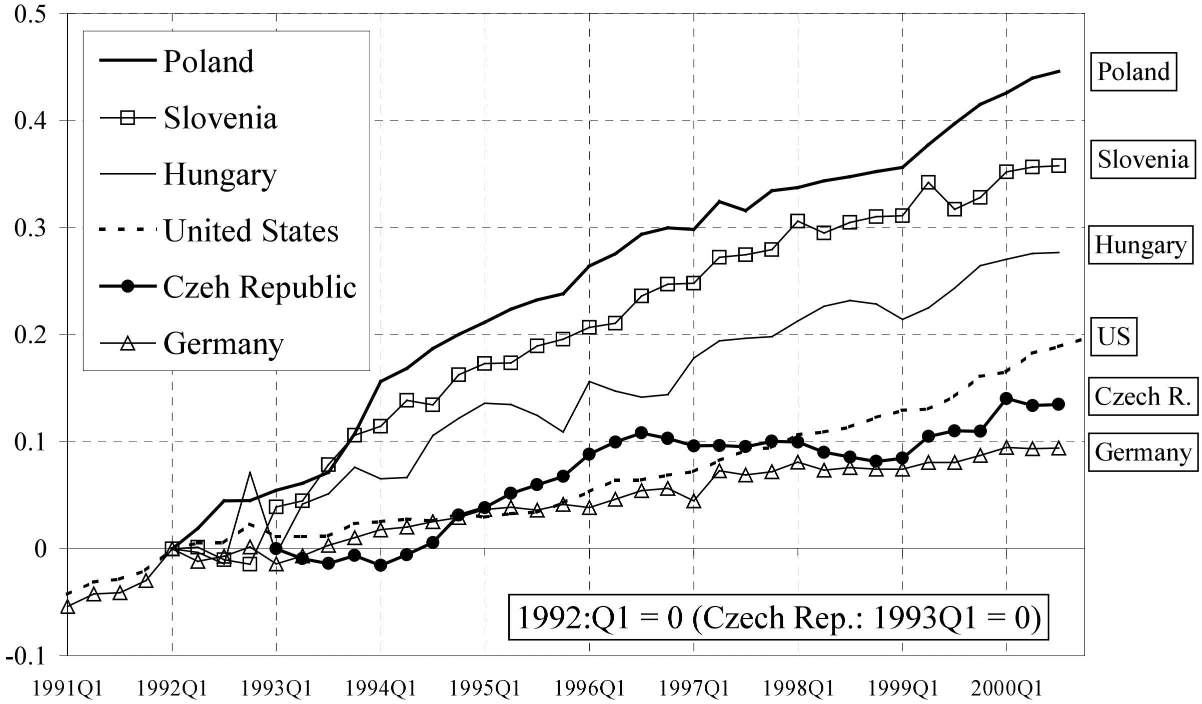


Figure 6. Labour productivity divergence in the tradable vs. non-tradable sector, Q1 1991 – Q4 2000



Note: values shown are calculated as $\ln[(Y^T/L^T)/(Y^{NT}/L^{NT})]$.

Figure 7. Labour productivity of GDP, Q1 1991 – Q4 2000



Note: values shown are calculated as ln(GDP/L).

5 Empirical results

Empirical analyses were conducted country by country. In this section of the paper we present the final results for all countries; some further details are shown in the annex. The analysis consists of the following steps.

- Analysis of seasonality and unit root tests

  The short time-span renders seasonal unit root testing inappropriate and we do not attempt to set up a seasonal error correction model. In addition, some production measures included in the long-run equations are available on a seasonally adjusted basis only. Therefore, we aimed to work with data that do not contain seasonal components, adjusting the series seasonally only if a strong seasonal pattern was determined by the correlogram of the series and its differences.[21]

  We used Phillips–Perron (PP) and augmented Dickey–Fuller (ADF) *t*-tests, with both (i) constant and (ii) constant and trend in the test equation. We selected the truncation lag for the variance estimate need for the PP test by the rule of thumb suggested by Newey–West. For the ADF, there are several methods for selecting the appropriate lag in the test equation. We used three of them: (1) starting from the specification having no lagged differences and increasing the number of lags one at a time if there is autocorrelation in the residuals, (2) staring with a specification having many lagged differences and decreasing their number one by one until the number becomes significant, (3) selecting the test equation with the smallest Schwarz criterion value. The longest possible lag length for all three methods was set to 8.

- Pass-through estimates with vector autoregression
- Estimation of the long-run real exchange rate with static regression for the levels (Engle–Granger)

  We experimented with the VECM approach of Johansen and several single equation ECM-based tests as well.[22] However, since these methods require the estimation of many parameters, in our short sample they proved to be very unstable and sensitive to minor changes in specification.

[21] As the inflation measure that we adopt eliminates food and administered prices, most sources of seasonality are probably removed. As a consequence, we did not automatically adjust our fundamental price indices but first tested for the presence of seasonality.

[22] For a comprehensive survey of different methods of estimating the cointegrating vectors, see Maddala–Kim (1998).

- Fixed parameter estimation of the pass-through relationship

  The full period estimation was tested using a Chow breakpoint test and was also performed at a rolling window of three years to give a feeling of the stability of parameters.

- Time-varying parameter estimates of the pass-through model.

Table 2 shows the definitions and abbreviations of the variables.

Table 2. Variable definitions [1,2]

Abbreviation	Description
p_t	Domestic fundamental prices, defined as non-food, non-energy and non-administered prices, last month of the quarter
s_t	Exchange rate, last month of the quarter
p_t*	Foreign fundamental prices, last month of the quarter
rpa_t	$= p^{(A)}_t - p_t$, Relative prices of administered goods, last month of the quarter
rpe_t	$= p^{(E)}_t - p_t$, Relative prices of energy, last month of the quarter
$idif_t$	$= i_t - i_t*$, Interest rate differential calculated from annualised three-month nominal interest rates, last month of the quarter
$prod^{(GDP)}_t$	$= (gdp_t-l_t)-(gdp_t*-l_t^*)$, Productivity differential measured as GDP per employee relative to abroad; individual GDP figures are seasonally adjusted
$prod^{(T-NT)}_t$	$= [(y_t^T-l_t^T)-(y_t^{NT}-l_t^{NT})]-[(y_t^{T*}-l_t^{T*})-(y_t^{NT*}-l_t^{NT*})]$, Productivity differential measured as tradable productivity minus nontradable productivity at home relative to abroad, individual figures are seasonally adjusted
tot_t	Terms of trade
$sgov_t$	Share of government expenditures in GDP, seasonally adjusted
nfa_t	Net foreign assets (last day of the quarter) over GDP, GDP is measured as a four-quarter centred moving average of seasonally adjusted figures
fdi_t	FDI liabilities (last day of the quarter) over GDP, GDP is measured as a four-quarter centred moving average of seasonally adjusted figures
$onfa_t$	$= nfa_t - fdi_t$, Other net foreign assets over GDP
r_t*	$= i_t* - (p_t*- p_{t-4}*)$, Foreign real interest rate, annualised yield on three-month treasury bill less four-quarter inflation, last month of the quarter

(1) With the exception of variables expressed as a percentage of GDP, all other variables are in logs (in case of nominal interest rates *ln*(1+*yield*) was used).
(2) For foreign variables needed, we weighted German and US data using the following weights: Czech Republic: 65%-35%, Hungary: 70%-30%, Poland: 60%-40%, Slovenia: 90%-10%. These weights closely represent the official currency baskets of the countries.

Data availability

Czech Republic: All data are available for the analysis and, with one exceptions, they cover the full period of 1993-2000. The exception relates is quarterly GDP, which are available from 1994. For 1993 we approximated quarterly GDP using the method developed at the National Bank of Hungary.[23]

Hungary: Quarterly national accounts data are available only from 1995. Our data up to 1995 was approximated. All other data are available for the full period.

Poland: Fundamental prices are measured with error. We had aggregates for administered prices, food prices, and energy prices, but not their weights, so we assumed that Polish weights are similar to those of the other three countries in order to be able to clean them away from CPI. Quarterly GDP was approximated for the full period.[24] The international investment position is available only at annual frequency and from 1997; NFA, FDI, and ONFA are therefore omitted. The remaining data were available.

Slovenia: The international investment position was available only at annual frequency, so we approximated quarterly figures using the quarterly balance of payments.[25] Terms of trade figures were approximated by the Research Department of the Bank of Slovenia. All other data are available for the full period.

5.1 Preliminary data analysis

Fundamental inflation in the Czech Republic, Poland and Slovenia showed no seasonal pattern, while Hungarian and foreign inflation did. Therefore, we made adjustments to the latter two only. Among the other variables, the share of government and relative administered prices needed seasonal adjustment. Unit root tests indicated that the series can be well approximated as I(1). (See details in the annex.)

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[23] I am grateful to Viktor Várpalotai, who approximated missing quarterly GDP figures in case of the Czech Republic, Hungary, and Poland.

[24] At a quarterly frequency current price quarterly national accounts data are available since 1995 and real growth rates compared to the same quarter of the previous year since 1996.

[25] We assumed that the movement of the stock within a given year is proportional to the accumulated current account balance (in the case of NFA) and accumulated FDI (in the case of FDI) of the same year.

5.2 VAR pass-through estimates

An impulse response function derived from a vector autoregression (VAR) might give a rough picture of pass-through. OLS estimation of a VAR for the level of some series delivers consistent, albeit inefficient, estimates even if the series have unit roots and are cointegrated. However, impulse response analysis is ambiguous when innovations are contemporaneously correlated. The usual solution for this problem is to set up a structural VAR (SVAR) in order to identify orthogonal shocks. However, we did not pursue this particular option (which seemed hopeless) but adopted the triangular factorisation of positive definite symmetric matrices to calculate orthogonalised innovations. The latter depends on the ordering of variables so we checked two different orderings. We have normalised impulse response functions to show the effect of a 1% increase in the orthogonalised innovation of the exchange rate.

We have estimated three-variable vector autoregressions for domestic fundamental prices, exchange rate and foreign fundamental prices.[26] We used two methods for selecting the optimal number of lags: (i) the likelihood ratio test and (ii) the Ljung-Box test for autocorrelation of all three residuals. In order to gain an impression of the stability of the inference, we used two different samples, the first starting in 1992 and the second in 1994.

Table 3 shows the selected lag lengths, the resulting sample period, the number of estimated parameters per equation, the correlation of (unadjusted) innovations and impulse responses of prices to a 1% shock to the orthogonalised exchange rate innovation in one year and in three years.

The general impression is that impulse response functions are imprecisely estimated, sensitive to the sample period, to the lag length and to the correlation between unadjusted innovations.

The point estimates of Czech figures are virtually zero in all cases and are never significant. Only Polish results seem significant in the vast majority of cases. In the case of Hungary and Slovenia, one-year impulses are mostly significant, whereas three-year impulses are not.

If any conclusion can be drawn from this experiment, the indication is that the point estimates of pass-through are higher in Hungary and Slovenia than in Poland and there is zero pass-through in the Czech Republic.

[26] The reported results refer to the case when a constant is included in the VAR. We also estimated VARs including a deterministic time trend that lead to qualitatively the same results.

Table 3. Impulse response functions derived from three variable VARs

	lag	*T*	*k*	ρ_{12}	ρ_{13}	ρ_{23}	*i*1(1y)	*i*1(3y)	*i*2(1y)	*i*2(3y)
Sample: Q1 1992 – Q2 2000, lag selection: LR										
Czech Rep.	5	29	16	-0.36	0.04	0.03	0.01 (0.03)	-0.08 (0.10)	-0.01 (0.03)	-0.08 (0.13)
Hungary	4	30	13	0.63	0.51	0.45	0.22 (0.11)	0.02 (0.14)	0.41 (0.12)	0.19 (0.14)
Poland	7	27	22	-0.06	0.02	-0.77	0.13 (0.03)	0.12 (0.06)	0.06 (0.02)	0.07 (0.03)
Slovenia	6	28	19	0.73	0.39	0.02	0.12 (0.08)	0.22 (0.32)	0.49 (0.17)	0.95 (1.16)
Sample: Q1 1992 – Q2 2000, lag selection: LB										
Czech Rep.	5	29	16	-0.36	0.04	0.03	0.01 (0.03)	-0.08 (0.10)	-0.01 (0.03)	-0.08 (0.13)
Hungary	3	31	10	0.50	0.50	0.24	0.31 (0.10)	0.19 (0.11)	0.47 (0.12)	0.22 (0.19)
Poland	1	33	4	0.15	0.30	-0.12	0.28 (0.09)	0.11 (0.10)	0.31 (0.10)	0.14 (0.09)
Slovenia	2	32	7	0.24	0.02	-0.22	0.13 (0.13)	0.06 (0.17)	0.27 (0.17)	0.16 (0.20)
Sample: Q1 1994 – Q2 2000, lag selection: LR										
Czech Rep.	4	22	13	-0.21	-0.21	-0.44	0.00 (0.04)	0.01 (0.01)	-0.01 (0.05)	-0.14 (0.12)
Hungary	3	23	10	0.41	0.26	0.29	0.28 (0.07)	0.11 (0.13)	0.39 (0.10)	0.19 (0.10)
Poland	1	25	4	-0.03	-0.08	-0.09	0.14 (0.06)	0.06 (0.06)	0.13 (0.07)	0.06 (0.03)
Slovenia	1	25	4	0.28	-0.01	-0.07	0.38 (0.15)	0.50 (0.44)	0.48 (0.17)	0.60 (0.50)
Sample: Q1 1992 – Q2 2000, lag selection: LB										
Czech Rep.	2	24	7	-0.02	-0.09	-0.23	0.04 (0.06)	-0.03 (0.03)	0.05 (0.06)	-0.01 (0.03)
Hungary	1	25	4	0.33	0.22	0.22	0.66 (0.20)	0.21 (0.21)	0.72 (0.21)	0.20 (0.23)
Poland	5	21	16	-0.48	-0.88	0.29	0.06 (0.03)	0.09 (0.05)	0.07 (0.03)	0.10 (0.05)
Slovenia	2	24	7	0.21	0.01	0.20	0.26 (0.16)	0.18 (0.24)	0.38 (0.18)	0.42 (0.39)

Notes: VARs were estimated for p_t, s_t, p_t*. The optimal number of lags was selected by the likelihood ratio test (LR) and the Ljung-Box test for autocorrelation of all three residuals (LB). *lag*: lag length; *T*: number of usable observation; *k*: number of parameters per equation; ρ_{ij}: correlation of unadjusted innovations; *i*1(.) and *i*2(.): impulse response of prices to a 1% shock to the exchange rate with variable ordering p_t, s_t, p_t* and p_t*, s_t, p_t respectively, in one year (1y) and in three years (3y), standard errors in brackets.

5.3 Long-run real exchange rate

The selected specifications are shown in Table 4. Among the two productivity measures, relative GDP productivity was selected for the Czech Republic and Poland, while the relative sectoral productivity variable proved to be better in case of Hungary and Slovenia. The huge Hungarian productivity increase and small real appreciation is reflected in the very low estimated coefficient.

The foreign real interest rate was important for all four countries with reasonable parameter estimates. The ONFA position was positive in the case of the Czech Republic and Slovenia and negative in the case of Hungary and Poland, indicating that the former two were net lenders and the latter two were net debtors in respect of the rest of the world; a rise in foreign rates therefore indicates that net interest income from abroad increases in the former two and decreases in the latter two countries. As a consequence, we would expect a rise in foreign interest rates to have a positive impact in the first two countries and a negative impact in the latter two.

