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Testing for Structural Change in Cointegrated Relationships: Analysis of Price–Wages Models for Poland and Hungary

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Abstract. In previous studies concerning short- and long-run relationships for price–wage models, the cointegration analysis has been developed assuming the existence of a unique cointegration parametrisation. These empirical results reveal the presence of significant relationships, both in the short and in the long run, among prices, wages, labour productivity and exchange rate. In this paper we intend to develop the possibility of a more general type of cointegration, allowing for a change at an unknown time period in the sample. At this end we will consider mainly the long-run relationship among these variables using the testing procedure suggested by Gregory and Hansen (1996a,b). This permits us to consider a multivariate extension of the endogenous break univariate approach and, in the meantime, this enables us to test for cointegration in the presence of possible structural breaking cointegrated relationships under the alternative. The empirical analysis of a multivariate model for price–wage relationship both for Poland and Hungary, over the period 1970–1996, is presented and discussed.

JEL Classification: C12, C52

Key words: structural change, cointegrated system, changes at unknown period, wage–price dynamic models

1. Introduction

The long-run relationships between prices and wages, through two different structural models, have been investigated in two previous papers (Golinelli and Orsi, 1994, 1998). More precisely, the first paper was devoted to the study of short- and long-run price–wage dynamics in Poland by examining the cointegration between domestic prices, import prices, wages and labour productivity. The second paper extended the analysis both by including other macroeconomic variables like exchange rate and employment, and by using the same model to analyse the Hungarian economy.

A common feature of the two previous papers is that the cointegration studies were developed assuming the existence of a unique cointegration parametrisation, in the sense that no regime change was allowed to the cointegration vector over the period considered, 1970–1994. The reason for that choice was mainly the fact that too little statistical information was available for the period after the break occurred at the end of the 1980s, preventing us to gather significant empirical evidence showing the existence of a new long-run equilibrium between the variables.

The aim of this paper is to perform a more general type of cointegration analysis, namely a model specification where the cointegrating vector is allowed to change at an unknown time period within the sample. The standard tests used for cointegration are not useful for such an analysis since they implicitly admit that the cointegration vector is time invariant under the alternative. In other terms, we extend our cointegration analysis by admitting the possibility of a break in the cointegrating relationship, allowing the data exhibiting the exact period of occurrence of the shift. Theoretical properties along with the distributional theory of testing procedures for the presence of a break in the cointegrating vector have been investigated by Gregory and Hansen (1996a). This constitutes our point of departure in the theoretical developments we intend to do in this paper. More precisely, by using this testing procedure we will do a multivariate extension of the endogenous break univariate approach and, in the meanwhile, this represents a test for cointegration in the presence of a structural-breaking cointegrated relationship. This test allows us to assess if the cointegration amongst the variables of interest holds over a first sample subperiod, at an unknown separation point, and if it persists in the second subperiod or if it shifts to another long-run relationship.

This approach has been applied to two cointegrated systems, the price–wage spiral and a dynamic model of prices, wages, productivity and exchange rate. For the period 1970.1–1997.1 quarterly time series, both of these two systems show strong evidence of structural changes. Section 2 contains a preliminary analysis of the variables of interest at univariate level, either by graphical inspection of data or by analysing the results of a formal unit roots test. In Section 3 we develop testing procedures for detecting the structural change in a cointegration system. Moreover, we apply these procedures to estimate the long-run relationships in the presence of structural changes. Section 4 summarises our main findings.

2. Preliminary univariate analysis

This section is devoted to analysing a common set of variables (the corresponding labels are in parentheses) of interest for both the economics of Poland and Hungary, namely: the consumer prices (*lpc*), the nominal per capita wage (*lw*), the exchange rate against the US dollar (*lex*), the real per capita wage (*lwr*), the labour productivity or per capita output (*leta*) and the employment-population ratio (*lnpop*). All variables are in logarithms, and their data sources and descriptions are reported in Appendix Al. The first three variables are expressed in nominal terms and measure the domestic and foreign sources of inflation; the other variables are in real terms and intend to represent the supply side effect of the economy.

The plots of the variables, both in levels and in first differences, are depicted in Figures 1 and 2, respectively for Poland and Hungary. In the present paper, we

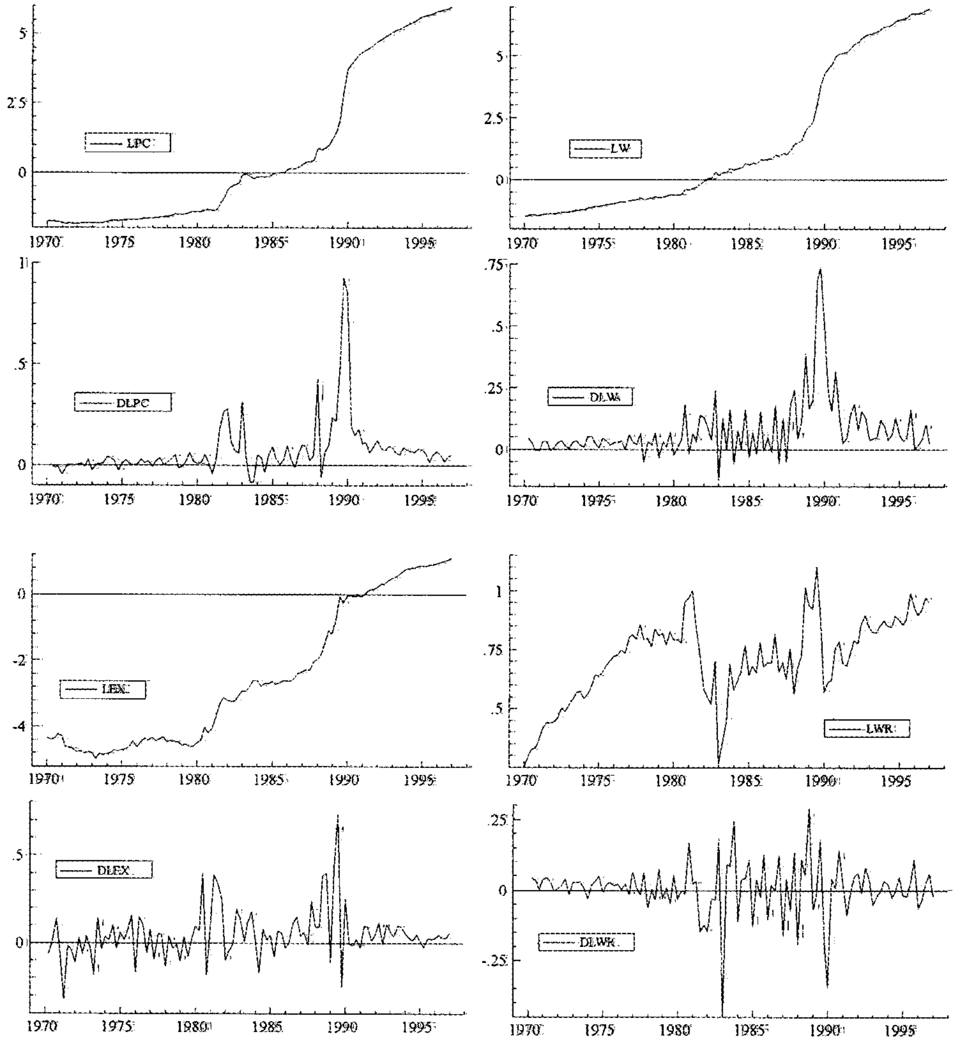

Figure 1a. Nominal variables (*lpc, lw, lex*) and real wages (*lwr*) in Poland, levels and first differences.

will analyse the structural models by using seasonally unadjusted data, since the use of filtered series by some filters associated with the X-11 seasonal adjustment program would cause a tendency to disguise the structural instability (Ghysels and Perron, 1996). A first look at graphs permits to ascertain the presence of breaks in the level of Polish price, wage and exchange rate variables reported in Figure 1a as well as in the labour productivity and employment–population ratio depicted in Figure 1b. Indeed, they grow during the whole period and their levels exhibit two major increases in both the beginning and the end of the 1980s. Looking at

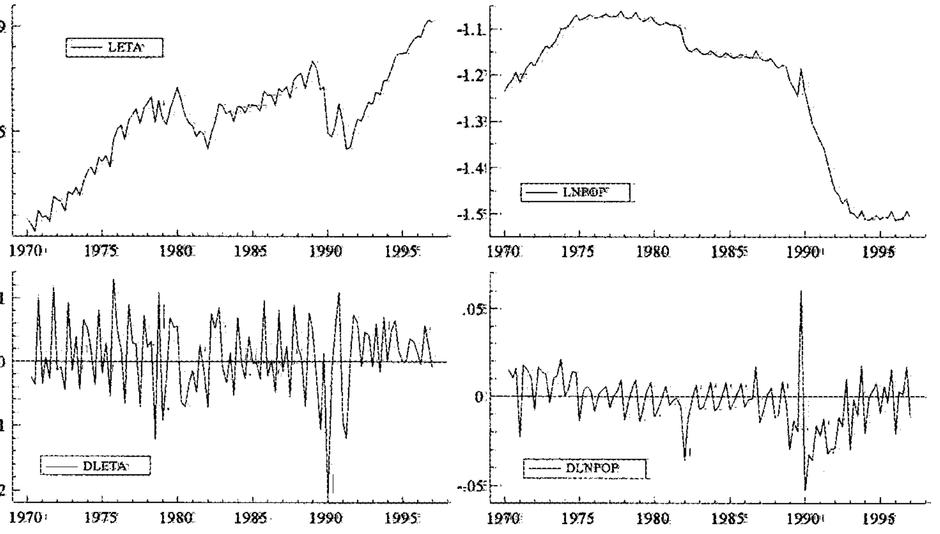

Figure 1b. Real variables (*leta, lnpop*) in Poland, levels and first differences.

the series in first differences, the above mentioned two acceleration periods are characterised by a fast increase in the variable fluctuations.

As far as stationarity is concerned, plots in Figures 1a and b need formal tests to ascertain the presence of one or more unit roots in their stochastic processes representation. In other terms, we want to inspect the possibility that the non-stationarity they show is of a deterministic or a stochastic nature. The use of Dickey-Fuller, augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests[2] reveal very similar results for all the variables, in the sense that a significant (not significant) value for the ADF test is, very often, associated with a significant (not significant) value for the PP test. Both ADF and PP tests provide evidence of a unit root for consumer prices, nominal per capita wage, labour productivity and exchange rate, but not for the labour–population ratio. The resulting order of integratedness higher than 1 for this last variable seems implausible, at least from an economic point of view. The plot in Figure 1b for the first differenced labour–population ratio (*dlnpop*) suggests the presence of a break around 1990 along with a slightly persistence of data over and below the corresponding sample mean; this could induce an apparent unit root even in the differenced data. For this reason, we will further analyse the data by using alternative unit root tests, which may take care of possible breaks in *lnpop*.