As we have discussed in the section on data availability, the stock variables of foreign assets and liabilities are properly measured at a quarterly frequency only in case of the Czech Republic and Hungary. In the Czech Republic only FDI and for Hungary only total NFA attained the correct sign. According to our estimates, FDI inflows cause the Czech equilibrium real exchange rate to appreciate.[27] In the case of Slovenia, for which quarterly frequency was approximated, these variables had correct signs but were insignificant, so we decided not to include them.

The parameter of the terms of trade had the wrong sign in the case of Hungary and Poland, and the parameter of the share of government had the wrong sign in the two countries for which this variable was available (Czech Republic, Hungary).

We tested for parameter stability using Chow breakpoint tests for the middle of the sample period (Q3 1996). Although the distribution of this test statistic is not known for non-stationary variables, with exception of Hungary the F-values are small enough for it to be safe to conjecture that there is no structural break in the parameters.

[27] Note that the stock of FDI is measured as a liability.

Table 4. Estimates of the equilibrium real exchange rate

	Czech Republic	Hungary	Poland	Slovenia
Const.	-3.19	-3.84	-0.06	-3.43
t	-193.88	-50.63	-1.55	-313.38
$prod^{(GDP)}$	1.15		1.81	
t	6.15		5.77	
$prod^{(T\text{-}NT)}$		0.073		0.312
t		1.76		3.22
tot	1.64			
t	5.79			
nfa		0.069		
t		1.83		
fdi	-0.055			
t	-2.13			
r^*	0.025	-0.013	-0.037	0.022
t	3.52	-1.27	-1.64	3.62
First obs.	Q1 1993	Q1 1993	Q1 1993	Q1 1993
Last obs.	Q1 2000	Q1 2000	Q1 2000	Q1 2000
T	29	29	29	29
R^2	0.854	0.306	0.859	0.713
Adjusted R^2	0.830	0.223	0.849	0.692
DW	1.420	1.048	0.902	0.979
AIC	-4.098	-4.712	-3.349	-4.589
BIC	-3.862	-4.523	-3.208	-4.448
SE	0.029	0.022	0.043	0.023
Chow	0.75	4.03	2.87	0.51
PP	-4.15**	-3.18	-3.31**	-3.36**
ADF	-4.71***	-3.40	-3.05*	-3.69**

Notes: *t*: *t*-ratio calculated with Newey–West heteroskedasticity and autocorrelation consistent standard error; *First obs.*: first observation; *Last obs.*: last observation; *T*: number of observations; *DW*: Durbin-Watson; *AIC*: Akaike's information criterion, *BIC*: Schwarz information criterion, *SE*: standard error of regression, *Chow*: F-value of the Chow breakpoint test for Q3 1996; *PP*: PP test for the residuals; *ADF*: ADF test for the residuals; ***, **, or *: significant at 1%, 5%, or 10%.

The last two rows of the table report cointegration tests (unit root tests for the residuals of the regressions). In case of Hungary, the null of no cointegration is not rejected at usual significance levels. We still insist on using this model for two reasons, because power and size distortions of unit root and stationarity tests are highly documented in the literature. Our sample is relatively short, so the scope for misjudgement is large. The test statistics are quite close to the critical value. Therefore, bearing in mind the problems with unit root testing, a statistic which is significant only at around 12% does not provide strong evidence against cointegration.

5.4 Fixed parameter pass-through estimates

First we estimated model (1) with OLS separately for the two equations in order to test for structural break. In the case of the Czech Republic, it is clear from Figure 1 that it does not make sense to estimate the model for the period 1993-2000 with fixed parameters. The exchange rate was virtually unchanged until February 1996 when a band was introduced and fluctuated sharply since May 1997 when it was allowed to float. Pass-through and the impact of other variables in the exchange rate are, therefore, bound to change. For this reason, we report results for the period starting in the third quarter of 1997, but we should keep in mind that the number of observations is only 12 in this case.

The coefficient of foreign prices was estimated to be very large (with standard errors given in brackets): Czech Republic: 2.23 (0.99), Hungary: 2.01 (0.51), Poland: 5.37 (1.46), Slovenia: 2.59 (0.56). These estimates are significantly different from zero and in two cases are even significantly larger than 1. Since there is no satisfactory economic explanation for a coefficient outside the [0.1] interval, we constrained the parameter to be 1 in all four cases. We can interpret this choice as modelling the inflation differential.

The coefficient of administrative prices was insignificant and even the point estimate had a wrong sign in case of Hungary and Poland, so we disregarded this variable.

Since the exchange rate enters the cointegrating vector with a negative sign, in the exchange rate adjustment equation a positive error correction coefficient indicates adjustment towards the equilibrium position. We tried to include other variables in the exchange rate adjustment equation (such as change of the interest rate differential, change of stock prices, inflation), but these were not significant and had the wrong sign. As a consequence, this equation differs from the random walk only by the time-varying drift and the error correction term.

Table 5 shows estimated results. The Chow test is significant for the price equation in all four cases, even for the short sample of the Czech Republic. The exchange rate equation indicates structural break for Hungary and Poland. Therefore, taking the full sample into account, seven of the eight short run equations have significant Chow tests.

To gain an impression of the variation of coefficients, we estimated these equations for rolling samples of three years. Estimated parameters in their two times standard error band are shown in Figures 8, 9, 10 and 11. (Note that in the case of the Czech Republic only samples starting in Q2 1993 and Q2 1996 - Q3 1997 can be regarded as covering homogenous periods.)

With regard to the price equation, the intercept shows a declining path in all countries, although it is more or less stable in Hungary for samples starting up to 1995. The declining path can be interpreted as diminishing autonomous inflation. The coefficient of the exchange rate is relatively stable in the case of the Czech Republic in samples starting after 1995 and varies substantially for the other three countries. The coefficient of the relative price of administered goods is insignificant in the case of the Czech Republic and Slovenia but has a steady path to a positive region, so we retain this variable in the final specification of these countries. Finally, the coefficient of the error correction term is quite sensitive to the sample in the case of Hungary, Poland and Slovenia, but more or less stable in the Czech case.

Parameters of the exchange rate equation reveal huge errors in all cases, which is a manifestation of the fact that modelling the nominal exchange rate is a complex problem. In the case of Hungary, Poland and to some extent Slovenia, the falling intercept indicates a declining autonomous source of depreciation. The error correction coefficients of the Czech Republic and Poland increases over time, which is reasonable since these countries made their exchange rate regime more flexible.

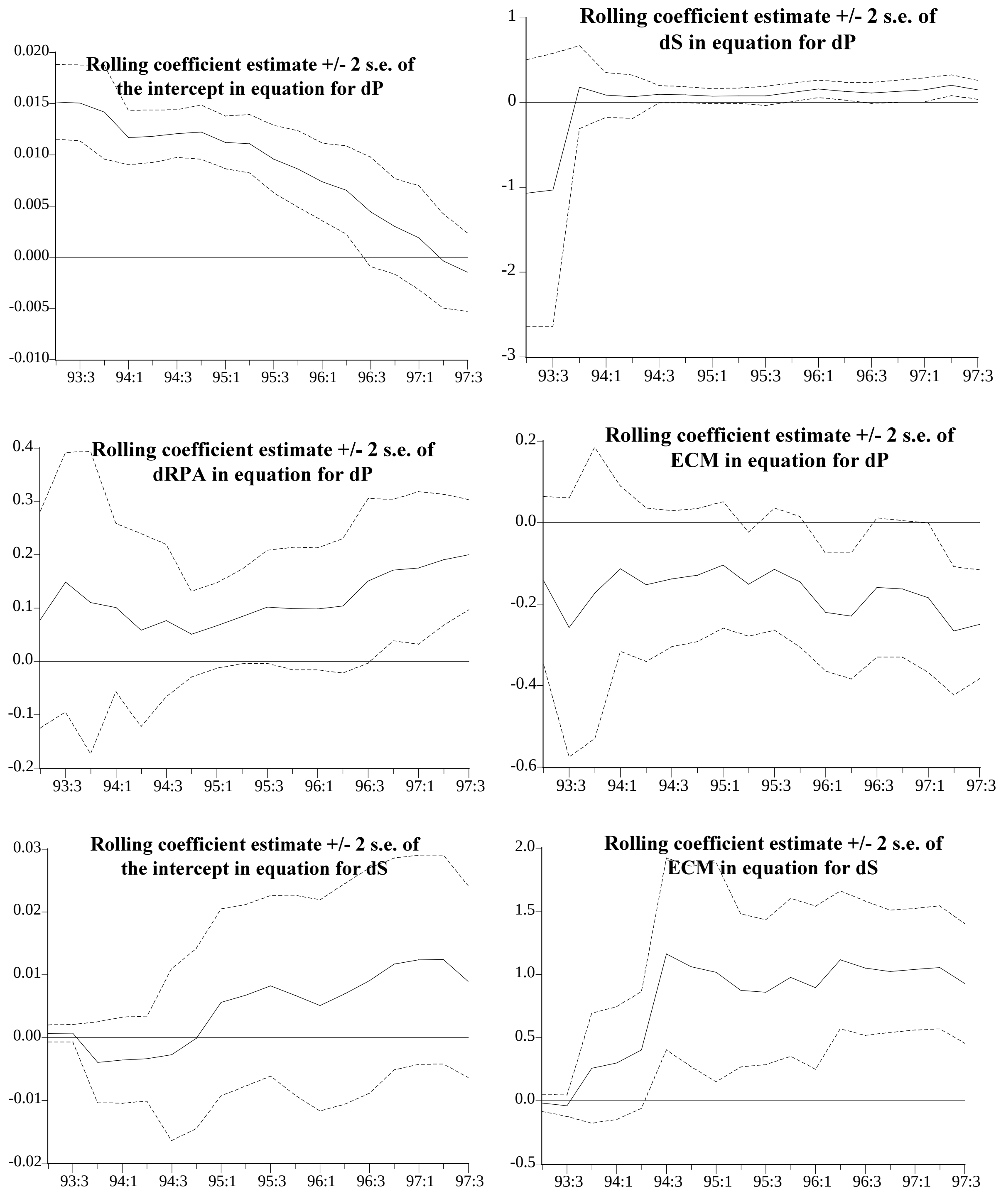
To sum up, some of the coefficients do seem to show substantial variability, necessitating time-varying estimation.

Table 5. OLS estimation of the pass-through model

	Price dynamics				*Exchange rate dynamics*			
	Czech Republic	Hungary	Poland	Slovenia	Czech Republic	Hungary	Poland	Slovenia
Const.	-0.001	0.020	0.031	0.014	0.009	0.034	0.029	0.018
t	-0.57	4.18	7.82	6.67	1.15	5.79	3.65	3.24
Δs	0.152	0.387	0.230	0.343				
t	3.47	4.20	2.78	5.31				
Δrpa	0.200			0.008				
t	5.70			0.079				
ecm	-0.250	-0.343	-0.244	-0.342	0.938	0.281	0.323	0.474
t	-2.69	-2.72	-3.37	-4.76	5.44	1.20	1.49	2.99
First obs.	Q3 1997	Q2 1993	Q2 1993	Q2 1993	Q3 1997	Q2 1993	Q2 1993	Q21993
Last obs.	Q2 2000	Q2 2000	Q2 2000	Q2 2000	Q2 2000	Q2 2000	Q2 2000	Q2 2000
T	12	29	29	29	12	29	29	29
R^2	0.783	0.532	0.406	0.675	0.606	0.069	0.127	0.197
Adjusted R^2	0.702	0.496	0.360	0.636	0.566	0.035	0.095	0.167
DW	1.47	1.17	0.71	1.81	1.58	1.15	1.93	0.91
SE	0.005	0.010	0.016	0.008	0.026	0.021	0.036	0.022
Chow	4.97*	5.62***	20.91***	4.80***	0.14	6.14**	3.22*	2.26

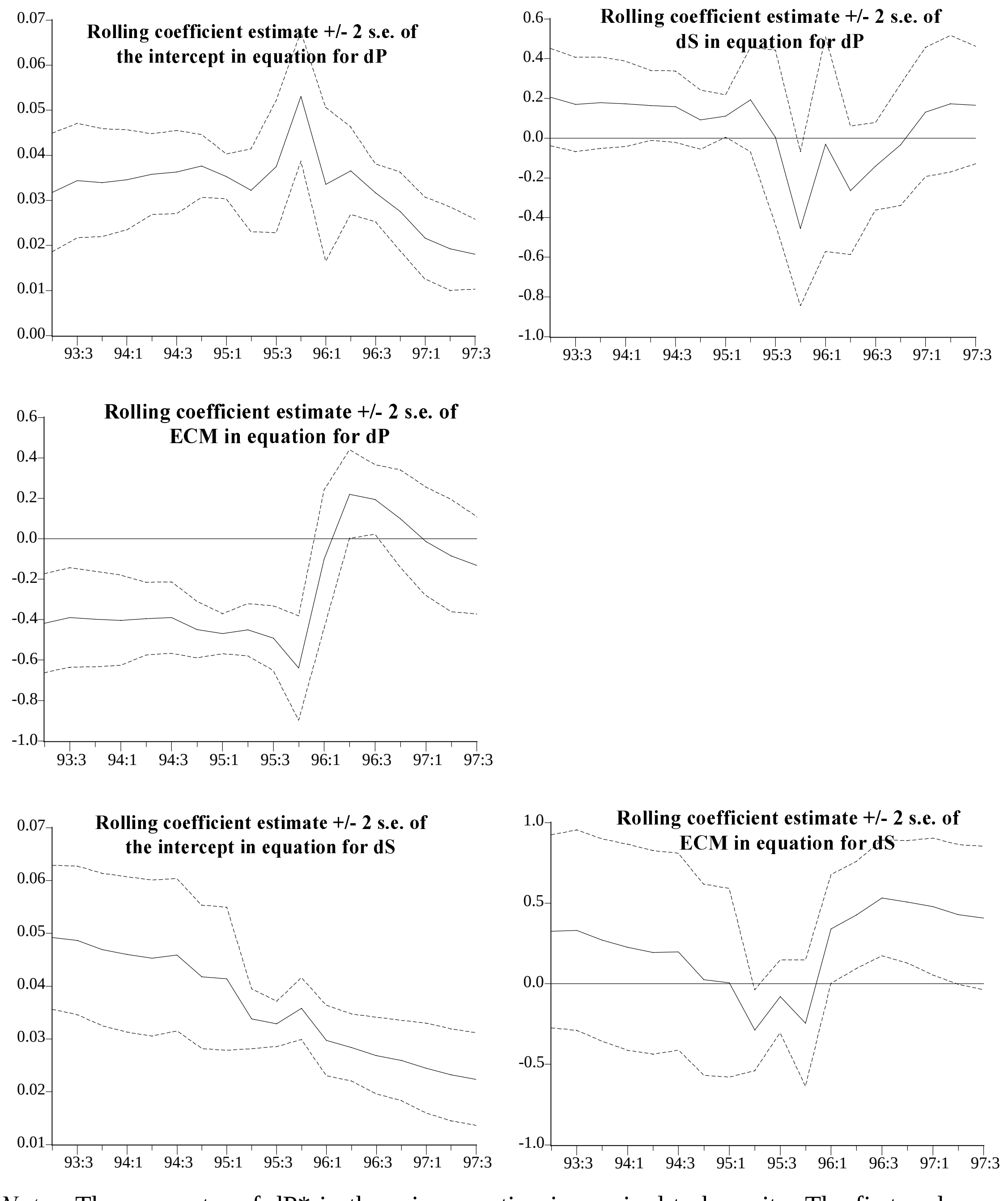
Notes: *t*: *t*-ratio calculated with Newey–West heteroskedasticity and autocorrelation consistent standard error; *First obs.*: first observation; *Last obs.*: last observation; *T*: number of observations; *DW*: Durbin-Watson; *AIC*: Akaike's information criterion, *BIC*: Schwarz information criterion, *SE*: standard error of regression, *Chow*: F–value of the Chow breakpoint test for Q3 1996 (Hungary, Poland, Slovenia) or 1998Q4 (Czech Republic); ***, **, or *: significant at 1%, 5%, or 10%.

Figure 8 (a-f). Czech Republic: Rolling sample single equation coefficient estimates



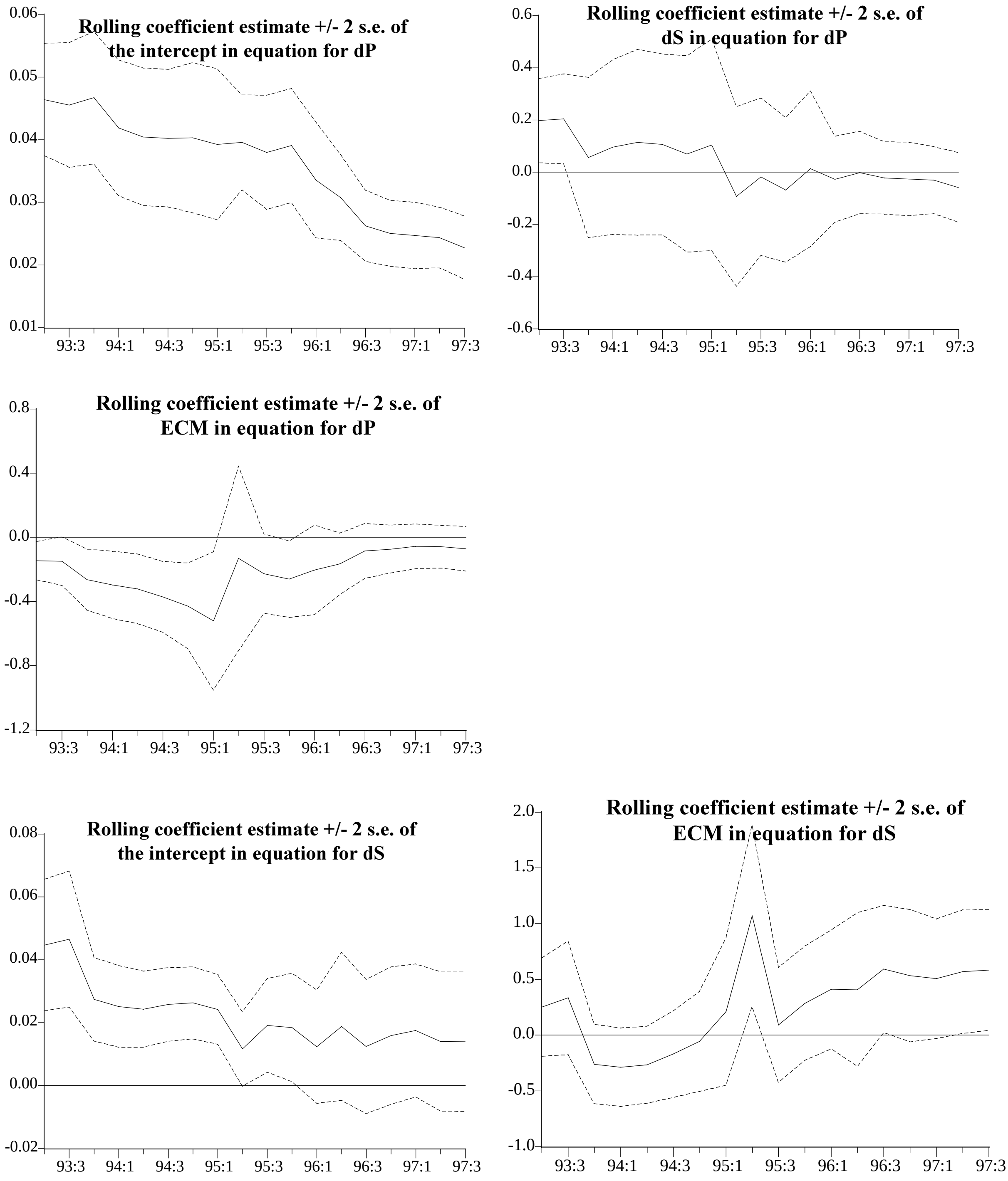
Notes: The parameter of dP* in the price equation is required to be unity. The first and second rows show parameter estimates for the price equation, and the third row for the exchange rate equation. The rolling window is three years long, e.g. the last observations show parameter estimates for the sample Q3 1997 – Q2 2000.

Figure 9 (a-e). Hungary: Rolling sample single equation coefficient estimates



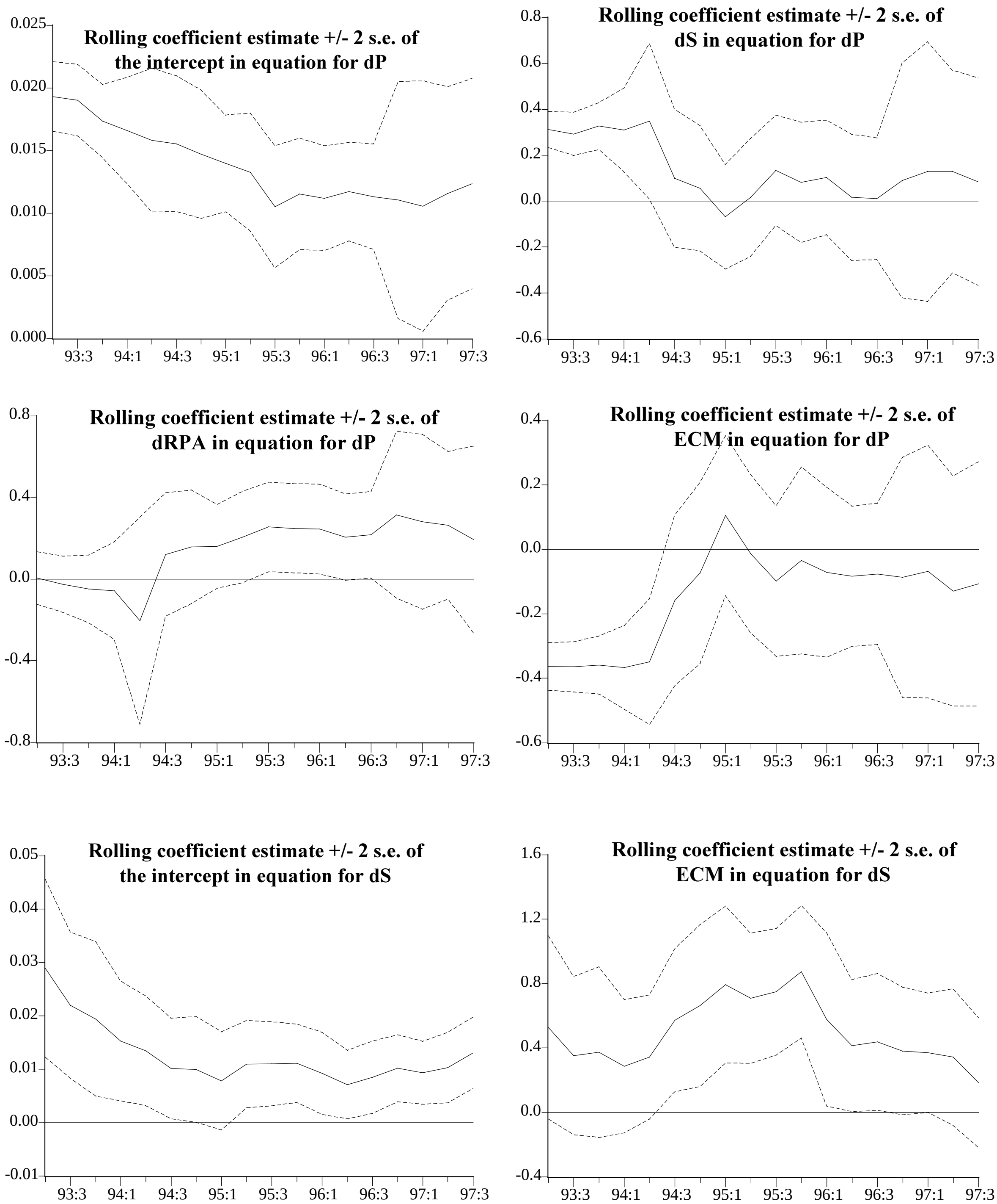
Notes: The parameter of dP* in the price equation is required to be unity. The first and second rows show parameter estimates for the price equation, and the third row for the exchange rate equation. The rolling window is three years long, e.g. the last observations show parameter estimates for the sample Q3 1997 – Q2 2000.

Figure 10 (a-e). Poland: Rolling sample single equation coefficient estimates



Notes: The parameter of dP* in the price equation is required to be unity. The first and second rows show parameter estimates for the price equation, and the third row for the exchange rate equation. The rolling window is three years long, e.g. the last observations show parameter estimates for the sample Q3 1997 – Q2 2000.

Figure 11 (a-f). Slovenia: Rolling sample single equation coefficient estimates



Notes: The parameter of dP* in the price equation is required to be unity. The first and second rows show parameter estimates for the price equation, and the third row for the exchange rate equation. The rolling window is three years long; e.g. the last observations show parameter estimates for the sample Q3 1997 – Q2 2000.

5.5 Time-varying parameter pass-through estimates

A time-varying estimation was performed for the model selected in the previous subsection. The estimation proved to be robust when various starting values were selected. Estimated time-varying parameters are shown in Figures 12, 13, 14 and 15 and some diagnostics in Table 6.

In accordance with expectations, in the case of the Czech Republic the parameter of the exchange rate in the price equation and the parameters of the exchange rate equations are virtually zero until Q1 1996 and take the right direction thereafter. The instantaneous pass-through is estimated to be around 0.1 and remains remarkably steady after Q1 1996, although the confidence band is wide.

In the case of Hungary, the autonomous source of inflation shows great variation (Figure 13). Its time path can be clearly interpreted: following the stabilisation programme in 1995, inflation increased and then declined gradually. Both the short-run impact of exchange rate changes on prices and the error correction coefficient are approximately twice as great as in the Czech Republic and have shown a slowly decreasing path (in absolute terms) in recent years. It is also interesting to note that depreciation during the Russian crisis (Q3 1998) was attributed to the intercept, i.e. to the autonomous part of the exchange rate depreciation, and for that date the model implies negative pass-through. In economic terms one can not find arguments suggesting that exactly during the Russian crisis exchange rate depreciation lowered inflation, so in our interpretation this negative coefficient is erroneous.

The autonomous source of inflation and exchange rate depreciation was also diminished in the case of Poland. The instantaneous pass-through also declined, especially after 1998, and does not differ much from zero. Looking at Figure 3 showing Polish exchange rate movements, we can observe that the pass-through coefficient arrived at the zero level when exchange rate fluctuations increased substantially.

In the case of Slovenia, the autonomous source of inflation shows a strong decline until 1997 but remains fairly stable thereafter (Figure 15). The magnitude of short-run pass-through and the error correction coefficients are about the same as in Hungary.

The long-run pass-through can be calculated by simulating the system under the assumption that the exogenous variables are unchanged. Figure 16 shows the dynamics of both variables in response to a 1% shock to the exchange rate using parameter values of Q1 1995 and Q2

2000.[28] In the Czech case, the values for Q1 1995 are only hypothetical: since the exchange rate did not change up to this period, it is pointless to ask what its long-run effects were on prices. Nonetheless, the answer given is in line with our expectations: prices must have adjusted fully, since this was the only way in this period to eliminate real exchange rate misalignment. The graph for Q2 2000 shows rapid convergence to the long-run value owing to the swift response of the exchange rate to misalignment. In the case of Hungary, Poland and Slovenia, the difference between short-run and long-run response is greater. Figure 17 shows instantaneous and long-run responses estimated at each date in the sample.[29] In the Czech case, there are two outlier long-run values in Q1 1996 and Q2 1996 which cannot be explained in economic terms. However, it is remarkable that the estimation of the long-run pass-through is relatively stable in the region of 0.15-0.2. In the case of Hungary and Slovenia, the long-run pass-through is estimated to be around 0.40-0.45; Poland is halfway between the Czech Republic and the other two countries with values around 0.25-0.3 at the end of the sample.

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[28] In the case of Hungary, we show the values corresponding to Q1 1994 because the period around Q1 1995 showed great instability (the stabilization program was introduced in this quarter).

[29] The long-run responses refer to 250 years, although there is much faster convergence.

Table 6. Diagnostics of time-varying parameter estimation

	Czech Republic	Hungary	Poland	Slovenia
T	29	29	29	29
k	8	7	7	8
LL	178.07	176.68	150.20	176.29
$\hat{\sigma}\left(\varepsilon_t^{(\Delta p_t)}\right)$	0.000 (s.e.: 0.003)	0.000 (s.e.: 0.004)	0.006 (s.e.: 0.001)	0.006 (s.e.: 0.001)
$\hat{\sigma}\left(\varepsilon_t^{(\Delta s_t)}\right)$	0.020 (s.e.: 0.003)	0.013 (s.e.: 0.003)	0.0231 (s.e.: 0.005)	0.013 (s.e.: 0.004)
$JB\left(\Delta p_t - \Delta\hat{p}_{t\|t-1}\right)$	1.84 (prob.: 0.397)	7.44 (prob.: 0.024)	0.94 (prob.: 0.62)	12.11 (prob: 0.002)
$JB\left(\Delta s_t - \Delta\hat{s}_{t\|t-1}\right)$	5.03 (prob.: 0.081)	7.11 (prob.: 0.029)	0.51 (prob.: 0.78)	0.14 (prob: 0.934)
One period ahead forecasts				
$MSE\left(\Delta p_t - \Delta\hat{p}_{t\|t-1}\right)$	0.384 (49.4 %RW)	0.873 (43.2 %RW)	0.982 (22.9 %RW)	0.442 (29.4 %RW)
$MAE\left(\Delta p_t - \Delta\hat{p}_{t\|t-1}\right)$	4.652 (65.4 %RW)	6.524 (54.7 %RW)	7.599 (44.6 %RW)	4.944 (48.9 %RW)
$MSE\left(\Delta s_t - \Delta\hat{s}_{t\|t-1}\right)$	6.020 (68.4 %RW)	3.453 (77.0 %RW)	11.685 (85.1 %RW)	3.684 (66.8 %RW)
$MAE\left(\Delta s_t - \Delta\hat{s}_{t\|t-1}\right)$	15.193 (86.8 %RW)	13.650 (95.4 %RW)	26.490 (97.2 %RW)	14.80 (91.2 %RW)