As far as Hungary data are concerned, the plots are reported in Figures 2a and b. The graphical inspection of the series suggests a growing pattern that is more stable than the corresponding Polish variables, while the seasonal profile of the data appears considerably more marked.

As we can see from the figures, the period of occurrence of (possible) breaks is different for the two countries; in Hungary, the nominal variables tend to break only at the end of the 1980s, while the real ones denote a unique change in levels at the

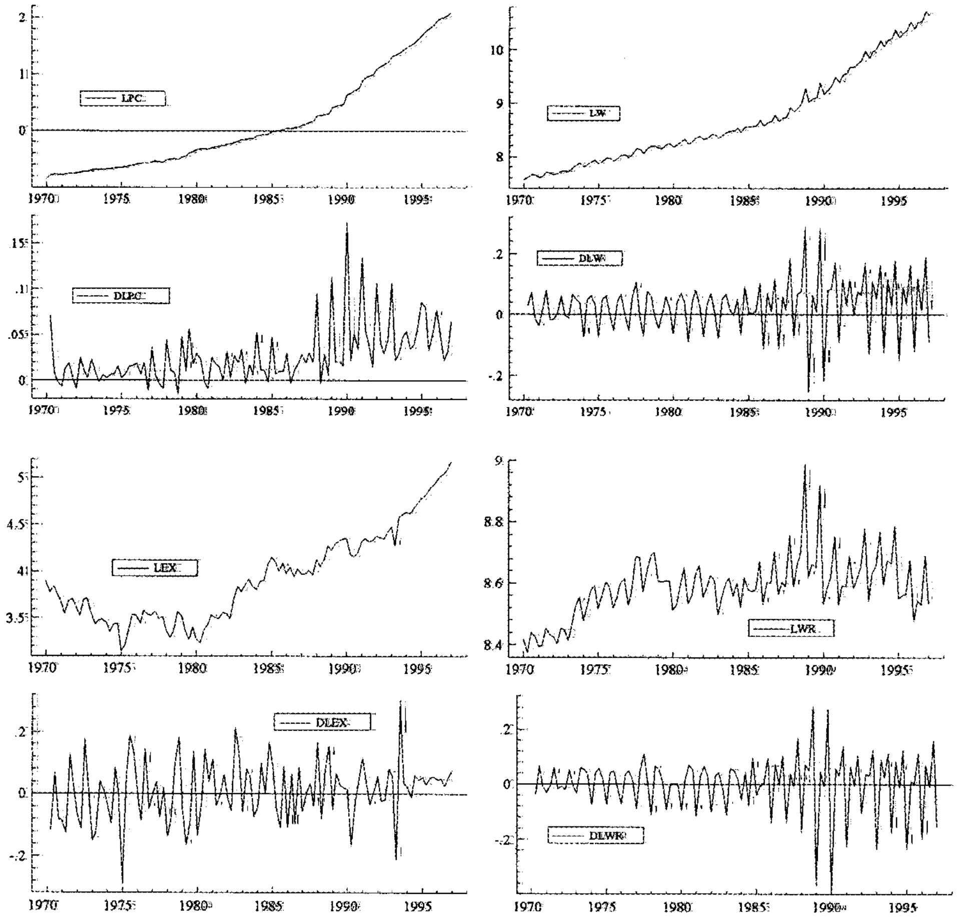

Figure 2a. Nominal variables (*lpc, lw, lex*) and real wages (*lwr*) in Hungary, levels and first differences.

beginning of the 1990s. The ADF and PP tests for Hungarian data provide evidence of a unit root for labour productivity and exchange rate, while the employment–population ratio seems integrated of higher order; all these results confirm what we found for Poland.

Moreover, differently from Poland, the ADF test for Hungary suggests that consumer prices and nominal per capita wages are integrated of an order higher than one. It is also important to note that in the case of Hungarian variables the number of times ADF and PP tests disagree is higher than in the Polish case. This outcome raises the already known serial correlation adjustment issue that the two testing procedures tackle in different ways (parametric and non-parametric). Following the ADF test results, we conclude that a number of variables in logarithmic first differences are integrated, and this fact would seem quite implausibly from an economic point of view, since they respectively measure the growth rate

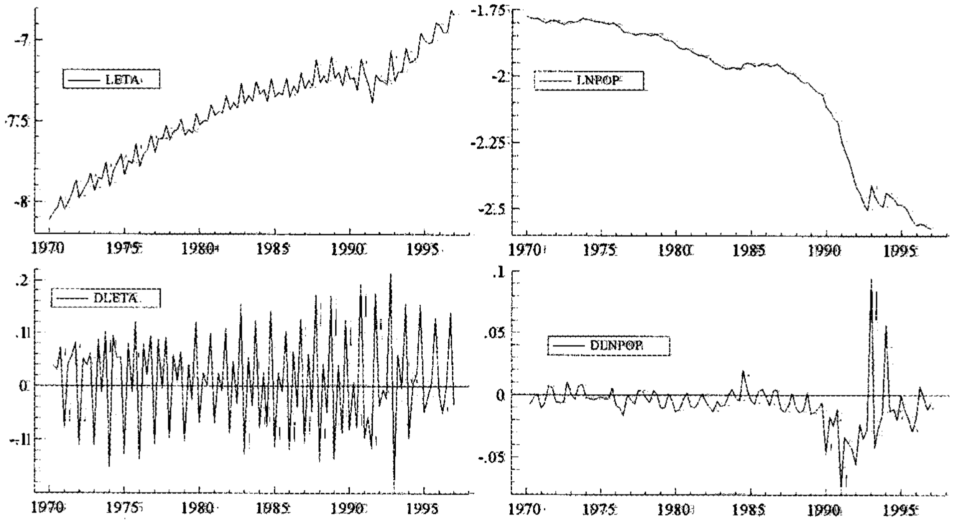

Figure 2b. Real variables (*leta, lnpop*) in Hungary, levels and first differences.

of an employment ratio, the inflation rate and the rate of growth of per capita nominal wages. On the other side, looking at the differenced variables in Figures 2a and b we suggest a possible stochastic nonstationarity of the first differences and, consequently, levels of the variables that are $I(2)$. Also in this case a possible explanation of this nonstationarity is related to the fact that all the variables show at least a relevant structural break, as well as some evidence for seasonal unit roots.

In order to use models for unit root tests that are congruent with the main statistical features we have found in the preliminary data analysis, we follow the procedure suggested by Perron (1997). It is well known that structural breaks in the deterministic components of the stochastic process tend to bias both ADF and PP tests towards the unit root null hypothesis; for this reason, Perron (1997) proposes both innovative (Model 1 and Model 2) and additive (Model 3) type outlier test of a $I(1)$ null hypothesis against a $I(0)$ alternative with a single break,[3] occurred at an unknown point in time denoted as T_b.[4]

The empirical evidence for the breaks reported in Table 1 deserves some specific comments, and in doing them we will try to combine such explanation with the informational content of the graphical inspection of data reported above. For Poland, the results provide evidence of possible breaks at the beginning or at the end of the 1980s which are characterised by an intercept change with or without a change in the slope of the trend function. According to the Perron procedure, Models 1 or 2 are recommended in this case, while Model 3 is completely mis-specified because it does not allow for the important shift in levels occurred in the 1980s. More precisely Model 1 imposes a common trend slope (before and after the break date); a clear perception we can get from a visual inspection of Polish data where the two trend slopes appear significantly different, and consequently we prefer to

Table I. The Perron sequential tests for the Polish variables

Name	Model 1			Model 2			Model 3		
	t_α^*	k	T_b	t_α^*	k	T_b	t_α^*	k	T_b
lpc	−7.42**	9	1989.1	−4.48	9	1989.1	−3.56	10	1983.2
lw	−8.10***	9	1989.1	−4.03	8	1981.3	−4.15	8	1984.1
lwr	−4.85	8	1981.2	−6.06**	8	1981.2	−3.00	8	1974.2
leta	−4.12	16	1988.4	−4.49	12	1989.3	−3.43	16	1975.4
lnpop	−6.54***	17	1989.3	−3.69	12	1990.1	−3.97	12	1984.4
lex	−4.33	10	1988.1	−3.73	5	1977.3	−3.70	5	1977.3

*, ** and *** respectively 10%, 5% and 1% significant. Critical values are from Perron (1997, Table 1, pp. 362–363).

Table II. The Perron test for the first difference employment-population ratio in Poland

Variable	T_b	k	t_α^* statistics	Critical values 1%	5%	10%
dlnpop	1989, 3rd	16	−5.94	−6.21	−5.55	−5.25

The critical values are from Perron (1997, Table 1, p. 362).

use the Model 2 since it appears the most adequate. In this case, the presence of breaks under the alternative would not change the conclusions suggested by the ADF and PP tests, i.e. many of the Polish variables of interest are first order integrated. The employment–population ratio (*lnpop*) is the only variable that, in the light of the classical ADF and PP tests, seems integrated of order bigger than one. The occurrence of a break, both in the intercept and in the slope of the trend function, makes possible the use of the Perron Model 2 to test the null of a unit root in the series in differences against an alternative of a stationary process with breaks in both the intercept and the slope of the trend function. It is also very important to note in Table I (Model 2) that per capita nominal wages and consumer prices seem to be cointegrated with parameters 1 and −1 in presence of a break, in the deterministic components, at the beginning of the 1980s. The results presented in Table II clearly show that *dlnpop* (first difference of the employment–population ratio) is a stationary variable with the deterministic component that breaks in the 3rd quarter of 1989.

The Hungarian data, as we have already seen, provide evidence in favour of a possible change in trend slope, consequently we think Models 2 and 3 represent the correct model specification in order to test for the presence of a unit root. Anyway the results reported in Table III, independently from the specific model chosen, reveal that the variables of interest are not stationary (in level), even taking into account for possible breaks. As a matter of fact, a 10% significant test for nominal wages and employment–population ratio represents too weak evidence to sustain

Table III. The Perron sequential tests for the Hungarian variables

Name	Model 1			Model 2			Model 3		
	t_{α}^{*}	k	T_b	t_{α}^{*}	k	T_b	t_{α}^{*}	k	T_b
lpc	−2.26	10	1989.2	−4.85	4	1985.3	−3.91	4	1994.4
lw	−3.42	20	1988.3	−5.45*	4	1985.3	−4.11	9	1987.1
lwr	−3.60	4	1972.4	−3.58	4	1972.4	−3.45	4	1976.2
leta	−3.68	9	1988.4	−3.52	9	1979.2	−3.40	9	1980.2
lnpop	−4.93*	8	1990.3	−5.03	17	1986.1	−4.76*	17	1988.3
lex	−4.12	10	1977.1	−3.49	12	1977.1	−3.47	12	1978.2

*, ** and *** respectively 10%, 5% and 1% significant. Critical values are from Perron (1997, Table 1, pp. 362–363).

Table IV. The Perron test for some Hungarian variables in first differences

Variable	T_b	k	t_{α}^{*} statistics	Critical values		
				1%	5%	10%
dlnpop	1989, 3rd	7	−6.37	−6.21	−5.55	−5.25
dlpc	1988, 3rd	6	−5.43	−6.21	−5.55	−5.25
dlw	1987, 2nd	6	−5.76	−6.21	−5.55	−5.25

The critical values are from Perron (1997, Table 1, p. 362).

the presence of a break, as it appears in the mid-1980s, in the trend function. As far as the *lnpop, lpc* and *lw* variables are concerned we use the Perron Model 2 to test for the presence of a unit root in the differences of these variables, against the alternative of a stationary process with breaks both in the intercept and in the slope of the trend function. The results reported in Table IV clearly show that all the Hungary variables in first differences are stationary. Incidentally, this provides an explanation for the first order integratedness of some variable in differences, as evidenced by classical ADF and PP tests we conduct in the preliminary analysis, in the sense that these findings are able to ascribe to the presence of relevant structural breaks in the deterministic components of the stochastic processes.

In summary, the use of the unit roots Perron test in presence of structural breaks reinforces the main conclusions drawn from the ADF and PP unit roots tests: at least all the variables of interest reveal to be first order integrated.[5]

As noted, a stationary process for real per capita wage in Poland would suggests a possible cointegrating vector (1, −1) between nominal wages and consumer prices, while the same is not true for Hungary. This result confirms the main findings for the wage equations and reported in Golinelli Orsi (1994, 1998) and introduces an important issue to be tackled in the next section, that will be devoted to the cointegration analysis in presence of structural breaks.

3. Testing for cointegration in presence of regime shifts

The short- and long-run relationships among prices, wages, labour productivity and exchange rate in Poland and Hungary by using the Johansen (1988) maximum likelihood approach for cointegrated systems, has been analysed in two previous studies by Golinelli and Orsi (1994, 1998). The cointegrating relationships have been studied by considering conditional models that are not affected by breaks in the short- and the long-run parameters. For this reason, both a number of dummies and some breaking variables (as the private to total share in employment and the employment population ratio) were included in the cointegrated VAR.

In order to detect the presence of a structural break in the cointegrating relationship, in the subsequent analysis we follow an alternative procedure recently proposed by Gregory and Hansen (1996a, b) and developed in the field of the Engle–Granger (1987) cointegration analysis (EG approach). The Gregory–Hansen (GH) statistics can be seen as a multivariate extension of the endogenous break univariate approach used in the previous paragraph, and enables us to test for cointegration by taking into account for a breaking cointegrated relationship under the alternative. The GH test allows assessing if the cointegration amongst the variables of interest holds over a first period of time and then, in a priori unknown period T_b, it shifts to another long run relationship.[6] Gregory and Hansen introduce four different models to take into account the structural change in the cointegrating relationship under the alternative. The first model is a *level shift* model, denoted by C and defined as:

$$y_t = \mu + \theta DU_t + \alpha' x_t + u_t \tag{1}$$

where y_t is a scalar variable, x_t is an m-dimensional vector of explanatory variables (both y_t and x_t are supposed to be I(1)), u_t is the disturbance term, DU_t is a step dummy variable defined as: $DU_t = 1(t > T_b)$, where 1(.) denotes the indicator function. Parameters μ and θ measure respectively the intercept before the break in T_b and the shift occurred after the break, while α are the parameters of the cointegrating vector.

The second model is the *level shift with trend* model, denoted as *C/T*:

$$y_t = \mu + \theta DU_t + \beta t + \alpha' x_t + u_t \tag{2}$$

where a β sloped time trend is added in the level shifting model (1).

In the third model the cointegrating vector parameters are allowed to shift and this model is labelled as the *regime shift* model, denoted by *C/S*:

$$y_t = \mu + \theta DU_t + \alpha' x_t + \delta' x_t DU_t + u_t \tag{3}$$

where δ measures the change in the cointegrating vector after the shift. These three models were introduced in Gregory and Hansen (1996a) and are followed by a fourth model, presented in Gregory and Hansen (1996b), labelled the *regime and* *trend shift* model, denoted by *C/S/T*:

$$y_t = \mu + \theta DU_t + \alpha' x_t + \delta' x_t DU_t + \beta t + \varphi DU_t t + u_t. \quad (4)$$

In this last model, with respect to the *C/S* model, a broken trend variable is added and the parameter φ measures the change in the trend slope after the break.

All the GH tests are residual based, and the null of no cointegration corresponds to a unit root in the OLS residuals of models *C, C/T, C/S* and *C/S/T*. In practice one would have to compute a cointegration test statistic for each possible regime shift and to take the smallest value across all possible break points. Like in the Perron (1997) univariate test, the timing of the shift T_b is treated as unknown. In order to test for the presence of unit roots in the residuals, Gregory and Hansen suggest using the classical ADF residual test and the Phillips (1987) test statistic. In this paper we use the ADF test statistic on residuals obtained from the *C, C/T, C/S* and *C/S/T* models, where the lag length k was set as in Perron (1997), following a general to specific procedure.

The distributional properties of the contegration with breaks test statistic are analysed in Gregory and Hansen (1996a); the approximate asymptotic critical values for these tests, obtained through simulation methods, are reported in Gregory and Hansen (1996a and Table I, and 1996b, Table 1). Since a major drawback of the GH tests (as for the EG approach) is the a priori hypothesis of unit cointegration rank, we applied the GH tests to the structural equations presented in Golinelli and Orsi (1998).

3.1. THE LONG-RUN PER CAPITA WAGE EQUATION FOR POLAND

At first, we apply the GH procedure to the wage equation for Poland as it was estimated in Golinelli and Orsi (1998, Table 3), where the employment–population ratio is introduced as a proxy for labour demand and supply effects in the labour market.[7]

$$lwr_t = \alpha_1 leta_t + \alpha_2 lnpop_t + u_t. \quad (5)$$

The four graphs in Figure 3 summarise the results of our analysis; each point in a single graph corresponds to a specific ADF statistic of each of the four breaking Models (1–4). A visual inspection of Figure 3 supports the idea that the real per capita wage equation for Poland is characterised by two possible structural breaks at the beginning and at the end of the 1980s. Clearly, it is to be remarked that each minimum value test statistic for each model has a specific critical value that the graphs omit to report.

A formal testing analysis is reported in Table V where the most relevant information is summarised. The second and the third columns respectively report the minimum ADF test statistic of no cointegration amongst the explanatory variables and the lag length. The endogenously detected period of break and the OLS estimates of the break-cointegrating relationships are also reported in the following

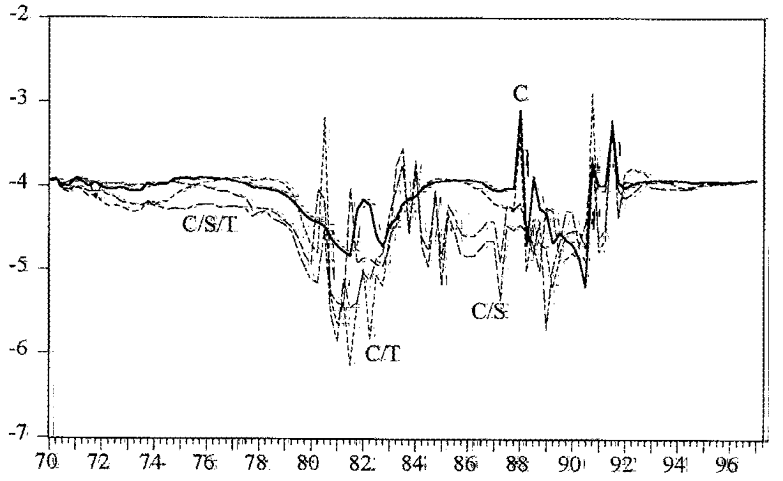

Figure 3. Regime shifts ADF plots for per capita real wage equation in Poland.

Table V. Testing for regime shifts in the real wage equation for Poland

	ADF	*k*	T_b	*leta* (α)	*lnpop* (α)	*leta* (δ)	*lnpop* (δ)
EG	−4.05	0					
EG/T	−4.04	0					
C	−5.19*	4	1990.3	0.561	0.758		
C/T	−6.14**	8	1981.3	0.172	0.480		
C/S	−5.43	4					
C/S/T	−5.85	6					

* and ** respectively 5% and 1% significant. δ parameters are reported only if *C/S* or *C/S/T* models breaks are significant.

columns only if the ADF test is at least 5% significant. The cointegration Engle and Granger test suggests that the null of no cointegration be not rejected for both models with (*EG/T*) and without (*EG*) trend against the alternative of a stable cointegrating relationship.

As far as the cointegrated breaking models are concerned, it is worth noting that both the intercept shift model (*C*) and the trended model with intercept shift (*C/T*) reject the null, but the significant breaking periods are situated in different point in time: the third quarter of 1990 and the third quarter of 1981, respectively.

On balance, we can conclude that there is some evidence in favour of the existence of a long-run cointegrating relationship between these variables even though all the conventional ADF tests fail to reject the null. In other terms, if standard ADF statistic does not reject the null while the GH test rejects the null, this implies that it is very important to allow for a structural change in the cointegrating vector, adopting a testing procedure that allows for an endogenous break identification.

Table VI. Real wage equation for Poland

Dependent variable is LWR			
109 observations used for estimation from 1970Q1 to 1997Q1			
Regressor	Coefficient	Standard Error	T-Ratio[Prob]
CONST	6.8832	.56461	12.1909[.000]
S1	−.050779	.029904	−1.6980[.093]
S2	−.035991	.030243	−1.1901[.237]
S3	−.013513	.030356	−.44514[.657]
LETA	.56084	.052649	−10.6525[.000]
LNPOP	.75811	.19303	3.9275[.000]
D903	.29711	.069185	4.2944[.000]
R-Squared	.64491	R-Bar-Squared	.62403
S.E. of Regression	.10985	F-stat. F(6, 102)	30.8756[.000]
Mean of Dependent Variable	.70949	S.D. of Dependent Variable	.17915
Residual Sum of Squares	1.2308	Equation Log-likelihood	89.6966
Akaike Info. Criterion	82.6966	Schwarz Bayesian Criterion	73.2769
DW-statistic	.68864	ADF (4)	−5.19
S1, S2, S3 seasonal dummies			

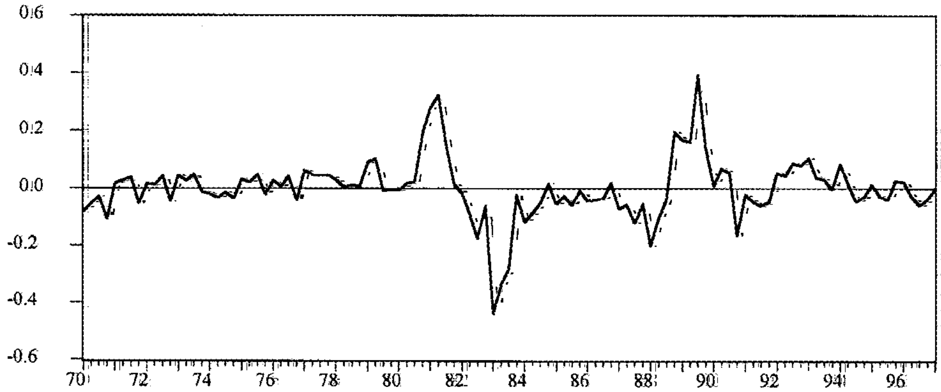

Figure 4. Residual plot of wage equation (*ecmwp*), Poland.