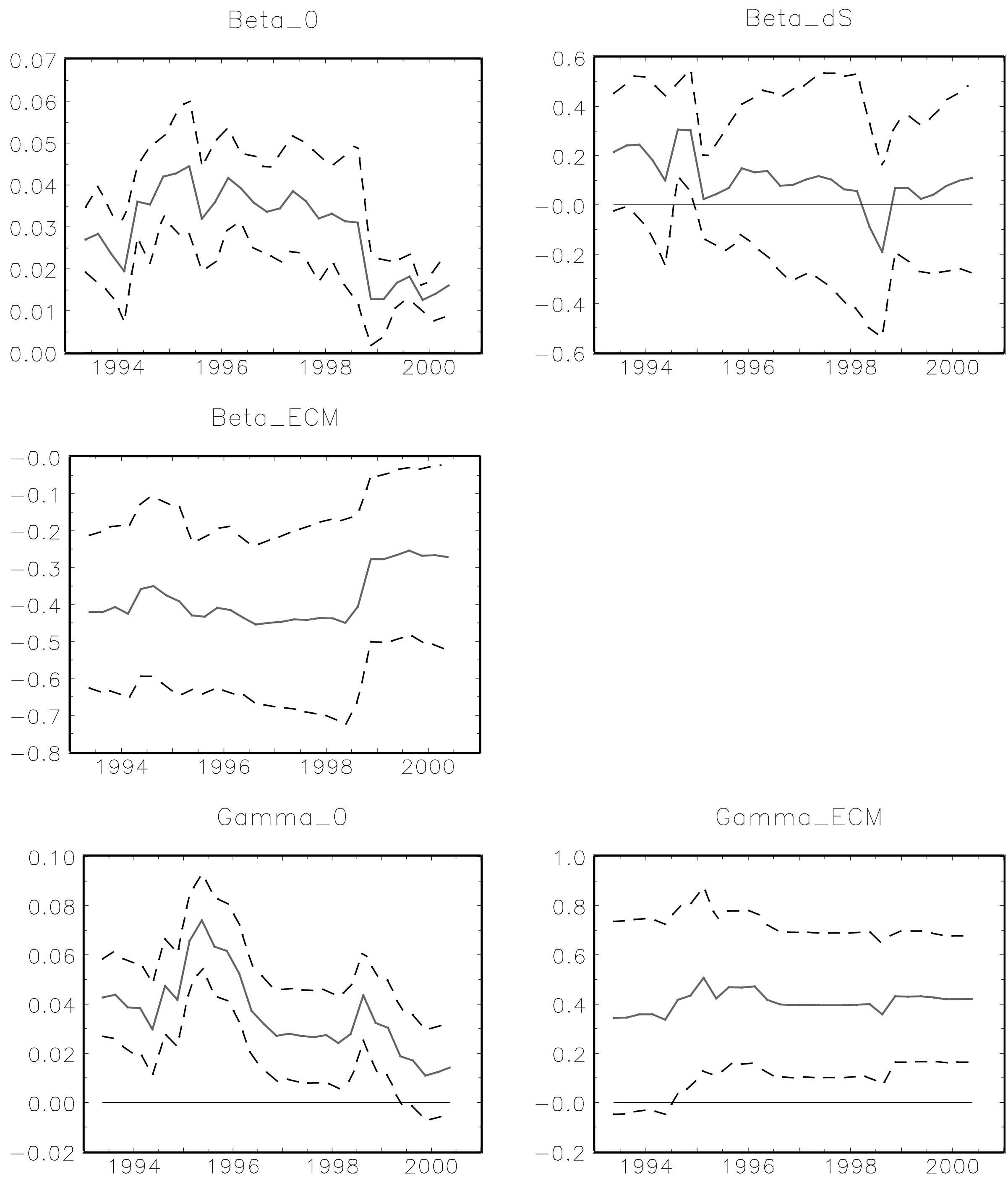
Notes: *k*: number of estimated parameters; *LL*: maximised value of the log likelihood function; *JB*: Jarque-Bera test for normality; *MSE*: mean squared error multiplied by 10,000; *MAE*: mean absolute error multiplied by 1,000; *%RW*: percentage of random walk with drift.

In evaluating the model, we compared its one period ahead forecasts to a random walk with drift. The last four rows of Table 6 show mean squared forecast error and mean absolute forecast error of both equations and their ratio to the values of the random walk. The improvement in forecasts is substantial in the case of the price equation, but also outperforms the random walk in the exchange rate equation.

Figure 12 (a-f). Czech Republic: Time-varying coefficient estimates

Note: The first two rows show time-varying parameter estimates of the price equation and the third those of the exchange rate equation.

Figure 13 (a-e). Hungary: Time-varying coefficient estimates

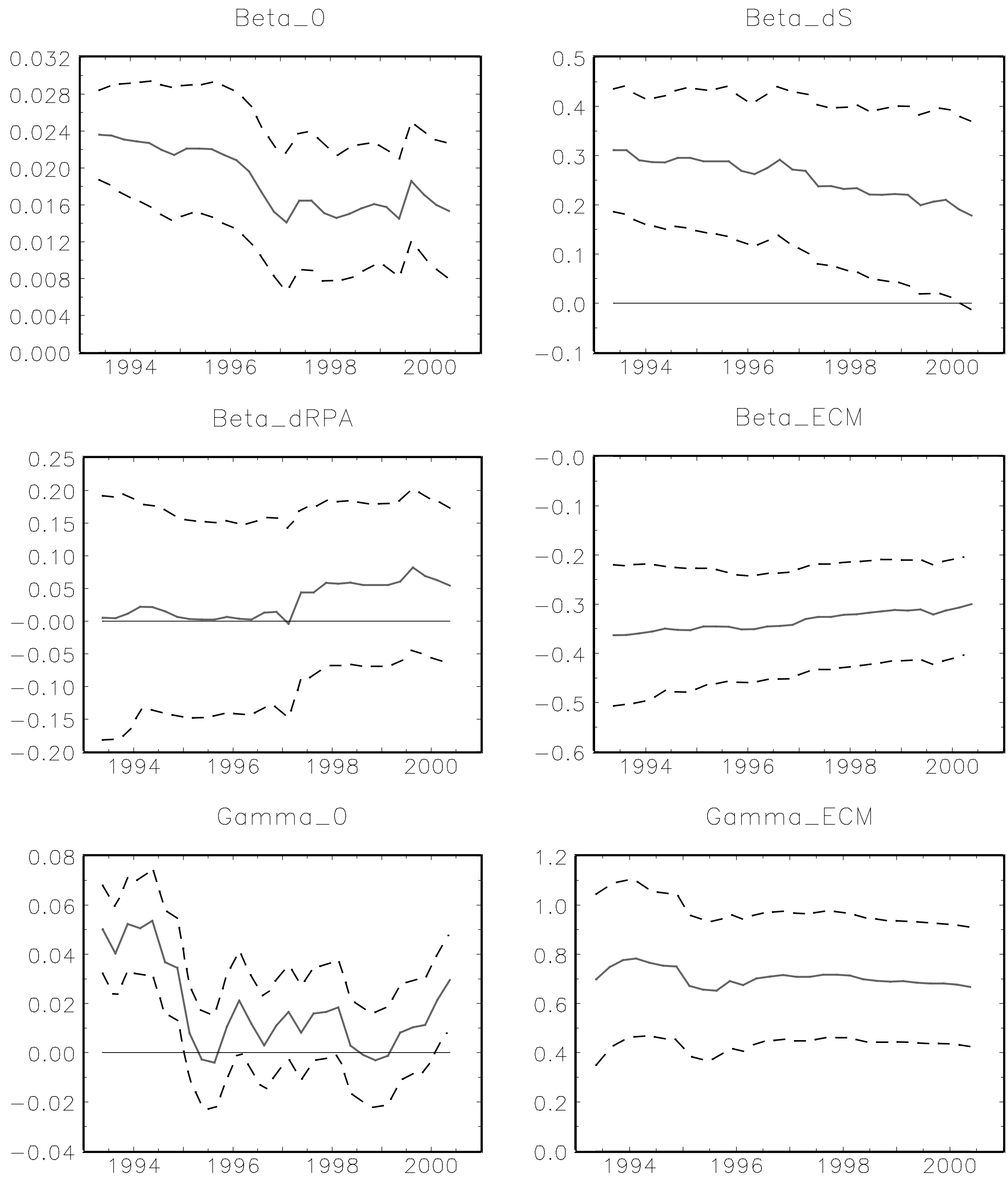


Note: The first two rows show time-varying parameter estimates of the price equation and the third those of the exchange rate equation.

Figure 14 (a-e). Poland: Time-varying coefficient estimates

Note: The first two rows show time-varying parameter estimates of the price equation and the third those of the exchange rate equation.

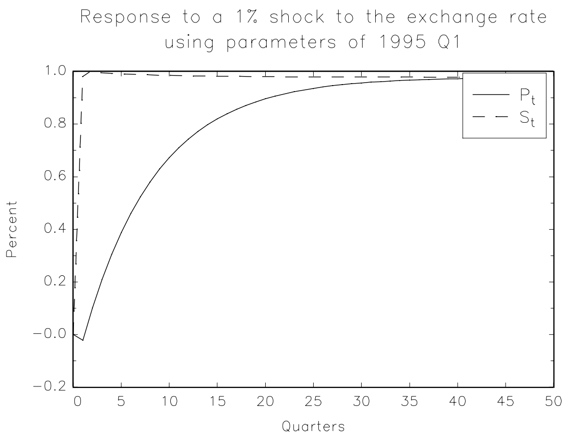
Figure 15 (a-f). Slovenia: Time-varying coefficient estimates

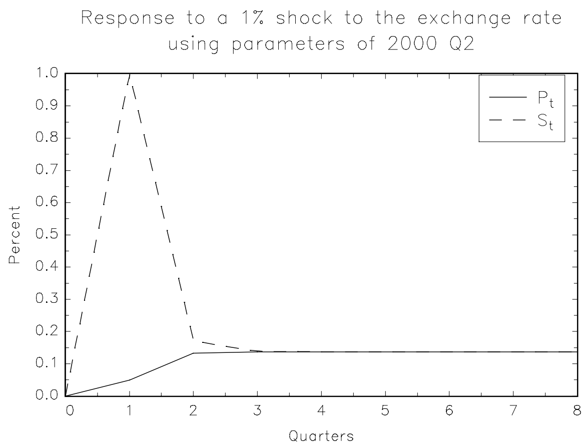


Note: The first two rows show time-varying parameter estimates of the price equation and the third those of the exchange rate equation.

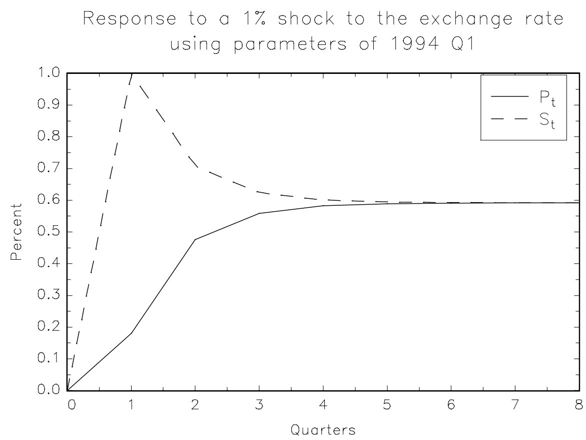
Figure 16 (a-h). Dynamics of the response to a 1% shock to the exchange rate on certain dates

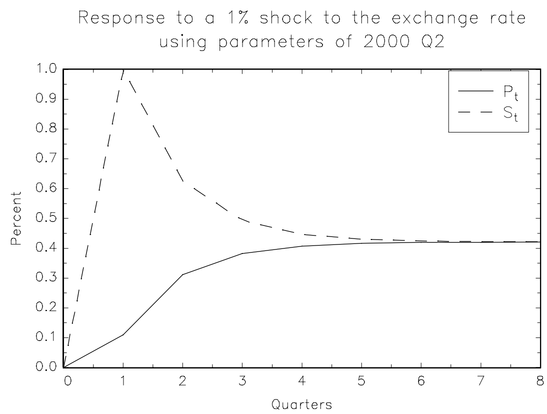
Czech Republic



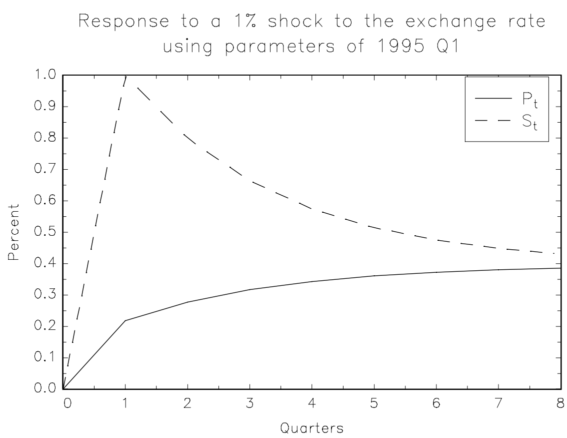


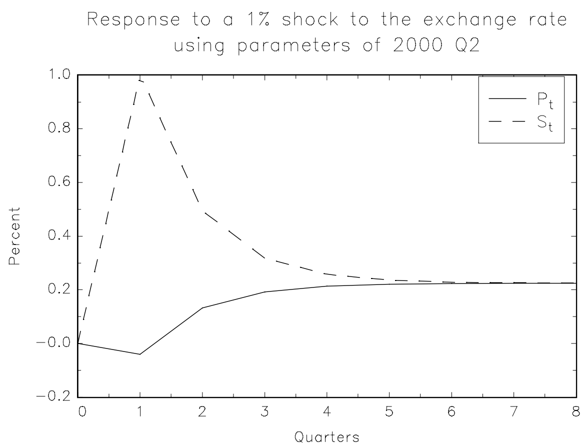
Hungary





Poland





Slovenia

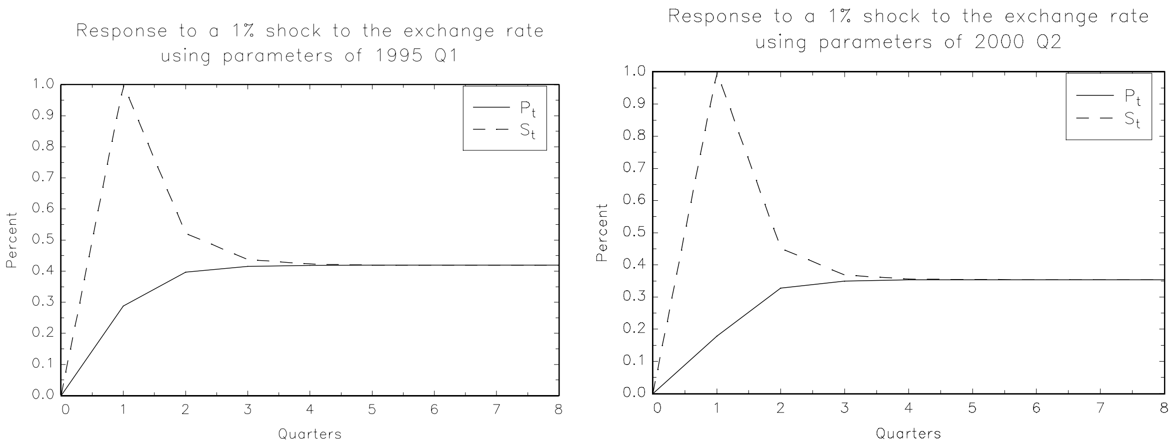
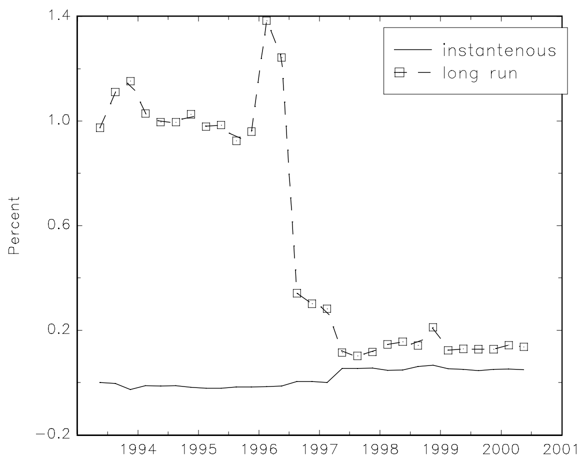


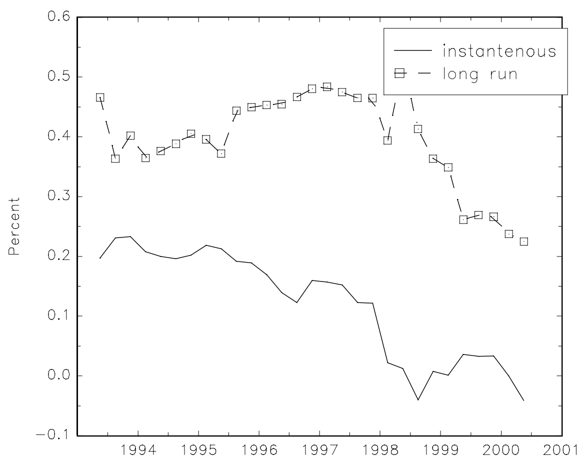
Figure 17 (a-d). Time-varying estimate of instantaneous and long-run response of prices to a 1% shock to the exchange rate

Czech Republic

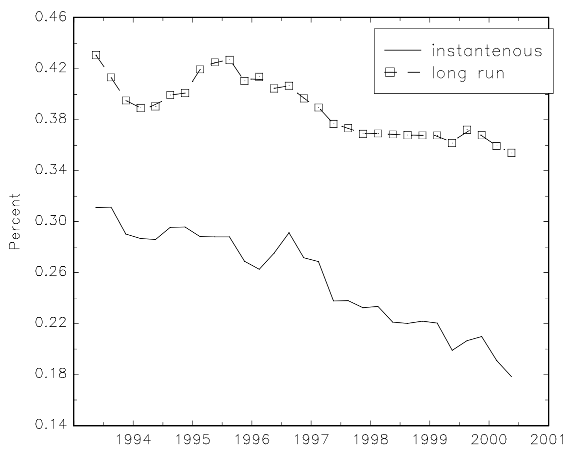


Hungary

Poland



Slovenia



6 Implications for exchange rate policy

At the last observation of the sample, Q2 2000, instantaneous (long-run) pass-through was roughly 10% (15%) in the Czech Republic and zero (20%) in Poland, which countries have floating exchange rate regimes, and approximately 10% (40%) in Hungary and 20% (40%) in Slovenia, which have managed exchange rate regimes. The response of prices to real exchange rate misalignment was also about twice as great in Hungary and in Slovenia as in the Czech Republic and Poland. There are some competing explanations for these results.