If we suppose the existence of a long-run relationship and one shift in intercept, it is interesting to note that the estimates of both the labour productivity and the employment–population ratio effects are not significantly different with respect to Golinelli and Orsi (1998) estimates, the same is not true if we specify a broken trend model.

From an economic point of view, the cointegrated model *C* estimates (where a simple intercept shift is added to the long run wage equation proposed in Golinelli and Orsi, 1998), seem to be the more acceptable. Given our preference to this

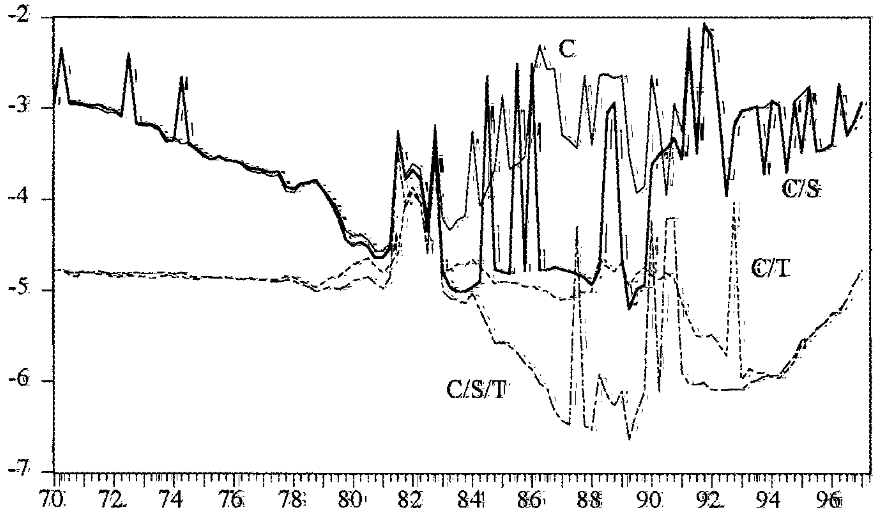

Figure 5. Regime shifts ADF plots for price equation in Poland.

long-run specification, Table VI reports the detailed estimation results and Figure 4 depicts the corresponding residuals.

3.2. THE LONG-RUN PRICE EQUATION FOR POLAND

Our point of departure is represented by the results reported in Golinelli and Orsi (1998, and Table III) where the exchange rate against the US dollar is also considered:[8]

$$lpc_t = \alpha(lw_t - leta_t) + (1 - \alpha)lex_t + u_t \tag{6}$$

The α estimate (0.108) suggests the relevant role played by foreign factors in driving the domestic prices, a typical situation for a command economy with repressed inflation.

From a methodological point of view, the use of the GH test allows to tackle the macroeconomic issue of possible regime shifts in the cointegrating relationships due to the transition from a centralised to a liberalised economy at the end of the 1980s: we would expect a reduction in foreign prices influence as a consequence of important reforms carried out during the 1880s.

Our main findings are summarised in Figure 5 where we report the point results for the GH test applied to previous Model (6) by considering the different regime shift models *C, C/T, C/S* and *C/S/T*. A first look at the graphs suggests that the price equation break tests have a greater variability than the wage equation tests previously depicted in Figure 3; in addition, though not clearly distinguishable in Figure 5, the tests based on the models *C/T* and *C/S/T* reveal to be more volatile than those related to the models *C* and *C/S*, the latter two being more acceptable than the first two from a theoretical point of view, since they do not include a linear trend in the long run relationship.

Table VII. Testing for regime shifts in price equation for Poland

	ADF	*k*	T_b	Parameter (α)	Estimates (δ)
EG	−2.95	4			
C	−4.56	4			
C/S	−5.20*	4	1989.2	0.67	0.43

Table VIII. Consumer price equation for Poland

Non-linear regression formula:
lpc=a0+a1*d892+a2*(lw−leta)*(1−d892)+(1−a2)*lex* (1−d892)+a3*(lw−leta)+d892+ a4*lex*d892+a5*s1+a6*s2+a7+a7*s3
109 observations used for estimation from 1970Q1 to 1997Q1
S1, S2, S3 seasonal dummies

Parameter	Estimate	Standard Error	T-Ratio[Prob]
A0	−5.9642	.55242	−10.7965[.000]
A1	−4.0300	.91166	4.4205[.000]
A2	.67224	.042804	15.7052[.000]
A3	.98162	.050997	19.2485[.000]
A4	.26071	.10162	2.5656[.012]
A5	.052976	.030915	1.7136[.090]
A6	.046080	.031238	1.4751[.143]
A7	.013561	.031291	.43338[.666]
R-Squared	.99845	R-Bar-Squared	.99834
S.E. of Regression	.11453	F-stat. F(7, 101)	9300.0[.000]
Mean of Dependent Variable	.72880	S.D. of Dependent Variable	2.8140
Residual Sum of Squares	1.3248	Equation Log-likelihood	85.6869
Akaike Info. Criterion	77.6869	Schwarz Bayesian Criterion	66.9215
DW-statistic	.61633	ADF(4)	−6.2920

The corresponding numerical results are reported in Table VII, where the restricted estimates are also reported. According to our previous comments, we just present the results for the models that reveal useful to explain the main stylised facts; namely the intercept shift (*C*) and the regime shift (*C/S*) models. It is worth noting that the break occurs in 1989 while no evidence of any other shift is signalled at the beginning of the 1980s (while it is in the wage equation case). The price elasticity to labour costs is 0.67 till the break, and then increases by 0.43 after the break. The previous analysis was conducted by imposing a homogeneity restriction to the parameter estimates; this fact could bias the results if the restrictions are not

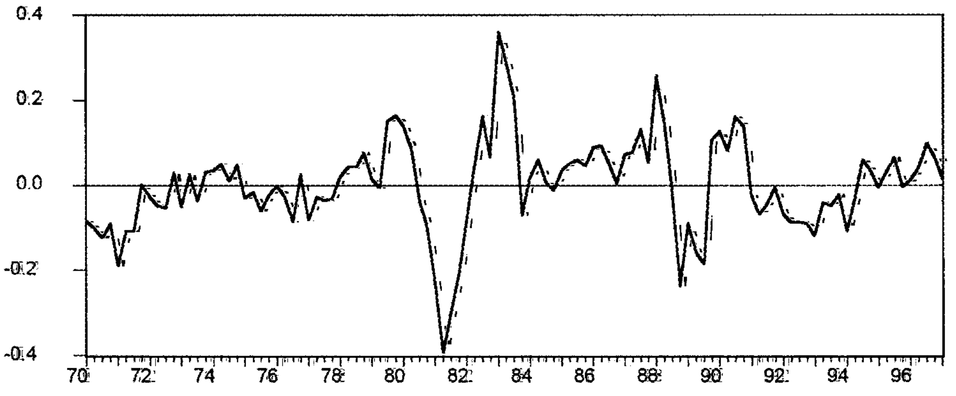

Figure 6. Residual plots of price equation (ecmpp), Poland.

data acceptable. In Table VIII we report the parameters estimates of a model with shifts in both the intercept and all other parameters. With respect to the results given in Table VII (Model *C/S* with the exchange rate as a proxy for the imported inflation effect) the model in Table VIII does not constrain the estimates after the shift. In this way we avoid that the price elasticity to the labour costs grows too much and consequently, because of the homogeneity restriction, the implicit elasticity to the exchange rate becomes negative.

The unconstrained parameter estimates in Table VIII allow for even more stationary residuals: the ADF statistic (bottom of Table VIII) reveals to be lower than the corresponding statistic in Table VII, and it is significant since the 5% corresponding critical value is −5.5. The consumer price elasticity to unit labour costs grows from slightly more than 0.6 to more than 0.9, while the elasticity to exchange rate moves from 0.33 to about 0.26. The fact that, after the regime shift, the unconstrained sum of wage costs and foreign elasticity is over one could suggest that, after the shock, the transitional high inflation phase is quite slowly absorbed by the system, influencing in this way a relevant part of the post 1989 sample. Perhaps a better estimate of the cointegrating relation after the shift would require additional data about the new regime, but this will be feasible only when future observations representing a stabilised inflation rate will be available. The residuals of the equation presented in Table VIII are plotted in Figure 6.

3.3. THE SHORT-RUN MODEL FOR POLAND

In the previous sections we discussed the results of modelling the cointegrated relationships in presence of possible structural shifts. With the aim of obtaining a valid representation of the long run, we allow the real per capita wage equation for an intercept shift in the third quarter of 1990 (see Table VI). On the other hand, as far as consumer price equation is concerned, the analysis leads to a complete regime shift, since the shock in the second quarter of 1989 gives cause for an

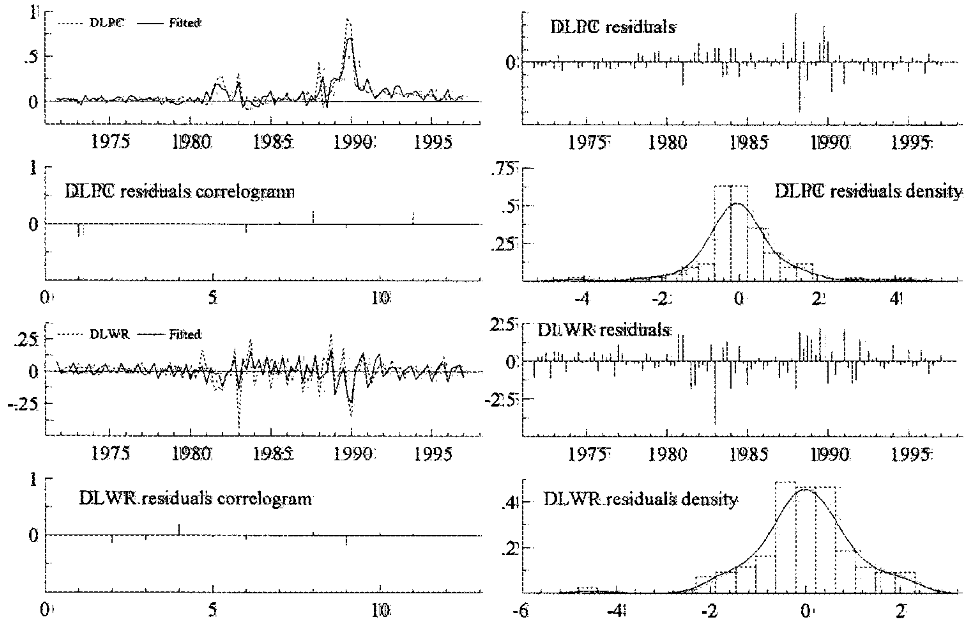

Figure 7. Price–wage model for Poland: the analysis of the residuals.

increase in the influence of domestic factors in explaining consumer prices (see Table VIII).

A complete analysis of the short-term aspects implies the specification of a vector equilibrium correction model (VECM) for the variables of interest,[9] the test fur weak exogeneity of the forcing variables and the system reduction to a more simplified structural model. A VAR model of order 6 is a reasonable point of departure of our analysis, as indicated by both the Akaike information criteria and the residuals diagnostics (not reported). According to this initial model the hypothesis of weak exogeneity of labour productivity, employment–population ratio and exchange rate in consumer price and real per capita wage equations are not rejected. Moreover we imposed two additional restrictions, namely the absence of the adjustment of short-run prices to the long-run wages (the parameter of *ecmwp* is set to zero in *dlpc* equation), and the absence of the adjustment of short-run wages to the long-run prices (the parameter of *ecmpp* is set to zero in *dlwr* equation); overall, we get a Likelihood-Ratio statistic $\chi^2[8]=14.7$, with a P-value=0.066.

These statistical characteristics enable us to simplify the analysis by avoiding to modelling a number of variables (labour productivity, employment–labour ratio and exchange rate) and by concentrating on a conditional two-equations model (prices and per capita real wage). After reducing the dimensionality of the parameter space by excluding the non-significant short term impulses, we obtain the structural model FIML estimates presented in Table IX; the over-identifying restrictions give a Likelihood-Ratio statistic $\chi^2[17]=22.9$, with a P-value=0.152. Residual diagnostics and parameter constancy tests are summarised in Figures 7 and

Table IX. Price–wage equation FIML estimates for Poland

Variable	Coefficient	Std.error	t-value	t-prob	HCSE[a]
Equation 1 for DLPC					
DLPC(−1)	0.98272	0.094669	10.381	0.0000	0.21225
DLPC(−2)	−0.28849	0.081623	−3.534	0.0006	0.15101
DLPC(−4)	0.21940	0.089832	2.442	0.0165	0.12143
DLWR(−1)	0.78141	0.12549	6.227	0.0000	0.22424
DLWR(−3)	0.18510	0.11141	1.661	0.1001	0.15074
DLWR(−4)	0.43431	0.12030	3.610	0.0005	0.15989
ECMPP(−4)	−0.20855	0.086563	−2.409	0.0180	0.11334
DLEX	0.22746	0.056617	4.018	0.0001	0.098013
Constant	−0.011458	0.020661	−0.555	0.5806	0.013655
Seasonal	0.0088418	0.028764	0.307	0.7592	–
Seasonal(−1)	−0.0032252	0.026988	−0.120	0.9051	–
Seasonal(−2)	−0.013343	0.028256	−0.472	0.6379	–
Equation 2 for DLWR					
Variable	Coefficient	Std.error	t-value	t-prob	HCSE[a]
DLPC(−1)	−0.35949	0.064821	−5.546	0.0000	0.075154
DLPC(−3)	0.28217	0.059969	4.705	0.0000	0.060041
DLPC(−5)	−0.19074	0.072351	−2.636	0.0099	0.13031
DLPC(−6)	0.19817	0.082330	2.407	0.0181	0.13617
DLWR(−1)	−0.44994	0.10147	−4.434	0.0000	0.092856
DLWR(−6)	0.27208	0.10452	2.603	0.0108	0.14468
ECMWP(−4)	−0.26224	0.073292	−3.578	0.0006	0.085420
Constant	0.056172	0.015244	3.685	0.0004	0.014512
Seasonal	−0.075789	0.021458	−3.532	0.0006	–
Seasonal(−1)	−0.059513	0.021776	−2.733	0.0075	–
Seasonal(−2)	−0.040692	0.020700	−1.966	0.0524	–

[a]Heteroschedastic consistent standard errors.

8. The results suggest an acceptable representation of the price–wage relationships for Poland in both long and short run. According to the Granger representation theorem, when cointegrating relationships are present, VAR model can be rewritten as a VECM, with the consequence that the long-run aspects of the system are identified with the *ecm* term. The long run of the system is hidden inside the *ecm* term and it was carefully discussed in the previous two paragraphs. As reported in Table IX, an exchange rate devaluation (a positive *dlex*) induces a short-term rise in the inflation rate.

The actual and fitted plots in Figure 7 remark the relevant explanatory power of the model proposed in Table IX, the residuals graphs show that some outliers

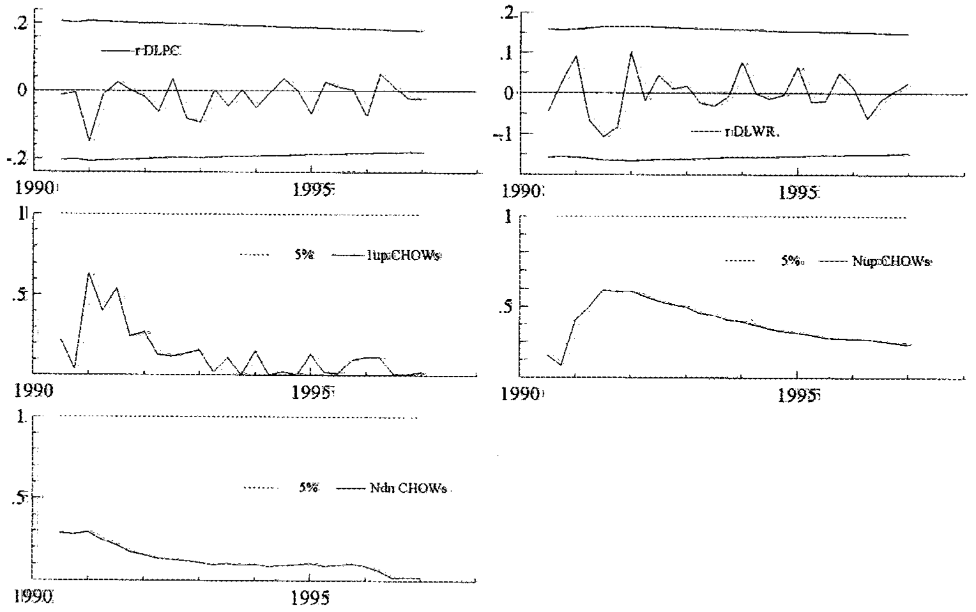

Figure 8. Price–wage model for Poland: 1-step residuals (first line) and Chow tests.

are clustered around the two periods of shock (at the beginning and at the end of the 1980s). The residual correlograms do not show any significant autocorrelation coefficient, the two histograms exhibit fat tails because of the outliers mentioned above.

Other useful information is depicted in Figure 8 where a number of graphical tests show a substantial parameter estimates stability after the shock.

A more specific issue of this analysis is to ascertain whether taking into account for breaks in the long run allows for a stable estimation of factor loading in each equation. As far as price equation is concerned, in Figure 9 both the recursive and the rolling OLS single equation estimates of the *ecm* coefficients are plotted. As it appears clearly from the plots, the break occurred at the beginning of 1980s reveals to have a much more important effect, on the coefficient stability, than the structural change that take place at the end of the 1980s. Moreover, the speed of adjustment of the price equation towards the cointegrated relationship after 1983 is rather stable, even in presence of relevant long-term structural breaks. During the second half of the 1970s (not plotted in Figure 9) the equilibrium relationship does not significantly influence (attract) the short-term movements. The rolling estimate, plotted in the lower area of Figure 9, confirms this stability over the period, but it coexists with a widening of the corresponding confidence interval since the shocks occurred during the 1980s induce bigger residuals (see also the residuals plots in Figure 7) and consequently less precise rolling estimates.

For what concerns the recursive and rolling estimates of the *ecm* variable in the wage equation the plots are reported in Figure 10. Though the estimate of factor loading seems reasonably stable since the second half of the 1980s, in this case the

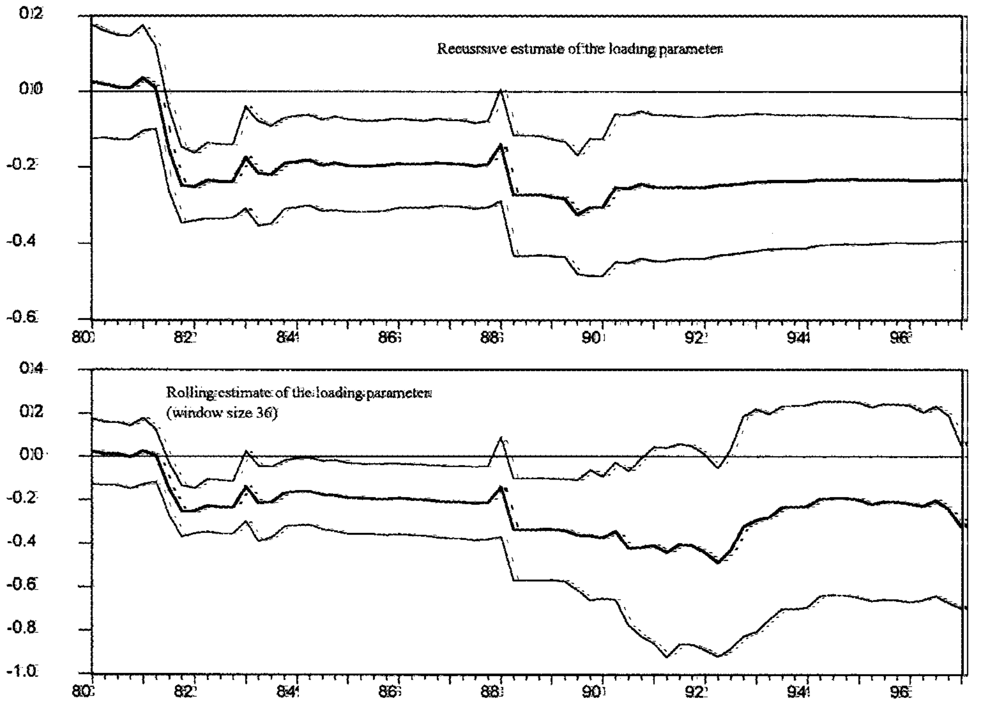

Figure 9. Price equation loading parameter estimates for Poland.