First, the exchange rate regime might effect pass-through. In an exchange rate targeting environment a change in the exchange rate might be regarded as more permanent than in a floating regime, implying higher pass-through.

Second, it is possible that the exchange rate regime itself is irrelevant, whereas the volatility of the exchange rate matters. In this case, even in a floating regime, if volatility were lower, then pass-through would be higher. Figure 18 confirms a negative relationship between exchange rate variability and pass-through.[30]

Third, inflation in Hungary, Poland and Slovenia was higher than in the Czech Republic and, with the exception of 1998-2000 in the case of Poland, pass-through was also higher. There are arguments suggesting that lower inflation leads to lower pass-through and it might also be argued that a higher level of inflation is associated with a higher price response to real exchange rate misalignment. As can be seen from Figure 19, inflation and pass-through are positively related, which relationship is clear in the cases of Hungary, Poland, and Slovenia.

The available evidence is insufficient to allow discrimination among the aforementioned hypotheses. However, one conclusion can be drawn for exchange rate policy. If the authorities can credibly reduce exchange rate volatility, then exchange rate targeting might play a positive role in reducing inflation. Credibility is crucial as the reasonable current account balance properly supported both Hungarian and Slovenian regimes.

The conclusion is neither new nor surprising. However, it is important as current thinking among many international economists tend to disregard the possible role of the exchange rate in curbing inflation in small open economies.

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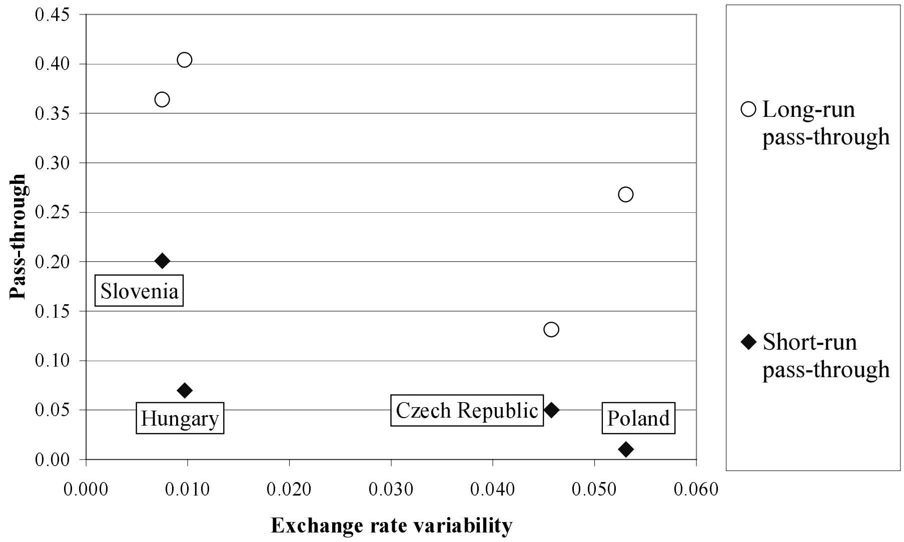
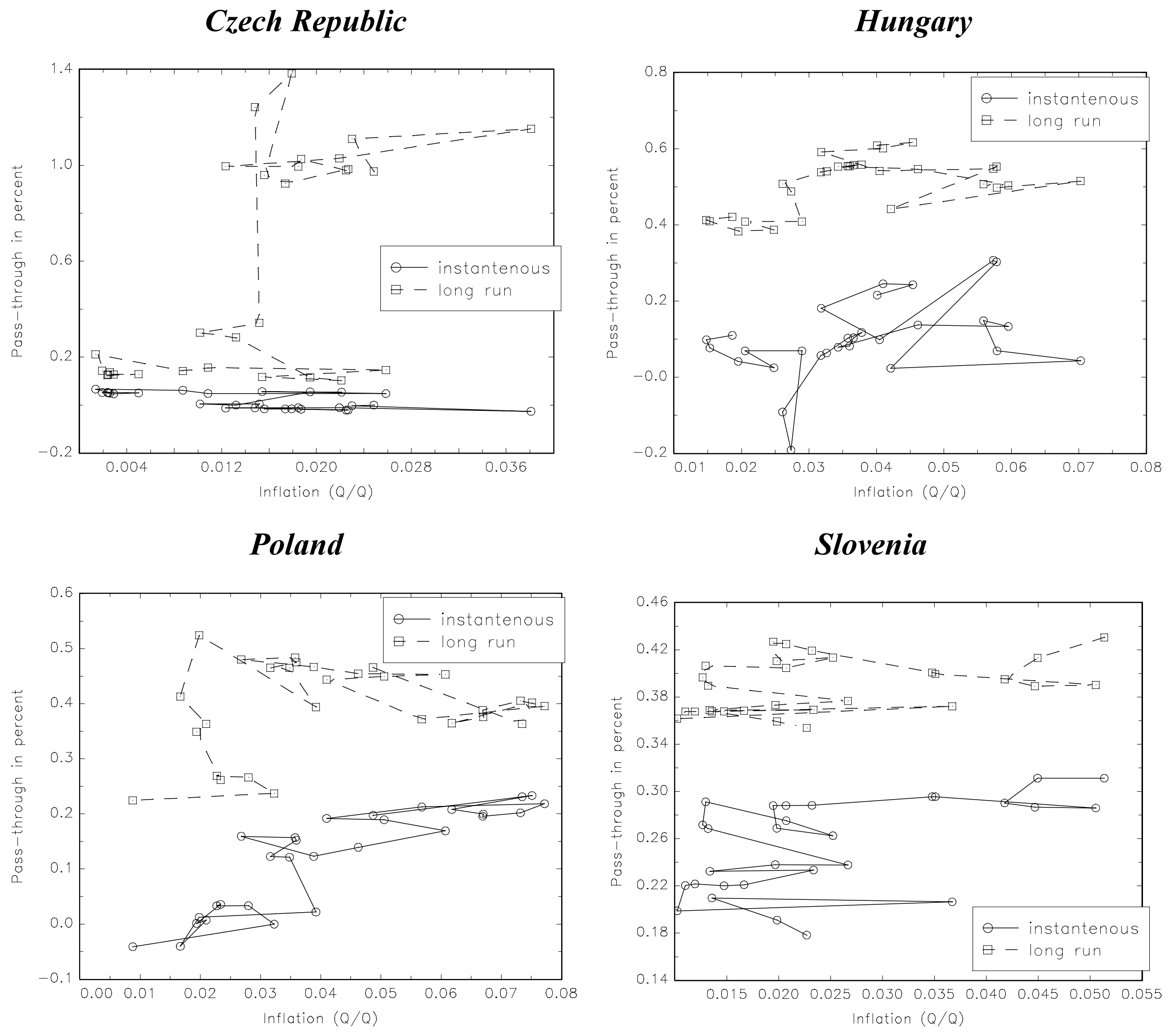
[30] Data shown refers for the period average in Q1 1999 – Q2 2000. As we have seen in the case of Hungary, the depreciation following the Russian crisis was inferred to have negative pass-through. We regarded this negative parameter as an erroneous inference, and the parameter indeed made a steady increasing path to a positive region reaching 10 percent by Q2 2000. As a consequence, the period average shown in Figure 18 biases downward Hungarian short-run pass-through.

Figure 18. Cross-plot of exchange rate variability and pass-through



Note: Pass-through: average of the period Q1 1999 – Q2 2000; Exchange rate variability: standard deviation of quarterly changes of the exchange rate in the period Q1 1999 – Q2 2000.

Figure 19 (a-d). Cross-plot of inflation and pass-through



7 Summary

In this paper we studied exchange rate pass-through into consumer prices of four EU candidate countries: the Czech Republic, Hungary, Poland and Slovenia. Our measure of consumer prices does not include administered prices and prices with seasonal and highly volatile patterns (food and energy prices). The model of this paper has three distinct features: (1) joint analysis of price and exchange rate changes, (2) error correction formulation incorporating the equilibrium real exchange rate, (3) time-varying parameters in the short-run equations.

The price levels in these countries are far lower than the EU average and their economies are growing faster, indicating that the equilibrium real exchange rates are on an appreciating path. The error correction formulation allows us to distinguish the sources of price changes between exchange rate shocks and equilibrating real appreciation. Joint modelling of inflation and exchange rate changes permits us to study the response of the exchange rate to both real exchange rate misalignment and to its own previous shocks. In addition, the error correction formulation with joint modelling of variables provides a convenient description of the dynamics of pass-through in contrast to many applications which adopt autoregressive distributed lag models with ad hoc restrictions. There are several reasons to expect time variation in coefficients: behavioural changes, aggregation, changing inflationary environment, exchange rate expectations and non-linear real exchange rate mean reversion.

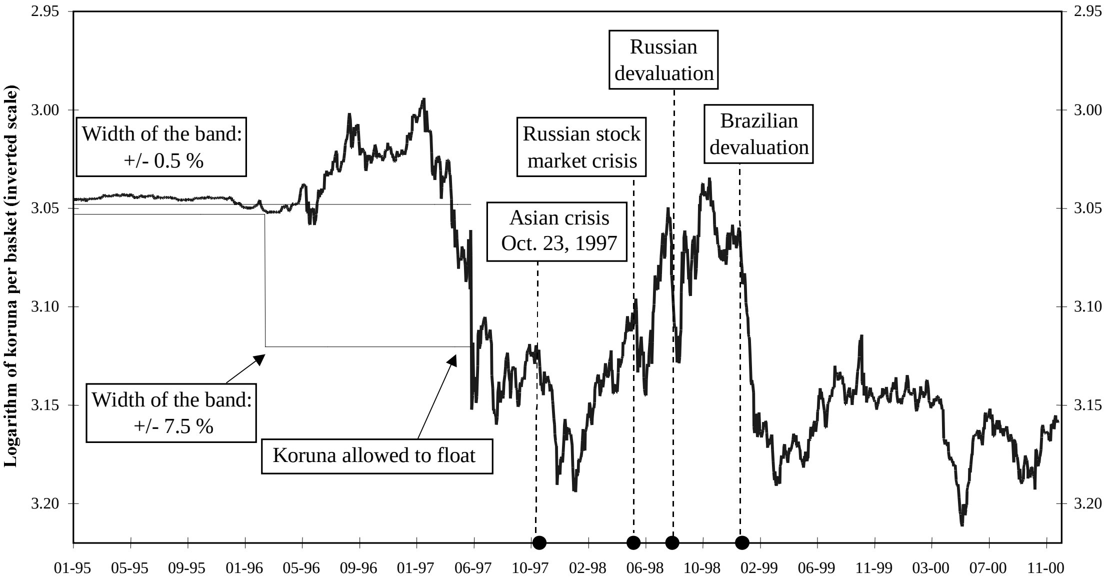
The model is supported by empirical analysis. Fixed coefficient estimates proved to be unstable and full of significant tests for structural breaks. Time-varying coefficient estimates led to correct coefficient signs and, in most cases, to significant estimates. The magnitudes of the time-varying coefficients seem reasonable in economic terms. The model beats the random walk in terms of one period ahead forecasting. Although this is no great achievement, the opposite result would have been devastating for the model.

At the last observation of the sample, Q2 2000, instantaneous (long-run) pass-through was roughly 10% (15%) in the Czech Republic and zero (20%) in Poland, which countries have floating exchange rate regimes, and approximately 10% (40%) in Hungary and 20% (40%) in Slovenia, which have managed exchange rate regimes. We were unable to determine whether country differences can be attributed to the exchange rate regime, to exchange rate volatility, or to the level of inflation. Nonetheless, it does seem that credible exchange rate management can play a useful role in reducing inflation.