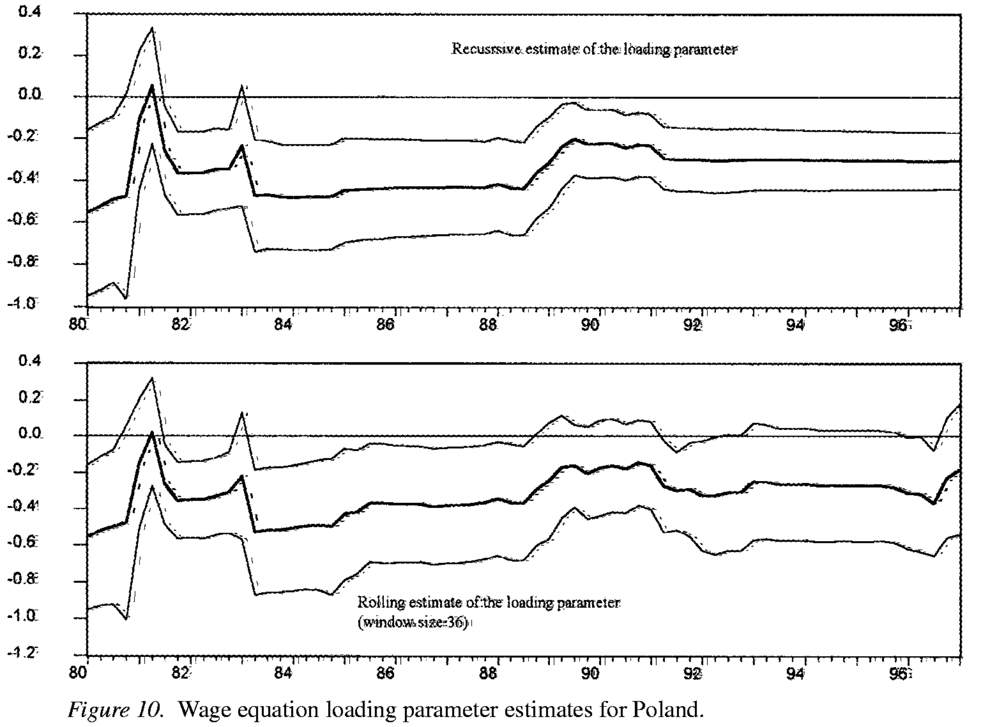

Figure 10. Wage equation loading parameter estimates for Poland.

Table X. Wage equation EG cointegration tests for Hungary

Test	*ADF*	*k*	*lpc*	*leta*
EG	−3.01	4	0.94	0.32
EG/T	−3.18	4	0.87	0.03

break occurrence in the first part of the 1980s produces a more relevant effect in terms of parameter instability. Moreover the stability of the wage equation loading factor after the 1985 estimate is associated with a precision that seems not heavily affected by the different shocks, as it were for prices.

The general evidence is that the two breaks (at the beginning and at the end of the 1980s) reflect a different structure because they exert different impacts on the short-run behaviours. More precisely, the end of the 1980s change exerts its effects on the long-run *equilibria* (through the *ecm* terms), while the beginning of the 1980s break seems to influence the short-run parameters and, among these, the speed of adjustment. This finding appears to be robust since it holds both in price and in wages equations, with recursive and rolling estimates.

After all, for both equations the beginning of the 1980s corresponds to a limited period of instability in the loading parameters but, while the wage equation loading parameter has a tendency to return at levels similar to those attained before the shock, a permanent shift (corresponding to an increase in the speed of adjustment) appears to characterise the price loading factor.

3.4. THE COINTEGRATION RELATIONSHIPS WITH REGIME SHIFTS FOR HUNGARY

In this section we iterate the analysis carried on for Poland in order to investigate the main characteristics of the long-run per capita nominal wage and prices relationships in Hungary. The wage equation is:[10]

$$lw_t = \alpha_1 lpc_t + \alpha_2 leta_t + u_t \tag{7}$$

where the parameters point estimates are found to be respectively 0.658 and 0.92 (see Golinelli and Orsi, 1998, Table 8). We start our analysis with the classical EG cointegration test, on the basis of variables appearing in Equation (7), and Table X reports the main findings. The outcomes clearly show that the assumption of cointegration is not supported by data, and this is true with or without the trend. With respect to previous results (Golinelli and Orsi, 1998, Table 8), Table X exhibits effects that are generally higher for prices and lower for labour productivity; these outcomes would not change if we include *lnpop* amongst the cointegrating variables (results not reported).

Then we use the procedure suggested by Gregory and Hansen for detecting the possible presence of significant breaks over the whole sample. At first, we specify a

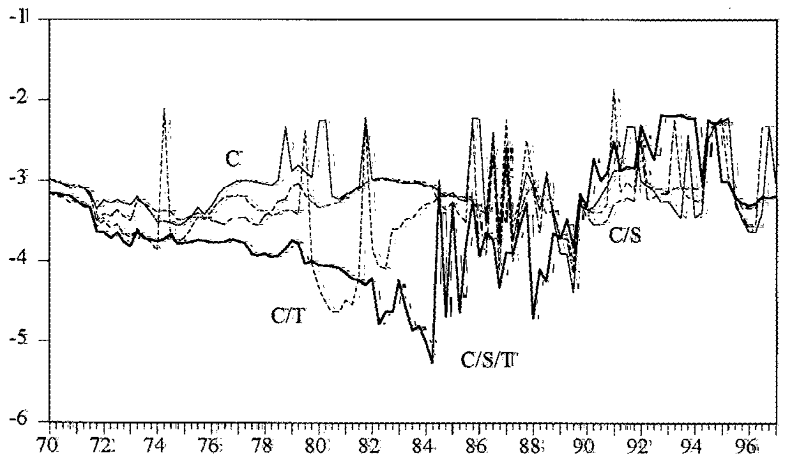

Figure 11. Regime shifts ADF plots for per capital wage equation in Hungary.

long-run and short-run model able to endogenise the expected structural shifts that occurred in the prices and wages relationships at the end of the 1980s. The single values of the GH test statistics for the wage equation are plotted in Figure 11; more precisely, the graph reports the GH statistics for the wage model as specified in Equation (7).

As we have seen before, Polish data are characterised by the presence of two evident and significant breaks at the beginning and at the end of the 1980s respectively. On the other side, in the case of Hungary we note a single and less pronounced break in the middle of the 1980s; as a matter of fact the univariate Perron tests never find strong evidence against the null. The GH cointegration test, in presence of break, for wage Equation (7) shows values that are not significant, In a separate analysis (not reported), the addition of the *lnpop* variable to Equation (7) provides economically meaningless estimates of the broken parameters. Divergent results for different models under the alternative or meaningless estimates suggest possible misspecification (omission of relevant variables).

A hypothetical explanation of these apparently negative results is in the first lines of Figures 2a and 2b where the levels of the four variables of interest are plotted. As already noted, each Hungarian variable exhibits just one break, often in different times: the nominal variables (Figure 2a) show a break in the second half of the 1980s, while the real ones (Figure 2b) are affected by a shock at the beginning of 1990s. The absence of synchronisation of the breaks could suggest that if we allow for a complete regime shift in the same point in time, the spurious interrelation amongst variables with different breaking times can lead to economically meaningless results.

As far as the price equation is concerned, we apply the GH tests to the long-run Hungarian consumer price equation, explained by the following structural relationship:[11]

$$lpc_t = \alpha_1(lw_t - leta_t) + (1 - \alpha_1)(lex_t + lpm59_t) + u_t \qquad (8)$$

Table XI. Price equation cointegration tests for Hungary

Test	*ADF*	*k*	T_b	α_1 estimate in eq. (8)	α_2 variation after T_b
EG	−0.04	1			
C	−5.08*	0	1991 4th	0.49	
C/S	−6.31**	0	1992 1st	0.50	−0.42

* and ** respectively 5% and 1% significant.

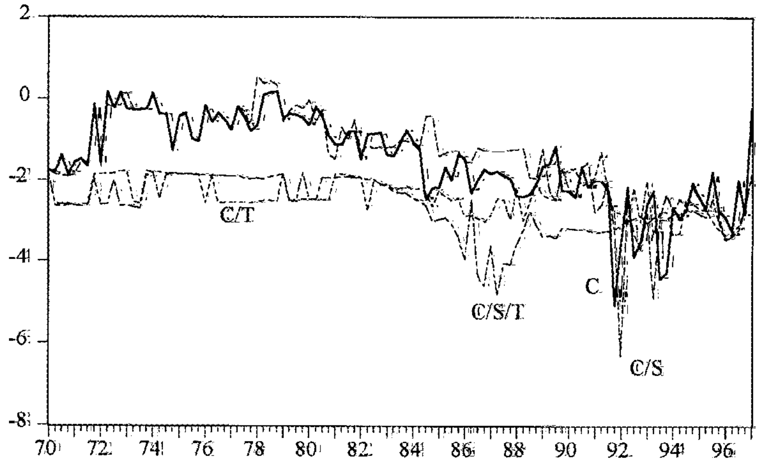

Figure 12. Regime shifts ADF tests for long-run price equation in Hungary.

where the parameter restricted estimate is equal to 0.382 (see Golinelli and Orsi, 1998, Table 8).

The results are depicted in Figure 12. From a visual inspection of the plots, it appears that the breaking cointegration hypothesis is more acceptable for models of type *C* and *C/S*. The evidence reported in Figure 12, along with the underlying theoretical economic guideline, is in favour of the estimation results of Equation (8) with respect to models *C* and *C/S*. These results along with the EG cointegration test (without the linear trend) are reported in Table XI. Moreover it seems important to draw attention on a number of important insights. Firstly, the breaking models are statistically significant while the EG tests are not: the detected break reveals essential in explaining a stationary cointegrated relationship, and the periods where the different models break are quite similar (the last quarter of 1991 or the first of 1992). For this reason we expect the breaks occur at the same time for all the variables of interest, while in the Hungarian wage equation it was not the case. Secondly, given the information about the existence of a breaking cointegration relationship, we must choose between Model *C* and Model *C/S* results. The tests are significant in both models, but it is to be noted that in Model *C/S* the estimated

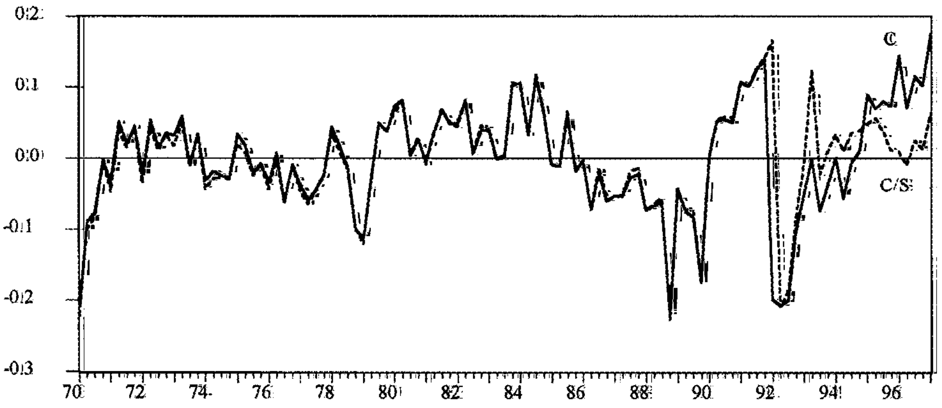

Figure 13. Error correction terms for Model *C* and Model *C/S* in Hungary.

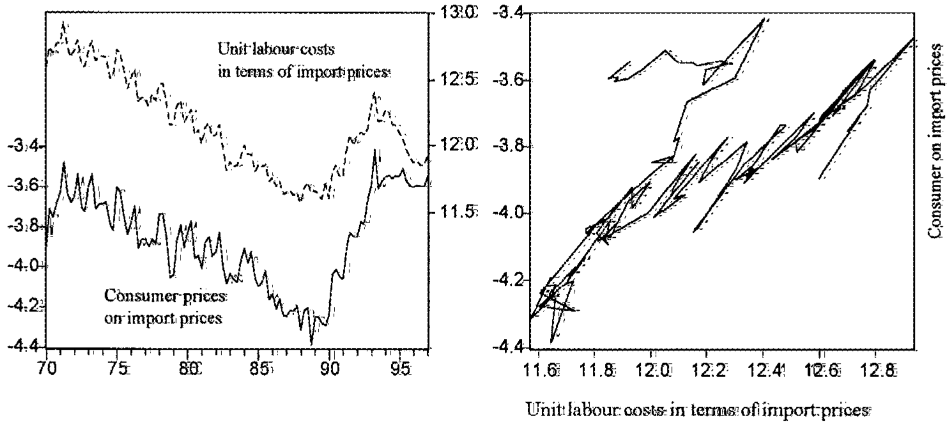

Figure 14. Consumer prices and unit labour costs in terms of import prices (Hungary).

parameters depicts a shift with an unexpected (and not theoretically acceptable) strong increase in foreign contribution to domestic prices; in addition we must emphasise that the statistical information after the break is very low.

As a consequence, the error correction terms (the residuals of the breaking cointegration models) in Figure 13 follow the same path till 1992, because the elasticity of prices to labour costs is almost the same in all models (see also Table XI). The *ecm* terms tend to differ markedly after 1994, because the lower long-run elasticity to domestic inflation sources (Model *C/S*) keeps near to equilibrium the corresponding error correction term.

In order to clarify ideas, Figure 14 reports the plots of both the variables of interest (on the left) and the scatter diagram of prices on unit labour cost, both in terms of import prices (on the right). From the plots on the left in Figure 14, we note that both the variables show similar path till the end of the 1980s, since then real consumer prices grow faster than real unit labour costs; this fact *per se* explains the break evidenced by the tests at the beginning of the 1990s. In the scatter diagram

Table XII. Price long-run equation for Hungary

Dependent variable is $(lpc - lex - lpm59)$
109 observations used for estimation from 1970Q1 to 1997Q1

Regressor	Coefficient	Standard Error	T-Ratio[Prob]
CONST	−9.8824	.25573	−38.6445[.000]
lw-leta-lex-lpm59	.48791	.021079	23.1472[.000]
S1	.029783	.021400	1.3918[.167]
S2	.050367	.021688	2.3224[.022]
S3	−.1504E-3	.021660	−.0069415[.994]
D914	.36979	.019348	19.1130[.000]

R-Squared	.89130	R-Bar-Squared	.88603
S.E. of Regression	.079221	F-stat. F(5, 103)	168.9180[.000]
Mean of Dependent Variable	−3.8602	S.D. of Dependent Variable	.23466
Residual Sum of Squares	.64643	Equation Log-likelihood	124.7921
Akaike Info. Criterion	118.7921	Schwarz Bayesian Criterion	110.7180
DW-statistic	.70072	ADF(0)	−5.0829

S1, S2, S3 seasonal dummies

on the right it is evident that, about at the same time of the break, the long-run relationship shows a clear upward shift in the constant, while, because of a lack of information, it is difficult to understand if also the slope changes after the break occurred.

Given the graphical evidence provided by Figure 14 and the more theoretical acceptability of Model *C* estimates in Table XI, in what follows we first analyse the Model *C* long-run results, as reported in Table XII, and subsequently we try to model the short-run adjustment movements. By comparing the results in Table XII with those presented in Golinelli and Orsi (1998), we observe that the inclusion of a significant break in the intercept of the model slightly increases the elasticity of prices to labour costs, from about 0.4 to 0.5.

The short-run analysis starts from a VECM model of order 6 for the variables of interest (prices, wages, exchange rate and labour productivity), conditional to the exogenous import prices in dollars variable and lagged the error correction term from the long-run estimates proposed in Table XII.

The residual tests do not reject the hypothesis of vector normal white noise errors, and the weak exogeneity assumption for wages, exchange rate and labour productivity in the price equation is not rejected (the corresponding Likelihood-Ratio statistic $\chi^2[3]$ is equal to 5.09, with P-value=0.165).

The results of the complete dynamic error correction model are reported in Table 13. The final form was obtained through a progressive parameter space reduction, starting from an autoregressive distributed lag model of order 6 (as

Table XIII. Price short-run equation for Hungary

Dependent variable is DLPC
104 observations used for estimation from 1971Q2 to 1997Q1

Regressor	Coefficient	Standard Error	T-Ratio[Prob]
CONST	.0050372	.0021706	2.3206[.022]
S1	.015474	.0054257	2.8520[.005]
DLPC(−4)	.67972	.084912	8.0050[.000]
DLEX	.037345	.019362	1.9288[.057]
ECMPH(−4)	−.049279	.021693	−2.2717[.025]

R-Squared	.67356	R-Bar-Squared	.66037
S.E. of Regression	.018438	F-stat. F(4, 99)	51.0684[.000]
Mean of Dependent Variable	.027781	S.D. of Dependent Variable	.031638
Residual Sum of Squares	.033656	Equation Log-likelihood	270.3008
Akaike Info. Criterion	265.3008	Schwarz Bayesian Criterion	258.6898
DW-statistic	1.8396		

in Poland), augmented by three seasonal dummies. The standard errors of parameter estimates are computed by using the White heteroscedasticity consistent covariance estimator because of potential heteroscedasticity due to the presence of outliers.

As also shown in Table XIII, the speed of adjustment of the price equation to the long-run broken cointegrating relationship is very slow and could be affected by parameter instability, probably due to an omitted effect of the wage equation *disequilibria*. This aspect is investigated by the recursive and rolling (with window size equal to 36) estimates of the loading parameter in the price equation. The corresponding graphs, presented in Figure 15, reveal a probable misspecification of the short-term model for prices.

4. Conclusions

In this paper we discuss and apply testing procedures useful for detecting structural changes in a cointegration system when these shifts occur at unknown periods. As a general comment we recommend some caution in the interpretation of a cointegration relationship when the cointegrating vector is allowed to change over time as a result of changes in the economic system under scrutiny, since these events have effects on the size of the cointegration test by affecting the power of the test.

Our analysis is illustrated using first the standard ADF and PP statistics, respectively proposed by Dickey–Fuller and Phillips–Perron, then the Perron testing procedure that is helpful for carrying out the analysis of stationarity for each variable along with testing for the occurrence of structural changes. We argued that

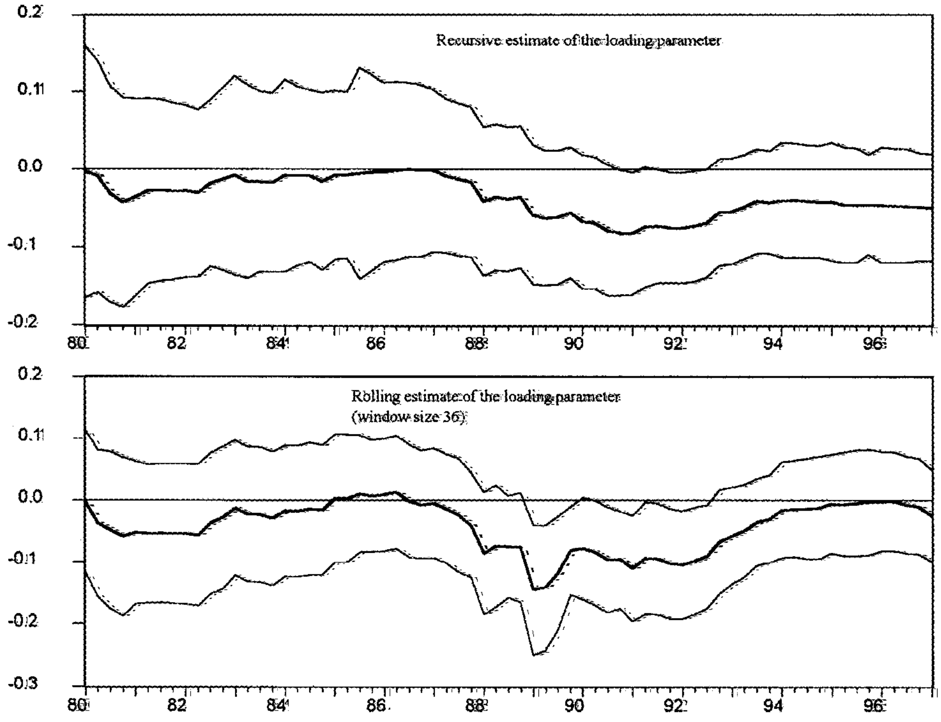

Figure 15. Price equation loading parameter estimates for Hungary.

although univariate representations of data are characterised by the presence of some breaks in different periods, the explicit consideration of these breaks in the stationarity analysis does not succeed in eliminating the integratedness of order one which characterises all the variables considered in our analysis. These results are important and, among other things, permit to ascertain that a broken-trend model for these variables is not fruitful for disentangle non-stationarity, except for the real wages for Poland. Moreover, these testing procedures are also useful for identifying the period of occurrence of a significant structural change.

In a multivariate framework, the cointegration analysis has been accomplished under the possibility of a break at an unknown period, and the existence of linear and stable long–run relationships between wages and prices was not rejected for Poland. In the short run a complete VAR model is reparametrized as a VECM, and the loading parameter estimates for the *ecm* terms are analysed over the sample. We use the testing procedure proposed by Gregory and Hansen, which can be view as a multivariate extension of the Zivot and Andrews test, and our empirical findings reveal significant evidence in favour of time varying cointegration relationship between wages and prices for Poland, when the long-run parameters are allowed to change between two different regimes. This cointegration relationship are not significant for Hungary, even if we found a significant evidence of a break around 1984–1985. Incidentally, the breaks affecting stationarity and the one affecting cointegration are not necessarily the same. These results appear robust against alternative model specifications.

These findings are meaningful and a conclusion we can reach is that the breaks are important for these countries, but they affect mainly the variable rate of growth rather than its non-stationarity. As a practical consequence we can say that the structural break analysis is an important component both for the non-stationarity analysis and cointegration, but it is not sufficient for providing a complete explanation of them. Correspondingly in the cointegration analysis the testing procedure used must allow for the possibility of significant structural breaks in the cointegration relationship, since they may affect the long-run equilibrium conditions over the sample.