8 Annex

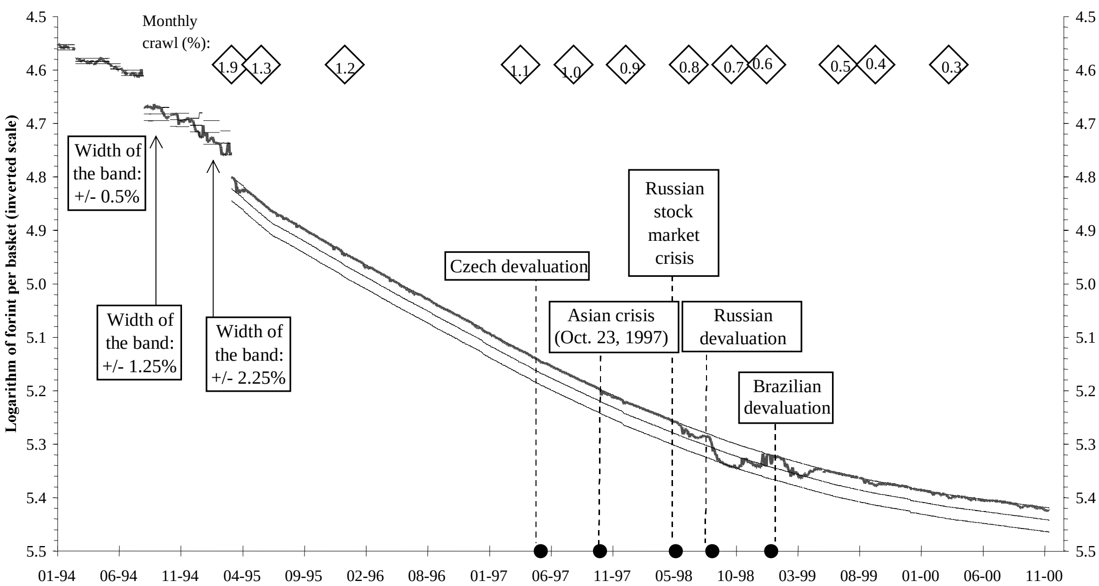
8.1 Nominal exchange rate movements

Figure 20. Czech Republic: Nominal exchange rate of the koruna against the official currency basket, January 1994 – December 8, 2000



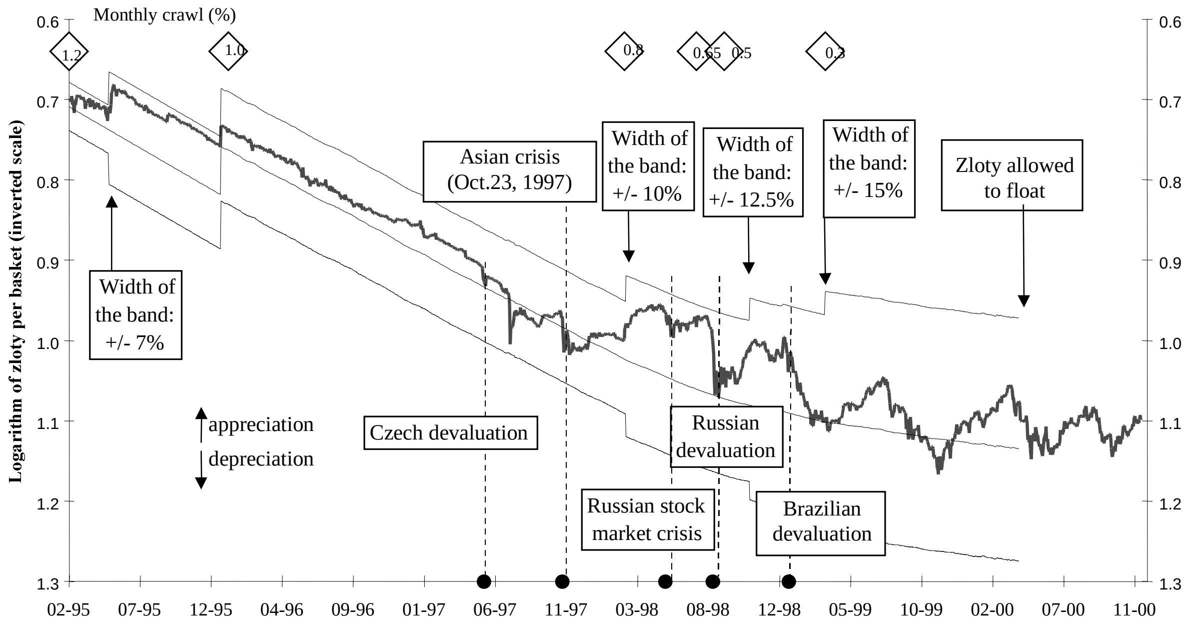
Source: Author's calculation based on data from the Czech National Bank and Reuters.
Notes: Basket used prior to May 1997: 65% DEM - 35% USD. We used the same basket for the floating period for comparative purposes.

Figure 21 .Hungary: Nominal exchange rate of the forint against the official currency basket, January 1994 – December 8, 2000



Source: Author's calculation based on data from the National Bank of Hungary.
Notes: Composition of the basket: 50% DEM + 50% USD for August 1993 – May 1994; 70% ECU + 30% USD for May 1994 – December 1996; DEM + 30% USD for January 1997 – December 1998; 70% EUR + 30% USD for January – December 1999; 100% EUR since 2000.

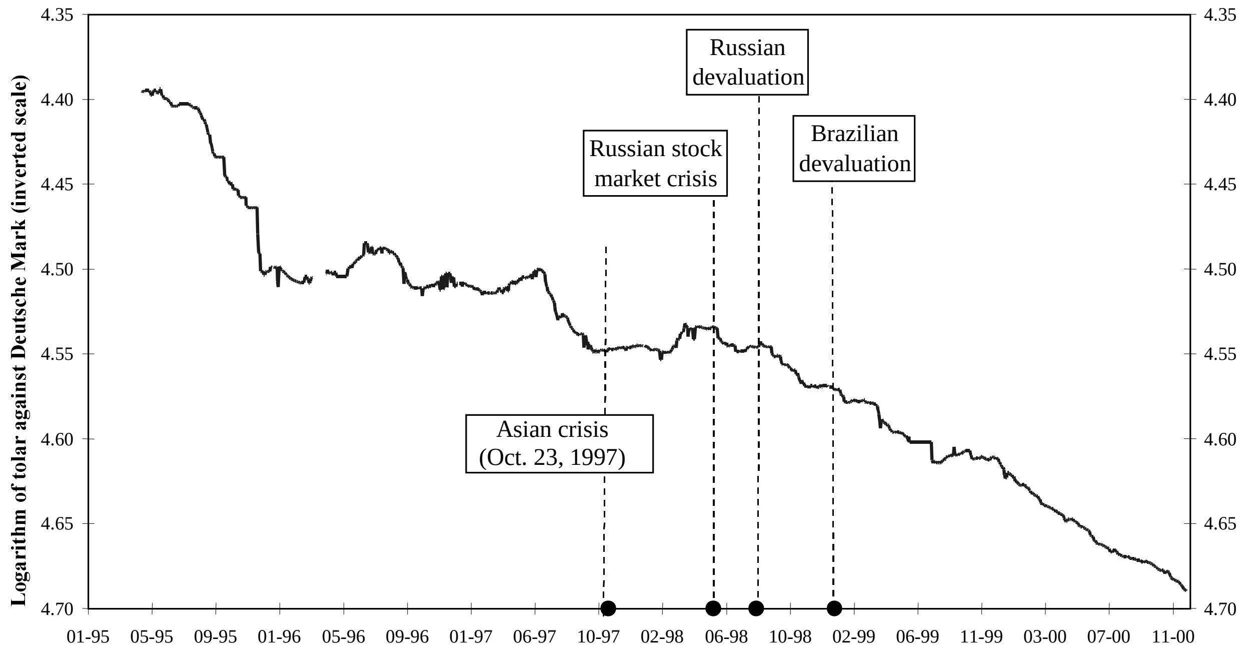
Figure 22. Poland: Nominal exchange rate of the zloty against the official currency basket, January 1994 – December 8, 2000



Source: Author's calculation based on data from Reuters.

Notes : Composition of the basket prior to 1999: 45% USD, 35% DEM, 10% GBP, 5% FRF, 5% CHF; since 1999: 55% EUR , 45% USD.

Figure 23. Slovenia: Nominal exchange rate of the tolar against the Deutsche Mark, April 1995 – December 8, 2000



Source: Author's calculation based on data from Reuters.

8.2 *Measurement of the equilibrium real exchange rate*

8.2.1 Concepts and methods

The concept of the equilibrium exchange rate is not without controversy. As shown in the survey carried out by Isard–Faruqee (1998), there are even arguments questioning the usefulness of this concept. Assuming that the concept makes economic sense, the equilibrium can be defined as a state in which "any transitory fluctuations in exogenous variables have been identified and discarded, all policy variables have been set at their sustainable long-run values, and *all* predetermined variables have been allowed to complete their endogenous adjustments and reach their steady-state levels. Unfortunately, however, a "purist" definition of this type would have little analytical or operational content" (Montiel, 1999, p. 226).

Theoretical and empirical literature on equilibrium real exchange rates is voluminous, while applications to EU candidate countries are rare. Several methods are surveyed in MacDonald (2000) and in the volumes of Williamson, ed. (1994), Stein–Allen, eds. (1995), MacDonald–Stein, eds. (1999), and Hinkle–Montiel, eds. (1999), which are labelled with abbreviations such as PPP, CHEER, BEER, PEER, DEER, FEER, NATREX, LRER, SRER. The simple purchasing power parity (PPP) idea is not applicable in the case of candidate countries owing to heavy structural changes. Many of the other methods (CHEER, BEER, PEER) are time series-based analyses adopting the Johansen VECM procedure and structural vector autoregressions with a small number of variables, or other univariate and multivariate trend-cycle decompositions. The atheoretical decompositions are heavily criticised (MacDonald). The applicability of VAR-based approaches to candidate countries are constrained by the short time span of data, the assumption of constant behavioural relationships and the uncovered interest rate parity (*UIP*) hypothesis, which is assumed in most cases and clearly violated in these countries.[31]

The fundamental equilibrium exchange rate (FEER) approach advocated by Williamson (1983) indicates an exchange rate that represents both internal and external balance. Internal balance is usually defined as high employment and low inflation, and the external balance as a sustainable current account balance and net foreign assets position. Empirical implementation of this approach requires a judgement on the sustainable external balance and estimation of the current account elasticities to the real exchange rate.[32] Begg–Halpern–Wyplosz (1999) argue that features of these countries "eliminate any hopes of computing the determinants of a sustainable current account" (p. 4). Krajnyak–Zettelmeyer (1998) add that "it would seem

[31] Introducing a time-varying endogenous risk premium into the UIP relationship is a burdensome if not impossible exercise, especially for candidate countries.

[32] The IMF's concept of a desired equilibrium exchange rate (DEER) is very similar to FEER.

difficult, if not impossible, to estimate the effects of real exchange rate movements on the current account at a time when fluctuations in exports are likely to be driven primarily by such changes as removal of export quotas, the break-down of traditional trading blocks, and changes in relative prices within the tradable sector" (p. 314).

Due to the deficiencies of the methods mentioned above, Halpern–Wyplosz (1997), Krajnyak–Zettelmeyer (1998) and Begg–Halpern–Wyplosz (1999) perform panel estimates for nominal monthly wages in USD terms for about eighty countries in the world in order to shed light on the equilibrium nominal dollar wage movements of transition countries. The explanatory variables are measures of development, such as PPP-adjusted GDP per worker or per capita, schooling, the share of agriculture, and some other variables and dummies. They calculate the equilibrium dollar wages by fitting the model for transition countries. Both papers document the original hypothesis of Halpern–Wyplosz that a considerable part of real appreciation was a correction of an initial strong undervaluation.

The panel approach is very similar to what is known as the reduced form approach to the exchange rate, which is also adopted in this paper. In this method the real exchange rate of a given country is regressed on some exogenous fundamentals, which equation is frequently claimed to be a reduced form implied by a theoretical model.[33]

The basic building block of this approach is the following model. There are two types of goods: non-tradable goods (*NT*) and tradable goods (*T*). We assume that the same good sold at home and at foreign markets does not necessarily have the same prices in common currency. The consumer price index is the weighted average of goods prices (all variables are in logarithms):

$$
(5) \qquad \begin{aligned} p_t &= \phi_{NT}\, p_t^{NT} + (1-\phi_{NT})\, p_t^{T} \\ p_t^* &= \phi_{NT*}^*\, p_t^{NT^*} + (1-\phi_{NT*}^*)\, p_t^{T^*} \end{aligned}
$$

where, of variables not previously defined, p_t^{NT} and p_t^{NT*} are the price level of non-tradable goods at home and abroad, p_t^{T} and p_t^{T*} are the price level of tradable goods at home and abroad, ϕ_X is the share of sector X in consumption, and an asterisk indicates foreign variable and weight.

We can decompose the real exchange rate as:

---

[33] See, for example, Feyzioglu (1997), Mongardini (1998), Alberola et al. (1999) and Baffes et al. (1999).

$$(6) \qquad q_t = p_t - p_t^* - s_t = \phi_{NT}\left\{\left(p_t^{NT} - p_t^T\right) - \left(p_t^{NT^*} - p_t^{T^*}\right)\right\} + \left[p_t^T(d) - p_t^T(f) - s_t\right] + w_t \quad ,$$

where, of the variables not defined previously, q_t is the actual real exchange rate and w_t is a discrepancy term reflecting the differences between the weights. We assume that w_t is a white noise.[34] The first element of the decomposition reflects the relative price of non-tradables to tradables at home versus abroad, which might change primarily due to the Balassa–Samuelson (BS) effect. We will denote this term as q_t^{BS}. The second term gives the real exchange rate of tradable goods, which might be different from zero because of either imperfect substitutability of domestically and foreign-produced tradable goods or due to pricing to market. We denote this rate as q_t^T. Using these notations we obtain

$$(7) \qquad q_t = \alpha_{NT} q_t^{BS} + q_t^T + w_t \quad .$$

The definition of the equilibrium real exchange rate is based on the condition of both internal and external balances for sustainable values of policy variables and exogenous variables. Internal balance is reached when the market for non-tradable goods clears, that is, when the following holds:

$$(8) \qquad y_t^{NT}\left(q_t^{BS},...\right) = c_t^{NT}\left(q_t^{BS},...\right) + g_t^{NT} \quad ,$$

where y_t^{NT} is the output of non-tradable goods under full employment, c_t^{NT} is private sector demand and g_t^{NT} is government consumption of non-tradable goods. q_t is at equilibrium when g_t^{NT} is set at its sustainable level and there is no excess demand for non-tradables.

The external equilibrium is usually defined as the achievement of the desired long-run level of net foreign assets. Change in net foreign assets, which equals the current account, can be written as:

$$(9) \qquad \Delta f_t = ca_t\left(q_t^T,...\right) = tb_t\left(q_t^T,...\right) + r_t^* f_t + gr_t = y_t^T\left(q_t^T,...\right) - c_t^T\left(q_t^T,...\right) - g_t^T + r_t^* f_t + gr_t \quad ,$$

where f_t is net foreign assets, ca_t is the current account, tb_t is the trade balance, r_t^* is the foreign real interest rate, gr_t is the foreign grants received by the government and the domestic private sector, y_t^T is the domestic output of tradable goods, c_t^T is the private sector demand and g_t^T is the government consumption of tradable goods.

34 Kovács–Simon (1998) looked at the effect of the weight discrepancy in the case of Hungary for changes in the real exchange rate. They have found that this discrepancy explains 0.2% per year movement on average, which is quite close to zero.

Equations (7-9) are standard building blocks of real exchange rate models. The left-hand side of (9) is set to zero and, with some further assumptions, the model can be solved for a reduced form in which the real exchange rate is the function of some exogenous and policy variables (see, for example, Alberola et al., 1999, Baffes et al., 1999). Having estimated the reduced form, the equilibrium real exchange rate is disclosed by fitting the model, using an estimated measure of the long-run components of the fundamentals.

As it is clearly discussed in Montiel (1999), selecting the long-run fundamentals poses severe difficulties. The most problematic variable is the net foreign assets (NFA) position, which is neither a policy nor an exogenous variable, and can be characterised as a predetermined variable. The key question is what kind of assumption we make for the adjustment speed of NFA. If the adjustment is relatively fast, then NFA is an endogenous variable in the system and there is no need to condition the estimate of the equilibrium real exchange rate on it. However, if the adjustment of the stock of NFA is judged to be too slow for policy relevance, then the natural procedure is to condition the equilibrium real exchange rate on the predetermined value of NFA. The former approach is adopted in Baffes et al. (1999) and the latter in Alberola et al. (1999).

For EU candidate countries the second approach seems more appropriate. Most of these countries accomplished their transition process during our sample period, which implies that their economies have undergone heavy structural changes. In a continuously changing economy the long-run or sustainable level of net foreign assets might have also changed. In addition, the starting position of some of the countries might have been far from equilibrium. For example, the communist regime in Romania had repaid almost all foreign debt by the end of the 1980s at the cost of pushing down living standards substantially, and Poland was able to achieve a significant debt reduction in the early 1990s. These arguments suggest that in our sample period of 1993-2000 the adjustment of NFA stocks toward long-run equilibrium might not have taken place.

There is another important consideration regarding the NFA positions of these countries that relates to its composition. While it is reasonable to assume that these countries earn the world interest rate on their foreign assets, their liabilities have some special characteristics. Inflows of capital can be divided into three major groups: (1) foreign direct investment, (2) interest rate sensitive flows into domestic currency denominated securities, and (3) official borrowing in foreign exchange.

(1) Foreign direct investment (FDI) plays a key role in "financing" the current account deficit in these countries (Table 1) and in building up competitive productive

capacities. We use the term "financing" in inverted commas as the order of causality is inverted: low wages compared to the quality of human capital and different government incentives attract foreign direct investment which is followed by imports of capital goods and raw materials. A significant part of the established productive capacities will produce exports. Therefore, FDI has various effects on the current account with different time lags: large imports of capital goods in the first stage, continuous imports of raw materials and exports of finished products starting in a second stage, and continuous profit repatriation in a third stage. Even profit repatriation effect cannot be captured by a simple term $r_t^* f_t$ in equation (9).

(2) Capital inflows into domestic currency denominated securities might be sensitive to the domestic interest rate. As uncovered interest rate parity does not hold in these countries, but there is a positive expected excess return that depends on various factors, the effective interest payments might be higher and not a linear function of $r_t^* f_t$.

(3) Most governments of these countries borrow in foreign currency to keep their international reserves at a desired level. The premium on these borrowings varies depending on the macroeconomic conditions and the willingness of foreign investors to take risks.

The above arguments suggest that for a large part of foreign liabilities the $r_t^* f_t$ income transfer poorly captures true flows and also highlight the fact that FDI has a distinctive role compared with other NFA positions. Therefore, we will condition the equilibrium real exchange rate separately on the stock of FDI liabilities and on the stock of other net foreign assets (ONFA). The reduced form of the hypothetical model leads to the following equation:

$$(10) \qquad q_t^{EQ} = f\left(g_t^T, g_t^{NT}, gr_t, r_t^*, FDI_t, ONFA_t, prod_t, tot_t, tp_t\right),$$

where, of variables not previously defined, $prod_t$ is a measure of productivity differentials related to the rest of the world, tot_t is terms of trade, and tp_t is the ratio of import duties to export subsidies representing the stance of trade policy.

In the reduced form of (10) data availability constrains us to use the share of government in GDP (g_t - y_t) instead of g_t^T and g_t^{NT} and to drop tp_t.

8.2.2 Some evidence for candidate countries

Analysing Hungarian data, Kovács–Simon (1998) found strong evidence in favour of the Balassa-Samuelson (BS) effect. Their results indicated that the two goods (tradable/non-tradable) model is a useful hypothesis only for differentiating between industries according to their rate of technological change, since the relative price of non-tradable goods reflect relative productivities between the two sectors. However, the tradable/non-tradable distinction does not separate good substitutes from poor substitutes for internationally traded goods, since prices of traded and non-tradable goods respond similarly to nominal exchange rate shocks. Canzoneri–Cumby–Diba (1999) reach the same conclusion for industrial countries. These results indicate that the two goods model might not be satisfactory for our purpose, but suggest clearly that the law of one price cannot be applied to tradable goods.

In addition to the Balassa–Samuelson effect, there might be a complementary explanation of the rise in the non-tradable/tradable price ratio, namely shifts in demand toward non-tradables. As argued by Halpern–Wyplosz (1997), the service sector was repressed in the era of planned economies and at least the early years of transition were characterised by large increases in the productivity of the non-tradable sector caused by improved distribution networks and the build-up of a banking system and other services that were previously non-existent. Repressed inherited non-tradable prices coupled with shifts in demand towards non-tradables might increase their relative price, even though there is a rapid productivity increase in this sector. However, since our sample starts in 1993, some three to four years after transition started in most countries, the initial build-up of services and their sharp relative price adjustments have already taken place to a large extent.

Table 7. Current account and foreign direct investment, 1990-99 (% of GDP)

		1990	1991	1992	1993	1994	1995	1996	1997	1998	1999
Bulgaria	*CA*		-2.6	-7.9	-22.1	-0.7	-0.4	0.3	7.5	-1.1	
	FDI		1.9	0.9	0.8	2.4	1.5	1.9	8.9	9.5	
Croatia	*CA*										
	FDI										
Cyprus	*CA*	-4,6	-11.1	-14.8	2.8	1.7	-3.0	-8.4	-5.9	-9.7	
	FDI	3.8	2.2	2.5	2.1	1.8	1.5	0.9	1.2	0.6	
Czech Rep.	*CA*				1.4	-2.1	-2.6	-7.4	-6.2	-2.5	
	FDI				1.9	2.2	4.9	2.5	2.4	4.9	
Estonia	*CA*				1.3	-7.3	-4.4	-9.1	-12.1	-9.2	
	FDI				9.9	9.4	5.7	3.4	5.7	11.2	
Hungary	*CA*	1.1	1.2	0.9	-11.0	-9.8	-5.7	-3.8	-2.2	-4.9	
	FDI		4.4	4.0	6.1	2.8	10.2	4.4	4.6	4.1	
Latvia	*CA*			14.0	19.2	5.5	-0.4	-5.4	-6.1	-10.2	-10.2
	FDI			2.2	2.1	5.9	4.0	7.4	9.2	5.6	5.9
Lithuania	*CA*				-3.2	-2.2	-10.2	-9.2	-10.2	-12.1	-11.2
	FDI				1.1	0.7	1.2	1.9	3.7	8.6	4.6
Poland♣	*CA*	5.2	-2.8	-3.7	-6.7	1.0	0.7	-2.3	-4.0	-6	-8
	FDI	0.2	0.4	0.8	2.0	1.9	2.9	3.1	3.4		
Romania	*CA*	-8.5	-3.5	-7.7	-4.7	-1.5	-5.0	-7.3	-6.1	-7.0	-3.8
	FDI		0.1	0.4	0.4	1.1	1.2	0.7	3.5	4.9	2.8
Slovak Rep.	*CA*				-4.8	4.9	2.2	-11.1	-10.1	-10.4	
	FDI				1.7	2.0	1.4	1.9	0.9	2.8	
Slovenia	*CA*				1.5	4.2	-0.1	0.2	0.2	0.0	-3.0
	FDI				0.9	0.9	0.9	1.0	1.8	0.8	0.4
Turkey	*CA*	-1.7	0.2	-0.6	-3.4	2.0	-1.4	-1.3	-1.4	0.9	
	FDI	0.5	0.5	0.5	0.3	0.5	0.5	0.4	0.4	0.5	
Greece	*CA*	-4.3	-1.8	-2.2	-0.8	-0.1	-2.4	-3.7	-4.0		
	FDI	1.2	1.3	1.2	1.1	1.0	0.9	0.9	0.8		
Portugal	*CA*	-0.3	-0.9	-0.2	0.3	-2.5	-0.1	-4.2	-5.4	-6.8	
	FDI	3.8	3.1	2.0	1.8	1.4	0.7	1.3	2.5	1.7	
Spain	*CA*	-3.7	-3.7	-3.7	-1.2	-1.3	0.1	0.1	0.5	-0.6	
	FDI	2.8	2.4	2.3	2.0	1.9	1.1	1.2	1.2	2.2	

Source of raw data: International Monetary Fund, *International Financial Statistics*. Data shown are CA (FDI) figures measured in USD in the balance of payments times annual average exchange rate against the USD divided by GDP at current prices.

♣: There was a methodological change in Poland's CA measurement in 1994: heavy cash inflows were moved from the financial account to the current account and were reclassified as unrecorded trade.

8.3 Some background calculations

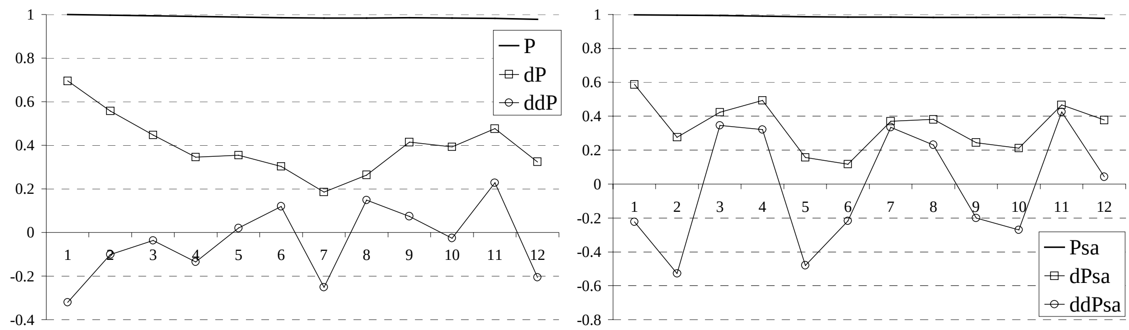
8.3.1 Seasonality

The fundamental price measure of the Czech Republic, Poland and Slovenia did not show a seasonal pattern, but the Hungarian and weighted foreign fundamental inflation did, so we adjusted the series accordingly. For comparison, we adjusted all time series, and as we cas seen in Figure 24, seasonal adjustment generated a strange autocorrelation pattern in the difference of Czech and Polish inflation data (ddPsa).

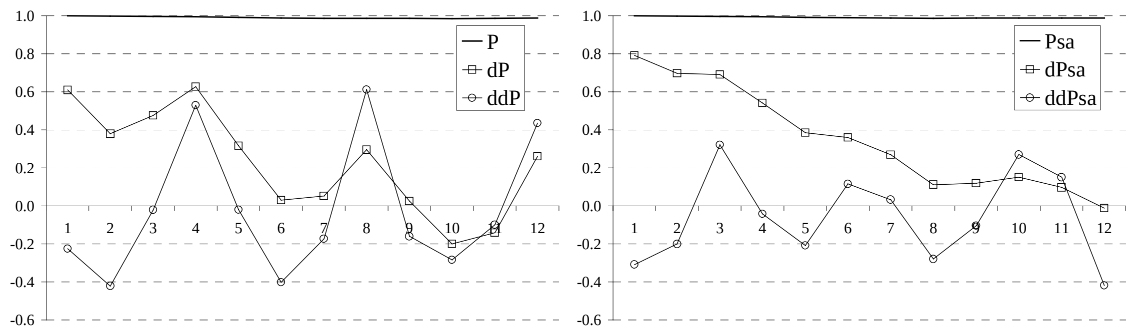
With the exception of administered prices no other variables seemed to be seasonal. Indeed, the very nature of administered prices is seasonal as most of them are set on a regular basis, i.e. at the beginning of the year. However, the seasonal pattern of these prices is probably known to the public, which is taken into account in the price setting behaviour.

Figure 24 (a-j). Correlogram of fundamental prices, original and seasonally adjusted

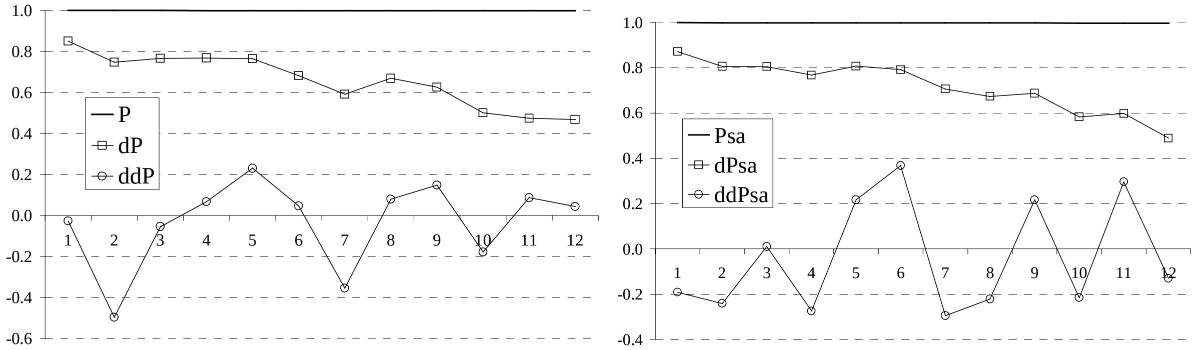
Czech Republic



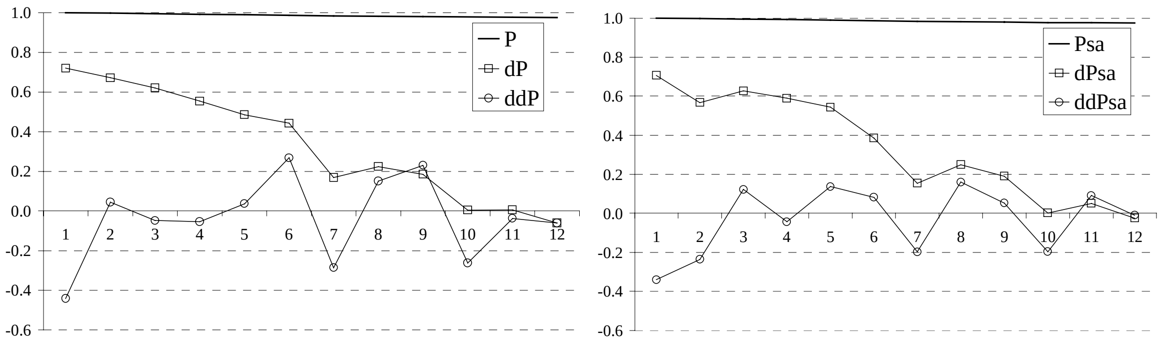
Hungary



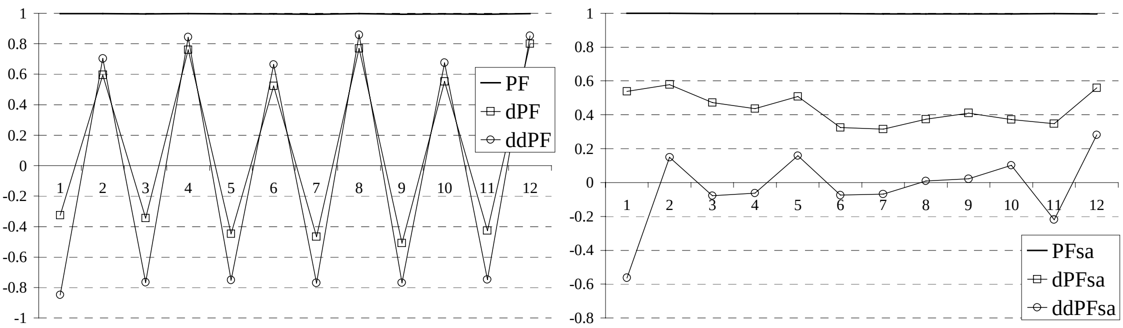
Poland



Slovenia



Foreign



8.3.2 Stationarity

Tables 8-9-**Fehler! Textmarke nicht definiert.**-10 show unit root tests. As frequently occurs, different unit root tests delivered different conclusions. Nonetheless, the broad picture suggests that most of the series can be appropriately approximated as I(1).

Table 8. Czech Republic: Unit root tests, 1993-2000

	PP(C)	lag1	DF(C)	lag2	DF(C)	lag3	DF(C)	PP(CT)	lag1	DF(CT)	lag2	DF(CT)	lag3	DF(CT)
P	-4.57 ***	1	-2.78 *	0	-5.78 ***	3	-1.66	0.27	0	0.55	0	0.55	3	1.48
ΔP	-1.77	0	-1.98	0	-1.98	4	-0.21	-3.69 **	0	-3.70 **	0	-3.70 **	4	-2.26
S	-1.38	0	-1.39	0	-1.39	1	-1.50	-2.63	0	-2.63	0	-2.63	0	-2.63
Δ	-5.04 ***	0	-5.02 ***	0	-5.02 ***	0	-5.02 ***	-4.97 ***	0	-4.96 ***	8	-3.00	0	-4.96 ***
PFSA	-4.91 ***	2	-2.80 *	0	-5.38 ***	4	-1.92	-0.75	2	0.74	0	-0.67	8	2.32
ΔPFSA	-5.14 ***	2	-3.32 **	2	-3.32 **	4	-0.24	-6.81 ***	3	-3.47 *	3	-3.47 *	4	-2.19
PRODA	-0.93	1	-1.27	7	-5.46 ***	7	-5.46 ***	-1.32	1	-1.56	7	-5.63 ***	7	-5.63 ***
ΔPRODA	-2.92 *	0	-3.01 **	7	-2.65	0	-3.01 **	-2.87	0	-2.97	7	-1.45	0	-2.97
PRODTNT	-2.43	0	-2.39	0	-2.39	0	-2.39	-3.19	0	-3.22	0	-3.22	0	-3.22
ΔPRODTNT	-6.10 ***	0	-5.98 ***	0	-5.98 ***	0	-5.98 ***	-5.92 ***	0	-5.81 ***	0	-5.81 ***	0	-5.81 ***
TOT	-1.28	2	-2.02	2	-2.02	2	-2.02	-1.90	2	-3.22	2	-3.22	2	-3.22
ΔTOT	-4.14 ***	1	-1.67	1	-1.67	1	-1.67	-4.08 **	1	-1.62	1	-1.62	1	-1.62
SGOVSA	-2.17	2	-1.86	8	-2.78 *	8	-2.78 *	-3.36 *	2	-1.86	8	-1.49	8	-1.49
ΔSGOVSA	-10.09 ***	0	-10.94 ***	0	-10.94 ***	7	-3.06 **	-10.37 ***	1	-3.60 **	7	-4.00 **	7	-4.00 **
FDI	4.92	0	4.21	0	4.21	0	4.21	0.26	0	0.04	0	0.04	0	0.04
ΔFDI	-2.82 *	0	-2.89 *	0	-2.89 *	1	-2.36	-4.62 ***	0	-4.63 ***	0	-4.63 ***	0	-4.63 ***
ONFA	-0.62	1	-0.86	1	-0.86	1	-0.86	-0.32	0	0.08	0	0.08	0	0.08
ΔONFA	-3.09 **	0	-3.02 **	0	-3.02 **	0	-3.02 **	-3.71 **	0	-3.68 **	0	-3.68 **	0	-3.68 **
FRIRB	-2.20	3	-6.34 ***	3	-6.34 ***	3	-6.34 ***	-1.95	3	-5.56 ***	3	-5.56 ***	3	-5.56 ***
ΔFRIRB	-6.50 ***	3	-3.25 **	3	-3.25 **	3	-3.25 **	-6.48 ***	3	-3.48 *	3	-3.48 *	3	-3.48 *
FRIRF	-2.02	1	-2.47	3	-3.83 ***	3	-3.83 ***	-2.20	1	-2.67	3	-5.10 ***	3	-5.10 ***
ΔFRIRF	-3.75 ***	1	-2.88 *	3	-3.54 **	4	-4.03 ***	-3.65 **	1	-2.77	3	-3.87 **	4	-4.10 **
ΔRPA	-7.50 ***	1	-3.36 **	0	-7.63 ***	0	-7.63 ***	-8.02 ***	2	-4.01 **	0	-8.02 ***	0	-8.02 ***
ΔRPE	-2.31	0	-2.27	0	-2.27	1	-1.06	-3.25 *	0	-3.14	7	-3.68 **	0	-3.14
ΔRPAE	-7.26 ***	1	-3.00 **	0	-7.48 ***	2	-3.38 **	-9.80 ***	2	-4.61 ***	0	-9.19 ***	0	-9.19 ***
IDIF	-1.28	0	-1.06	0	-1.06	2	-1.15	-1.13	0	-0.97	0	-0.97	2	-0.58
ΔIDIF	-5.88 ***	0	-5.92 ***	0	-5.92 ***	2	-2.39	-5.90 ***	0	-5.95 ***	0	-5.95 ***	2	-3.01
STP	-1.34	0	-1.34	0	-1.34	0	-1.34	-1.96	0	-1.78	0	-1.78	0	-1.78
ΔSTP	-5.69 ***	0	-5.70 ***	0	-5.70 ***	0	-5.70 ***	-5.65 ***	0	-5.66 ***	0	-5.66 ***	0	-5.66 ***

Notes: PP: Phillips–Perron *t*-test with truncation lag selected as the integer part of $4(T/100)^{2/9}$, T is the sample size; DF: (augmented) Dickey–Fuller *t*-test; C: constant in the test equation; CT: constant and trend in the test equation; lag1-2-3: lag selection for DF with (1) no autocorrelation in the residuals, (2) the highest significant lagged difference, (3) minimal Schwarz criterion, the longest possible lag length was set to 8; ***, **, or *: significant at 1%, 5%, or 10% respectively.

Table 9. Hungary: Unit root tests, 1992-2000

	PP(C)	lag1	DF(C)	lag2	DF(C)	lag3	DF(C)	PP(CT)	lag1	DF(CT)	lag2	DF(CT)	lag3	DF(CT)
PSA	-2.49	4	-1.54	6	-2.38	1	-2.15	0.90	4	0.75	6	-0.29	1	-0.28
ΔPDA	-3.67 ***	0	-3.34 **	6	-1.51	6	-1.51	-5.96 ***	0	-6.08 ***	5	-4.45 ***	6	-2.65
S	-1.42	0	-1.74	2	-1.37	0	-1.74	-0.33	0	0.24	2	-1.03	0	0.24
ΔS	-4.37 ***	0	-4.20 ***	1	-2.03	1	-2.03	-4.80 ***	0	-4.71 ***	0	-4.71 ***	0	-4.71 ***
PFSA	-2.49	4	-1.54	6	-2.38	1	-2.15	0.90	4	0.75	6	-0.29	1	-0.28
ΔPFSA	-3.67 ***	0	-3.34 **	6	-1.51	6	-1.51	-5.96 ***	0	-6.08 ***	5	-4.45 ***	6	-2.65
PRODA	-1.44	1	-0.26	8	1.29	4	0.60	-3.25 *	0	-3.30 *	0	-3.30 *	5	-0.88
ΔPRODA	-10.51 ***	0	-9.58 ***	0	-9.58 ***	4	-2.70 *	-12.01 ***	0	-9.83 ***	7	-2.79	7	-2.79
PRODTNT	0.07	0	-0.28	4	2.14	8	3.38	-3.00	0	-3.05	4	0.00	8	0.05
ΔPRODTNT	-7.18 ***	0	-6.84 ***	6	-0.82	3	-6.29 ***	-7.04 ***	0	-6.72 ***	6	-1.95	8	-4.76 ***
TOT	-2.32	4	-2.23	4	-2.23	4	-2.23	0.34	3	0.05	4	-0.90	2	-0.03
ΔTOT	-3.73 ***	3	-1.05	3	-1.05	3	-1.05	-4.47 ***	2	-4.69 ***	3	-2.32	1	-6.59 ***
SGOVSA	-2.31	4	-2.56	4	-2.56	4	-2.56	-3.35 *	4	-1.77	4	-1.77	4	-1.77
ΔSGOVSA	-9.15 ***	3	-2.07	3	-2.07	4	-3.21 **	-8.81 ***	3	-2.01	3	-2.01	4	-3.22
FDI	-0.92	0	-0.91	0	-0.91	0	-0.91	-1.70	0	-1.54	0	-1.54	0	-1.54
ΔFDI	-5.13 ***	0	-5.15 ***	0	-5.15 ***	0	-5.15 ***	-5.11 ***	0	-5.13 ***	0	-5.13 ***	0	-5.13 ***
ONFA	-1.15	0	-1.09	0	-1.09	0	-1.09	-1.98	0	-1.77	0	-1.77	2	-2.50
ΔONFA	-5.13 ***	0	-5.15 ***	0	-5.15 ***	0	-5.15 ***	-5.06 ***	0	-5.07 ***	0	-5.07 ***	0	-5.07 ***
FRIRB	-2.90 *	3	-4.03 ***	3	-4.03 ***	6	-4.27 ***	-2.88	4	-2.81	3	-5.68 ***	6	-3.69 **
ΔFRIRB	-6.00 ***	4	-4.51 ***	0	-5.99 ***	7	-2.90 *	-5.87 ***	4	-4.20 **	0	-5.86 ***	7	-2.93
FRIRF	-2.07	1	-2.53	3	-3.65 **	1	-2.53	-2.41	1	-2.89	3	-4.95 ***	3	-4.95 ***
ΔFRIRF	-4.72 ***	3	-3.73 ***	7	-4.24 ***	0	-4.70 ***	-4.63 ***	3	-3.67 **	7	-4.19 **	0	-4.60 ***
ΔRPASA	-7.76 ***	0	-6.85 ***	3	-5.34 ***	1	-6.10 ***	-7.97 ***	0	-6.87 ***	3	-5.59 ***	3	-5.59 ***
ΔRPE	-4.76 ***	1	-2.65 *	6	-1.95	0	-4.67 ***	-5.44 ***	1	-3.18	6	-2.50	0	-5.44 ***
ΔRPAESA	-6.88 ***	0	-6.53 ***	0	-6.53 ***	0	-6.53 ***	-6.79 ***	0	-6.43 ***	0	-6.43 ***	0	-6.43 ***
IDIF	-1.05	0	-0.53	1	-1.51	5	-1.16	-1.17	1	-1.90	1	-1.90	8	-4.16 **
ΔIDIF	-3.59 **	0	-3.66 ***	0	-3.66 ***	5	-2.28	-3.71 **	0	-3.81 **	0	-3.81 **	5	-2.03
STP	-0.78	0	-0.76	0	-0.76	0	-0.76	-2.23	0	-2.00	8	-2.80	0	-2.00
ΔSTP	-5.76 ***	0	-5.75 ***	0	-5.75 ***	0	-5.75 ***	-5.65 ***	0	-5.63 ***	0	-5.63 ***	0	-5.63 ***

Notes: PP: Phillips–Perron *t*-test with truncation lag selected as the integer part of $4(T/100)^{2/9}$, T is the sample size; DF: (augmented) Dickey–Fuller *t*-test; C: constant in the test equation; TC: constant and trend in the test equation; lag1-2-3: lag selection for DF with (1) no autocorrelation in the residuals, (2) the highest significant lagged difference, (3) minimal Schwarz criterion, the longest possible lag length was set to 8; ***, **, or *: significant at 1%, 5%, or 10% respectively.

Table 10. Slovenia: Unit root tests, 1993-2000

	PP(C)	lag1	DF(C)	lag2	DF(C)	lag3	DF(C)	PP(CT)	lag1	DF(CT)	lag2	DF(CT)	lag3	DF(CT)
P	-5.38 ***	0	-6.87 ***	0	-6.87 ***	0	-6.87 ***	-5.04 ***	0	-5.03 ***	0	-5.03 ***	0	-5.03 ***
ΔP	-2.72 *	0	-2.78 *	0	-2.78 *	0	-2.78 *	-2.91	0	-2.97	0	-2.97	0	-2.97
S	-3.15 **	0	-3.35 **	0	-3.35 **	3	-0.62	-5.47 ***	0	-5.27 ***	0	-5.27 ***	2	-5.09 ***
Δ	-5.98 ***	2	-5.35 ***	0	-5.61 ***	2	-5.35 ***	-5.63 ***	2	-4.86 ***	0	-5.38 ***	2	-4.86 ***
PFSA	-5.48 ***	2	-2.73 *	7	-3.22 **	1	-4.67 ***	-1.27	2	-1.16	7	-1.85	1	-1.05
ΔPFSA	-5.60 ***	3	-1.95	1	-2.18	3	-1.95	-8.97 ***	2	-3.82 **	6	-3.08	0	-8.81 ***
PRODA	-1.98	2	-0.17	8	-0.07	1	-0.88	-3.99 **	2	-3.71 **	7	-4.80 ***	7	-4.80 ***
ΔPRODA	-3.35 **	1	-2.80 *	0	-3.45 **	1	-2.80 *	-3.31 *	1	-2.39	0	-3.39 *	1	-2.39
PRODTNT	-0.84	0	-0.94	0	-0.94	1	-0.10	-3.41 *	0	-3.27 *	0	-3.27 *	0	-3.27 *
ΔPRODTNT	-7.09 ***	0	-6.73 ***	0	-6.73 ***	0	-6.73 ***	-6.91 ***	0	-6.60 ***	0	-6.60 ***	2	-4.67 ***
TOT	-3.47 **	7	-2.07	0	-3.98 ***	7	-2.07	-0.83	7	-0.51	0	-0.49	7	-0.51
ΔTOT	-2.07	7	-0.48	2	-0.77	2	-0.77	-3.85 **	7	-1.79	0	-3.77 **	4	-2.28
SGOVSA														
ΔSGOVSA														
FDI	0.05	0	0.13	0	0.13	0	0.13	-1.70	0	-1.51	7	-3.51 *	0	-1.51
ΔFDI	-4.46 ***	0	-4.45 ***	0	-4.45 ***	0	-4.45 ***	-4.44 ***	0	-4.43 ***	0	-4.43 ***	0	-4.43 ***
ONFA	-2.12	2	-2.11	2	-2.11	0	-1.97	-1.19	2	-1.47	6	-0.78	0	-1.25
ΔONFA	-5.01 ***	0	-4.93 ***	5	-2.18	0	-4.93 ***	-6.14 ***	1	-3.05	8	-4.00 **	0	-6.20 ***
FRIRB	-2.45	3	-4.59 ***	3	-4.59 ***	4	-3.26 **	-2.39	3	-4.05 **	3	-4.05 **	4	-3.13
ΔFRIRB	-7.41 ***	0	-7.15 ***	3	-3.69 **	4	-2.68 *	-7.10 ***	0	-6.78 ***	3	-3.75 **	4	-2.60
FRIRF	-1.83	1	-2.25	1	-2.25	4	-1.59	-2.30	1	-2.86	3	-5.16 ***	4	-3.36 *
ΔFRIRF	-3.81 ***	0	-3.75 ***	3	-3.40 **	4	-3.96 ***	-3.71 **	0	-3.64 **	3	-3.49 *	4	-3.74 **
ΔRPASA	-4.17 ***	0	-4.27 ***	0	-4.27 ***	0	-4.27 ***	-4.35 ***	0	-4.39 ***	7	-2.73	0	-4.39 ***
ΔRPE	-7.61 ***	3	-2.21	1	-9.70 ***	3	-2.21	-7.51 ***	3	-2.10	5	-2.55	2	-5.32 ***
ΔRPAESA	-5.13 ***	1	-4.93 ***	7	-3.01 *	7	-3.01 *	-5.15 ***	1	-4.93 ***	7	-3.02	7	-3.02
IDIF	-1.80	5	-3.82 ***	6	-3.40 **	8	-1.70	-1.83	1	-3.85 **	1	-3.85 **	8	-0.39
ΔIDIF	-4.23 ***	1	-5.72 ***	1	-5.72 ***	8	-4.64 ***	-4.31 **	2	-5.51 ***	2	-5.51 ***	8	-3.98 **
STP	-1.62	0	-1.71	0	-1.71	0	-1.71	-2.32	0	-2.35	0	-2.35	0	-2.35
ΔSTP	-5.25 ***	0	-5.14 ***	0	-5.14 ***	0	-5.14 ***	-5.12 ***	0	-5.04 ***	0	-5.04 ***	0	-5.04 ***

Notes: PP: Phillips–Perron *t*-test with truncation lag selected as the integer part of $4(T/100)^{2/9}$, T is the sample size; DF: (augmented) Dickey–Fuller *t*-test; C: constant in the test equation; CT: constant and trend in the test equation; lag1-2-3: lag selection for DF with (1) no autocorrelation in the residuals, (2) the highest significant lagged difference, (3) minimal Schwarz criterion, the longest possible lag length was set to 8; ***, **, or *: significant at 1%, 5%, or 10% respectively.

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