Appendix

A1. THE BASIC DATA SOURCES

As far as Poland data are concerned, Table A1 lists the variables object of the study; note that all the variables are not seasonally adjusted. The main data source is the database from *The Central and East European Economic Research Center* at Warsaw University; in addition, another database is currently available from OECD.[12] At the first stage of the analysis we made a number of checks in order to assess the compatibility of the two sources; the results were quite encouraging: in the most part of the cases, data are coherent.

The first column of Table AI lists the labels of the variables that are used in subsequent tables, in the second column there is the *legenda* of each variable and the third column specifies some particular notes of interest about the link of old and new data from national sources. In fact, the database for Poland was obtained by merging two data sources: the old information was from a 1970–1992 database,[13] the new one is from 1990–1997 LAM Model for Poland database[14] Where we do not mention any note, old and new data were perfectly compatible.

In the first row of Table AI, the industrial production data are from OECD for the period 1985–1996, the latter quarters (from the end of 1996) are updated by using LAM Model industrial production levels at constant prices. For all other variables in Table AI, the statistical information from OECD source was coherent with the existing information about LAM database.

The transformations imposed to the basic data in Table AI in order to obtain the final series to be used in the econometric analysis, can be summarised as follows. When we take the logarithms of a variable, we add an 'L' top of the corresponding variable label; for example, *lpc* means the logarithm of the consumer prices (PC). In general, in the paper the definition of the variables is not case sensitive. The result of the first difference transformation of a variable is labelled by a name starting with a 'D'; for example, $dipc_t = lpc_{t-1}$.

The same information about Hungary is reported in Table AII. As far as the listed Hungarian variables are concerned, given a lack of basic information about the

Table AI. The basic data for Poland

Label	Description	Notes
Y	Industrial production index (1985=1)	1985–1996 IIP from OECD database, 1996–1997 industrial production (X in LAM Model for Poland database)
N	Employment in national economy (thousands) L in LAM Model for Poland database	1970–1990 old average employment in State sector (in 1991 it is equal to L)
W	Gross monthly wages in national economy (Plz), W in LAM Model for Poland database	1970–1990 old average monthly wage in State sector (in 1991 it equals W)
PC	Consumer price index (1985=1) P in LAM Model for Poland database	
POP	Population (thousands)	Old quarterly data were interpolated by IMF, *International Financial Stats*
EX	US dollar market exchange rate against Plz D in LAM Model for Poland database	

1970s and the 1980s, the wage and employment data are referred to the industrial sector, instead of the whole economy. The data sources are basically the same as for Poland. The coherence between the alternative sources was successfully verified; in fact, apart for the industrial production index variable, the differences between levels were less than 3%.

As for the Poland case, a relevant subperiod (from 1980 to 1996) of industrial production index data was based on the OECD source; the LAM Model database source (production in levels) was just used for updating the latter quarters of the sample, starting from the end of 1996.

The rules followed in setting the label names of variables after transformation in the Hungarian case are exactly the same as for the Polish case.

A2. THE ADF AND PP UNIT ROOTS TESTS

The full results output of the ADF and PP testing procedures whose results are analysed in the paper are presented in Tables AIII and AIV, respectively for Poland and Hungary.

In understanding the empirical results reported in tables below it is important to remember that the Dickey–Fuller test is based on the hypothesis that the variable is generated by a simple AR(1) process; for this reason, the ADE and PP statistics will be used for testing the presence of a unit root in a higher order process (in general, of order $p > 1$). The ADF testing procedure makes a parametric correction for higher order AR processes by adding some lags of the dependent variable in the equation model.

Table AII. The basic data for Hungary

	Label and description	Notes
Y	Industrial production index (l985=1) X in LAM Model for Hungary database	1970–1979 industrial production index from the old database, 1980–1996 OECD data, from 1996 X variable
N	Employment in industry (thousands) L in LAM Model for Hungary database	1970–1990 employment in State sector
W	Gross monthly wages in industiy (Ft) W in LAM Model for Hungary database	1993–1997 gross monthly wages in national economy (variable W in LAM Model for Hungary database)
PC	Consumer price index (1985=1) P in LAM Model for Hungary database	
POP	Population (thousands)	Old quarterly data were interpolated by IMF, *International Financial Stats*
EX	Us dollar market exchange rate against Ft D in LAM Model for Hungary database	1970–1990 market exchange rate from Pick's *Currency Yearbook*
PM59	International manufacturing products in $	OECD-SITC codes from 5 to 9

As it is well known, the ADF testing procedure results are heavily influenced by two practical issues, namely: (a) the specification of the deterministic components in the equation and (b) the number of lagged differences to be added in the model, the so called order of the ADF test. As far as problem (a) is concerned, we choose to include both constant and trend in the equation model when the tested variable shows a trending path (mainly testing levels), while tests on first difference variables were based on a model only including the constant term. Hamilton (1994, p. 501) suggests in principle to choose a specification that is a plausible description of the data under both the null and the alternative; for this reason we performed the test both with and without seasonal dummies since we work on seasonally unadjusted series. Problem (b) was tackled by adding lagged terms in the equation model till the higher lag had a 10% significant parameter on the basis of the normal distribution critical values.

In alternative to the ADF parametric approach, Phillips and Perron (1988) suggests a non-parametric approach to control for higher order serial correlation problem; their proposal is labelled as the PP test. By considering the simple AR(1) model, the PP test makes a correction to the usual t statistic in order to take into account for possible serial correlation in the error term. The correction is nonparametric since the PP test uses an estimate of the spectrum of the error process at frequency zero that is robust to both heteroskedasticity and autocorrelation of unknown form. The Newey and West (1987) procedure for adjusting the standard errors is often used in the applied work, since it constitutes an easy approach to

Table AIII. Univariate unit root tests for the Polish variables

Label	Model	Statistic	ADF test lag	PP test statistic	5% critical value
lpc	c, t	−1.96	1	−1.75	−3.45
lpc	c, t, s	−1.98	1		−3.45
dlpc	c	−4.57*	0	−4.54*	−2.89
dlpc	c, s	−4.28*	0		−2.89
lw	c, t	−1.81	5	−1.52	−3.45
lw	c, t, s	−2.11	4		−3.45
dlw	c	−2.87	4	−5.78*	−2.89
dlw	c, s	−2.77	4		−2.89
lwr	c, t	−3.43	4	−3.99*	−3.45
lwr	c, t, s	−3.44	0		−3.45
lwr	c	−2.47	5	−3.29*	−2.89
lwr	c, s	−2.91*	0		−2.89
dlwr	c	−5.53*	4	−12.76*	−2.89
dlwr	c, s	−7.22*	2		−2.89
leta	c, t	−2.35	4	−2.45	−3.45
leta	c, t, s	−2.24	0		−3.45
dleta	c	−4.59*	3	−11.88*	−2.89
dleta	c, s	−6.94*	2		−2.89
lnpop	c, t	−2.63	5	−2.41	−3.45
lnpop	c, t, s	−2.53	5		−3.45
dlnpop	c	−2.15	4	−9.69*	−2.89
dlnpop	c, s	−2.08			−2.89
lex	c, t	−2.49	3	−2.51	−3.45
dlex	c	−4.23*	2	−9.88*	−2.89

*5% significant test; c=constant, t=trend; s=seasonals.

get consistent estimates of the parameters covariance matrix in the presence of both autocorrelation and heteroskedasticity of the errors. The distribution of the PP statistic is the same as the ADF statistic. As for the ADE test, the implementation of the PP test also rests on some important choices: (a) again, the specification of the deterministic components of the test in the regression model used; (b) the truncation point q for the Newey–West correction, i.e. the number of periods of serial correlation to include. As far as the problem (a) is concerned, we acted as in the previous ADF case. Following the suggestion in Newey–West (1987), the

Table AIV. Univariate unit root tests for the Hungarian variables

Label	Model	Statistic	ADF test lag	PP test statistic	5% critical value
lpc	c, t	−0.61	4	−0.79	−3.45
lpc	c, t, s	−0.55	4		−3.45
dlpc	c	−1.39	3	−9.09*	−2.89
dlpc	c, s	−1.41	3		−2.89
lw	c, t	0.64	4	−0.85	−3.45
lw	c, t, s	0.53	4		−3.45
dlw	c	−1.73	5	−18.71*	−2.89
dlw	c, s	−1.72	5		−2.89
lwr	c, t	−2.66	4	−8.65*	−3.45
lwr	c, t, s	−2.73	4		−3.45
lwr	c	−2.75	4	−7.50*	−2.89
lwr	c, s	−2.71	4		−2.89
dlwr	c	−5.05*	3	−24.89*	−2.89
dlwr	c, s	−4.72*	3		−2.89
leta	c, t	−2.08	5	−6.96*	−3.45
leta	c, t, s	−1.90	4		−3.45
dleta	c	−3.89*	4	−30.11*	−2.89
dleta	c, s	−4.88*	3		−2.89
lnpop	c, t	−2.12	4	−1.25	−3.45
Inpop	c, t, s	−2.05	4		−3.45
dlnpop	c	−2.57	3	−8.99*	−2.89
dlnpop	c, s	−2.55	3		−2.89
lex	c, t	−1.51	2	−1.81	−3.45
dlex	c	−9.66*	1	−11.53*	−2.89
lpm59	c, t	−2.13	3	−1.60	−3.45

*5% significant test; c=constant, t=trend; s=seasonals.

problem (b) was tackled by using a truncation point selection based on the largest integer not exceeding $4(T/100)^{2/9}$; in this way, our choice (q=4) only depends on the available number of observation T in the sample.

Notes

[1]The research was undertaken with support from the European Union's Phare ACE Programme 1996. We are much indebted to Prof. Krystyna Strzala of the University of Gdansk for providing us with data and useful information, participants at the Leicester Conference *Econometric Inference into the Macroeconomic Dynamics of East European Economies*, Leicester, June 20–21, 1998 and Lucio Picci for helpful comments.

[2]See Dickey and Fuller (1979) and Phillips and Perron (1988). The results of these tests are reported in Appendix A2.

[3]In more detail, Perron introduces three different models, labelled Model 1, Model 2 and Model 3. Models 1 and 2 are *innovational outlier* models, where the change occurs gradually. These two models can easily be obtained from the classical ADF(k) test equation, allowing for a time varying deterministic component that, in Model 1, is defined as a constant linear trend and a change in intercept at time T_b, while in Model 2 is defined as a change in both the intercept and the slope of the linear trend. In other terms, Model 1 states that a variable is defined as a trended variable with a one time change in the intercept, while Model 2 is a trended variable with both a change in intercept and the slope of the trend. Model 3 considers an *additive outlier* test, where the break only occurs in the trend parameter (and not in the intercept).

[4]For what concerns the change point T_b, we can relay on our a priori knowledge of the period of occurrence of the break, but as shown by many authors, Zivot and Andrews (1992) among others, a test of the presence of a unit root versus a structural change conducted conditional on the a priori known change point may cause a serious size distortion of the test. In any case they suggest assuming that the change data T_b is unknown.

[5]It is important to stress that the proposed Perron test has an important drawback: it allows for a unique break during the whole period, while, at least for the Polish case, the graphical inspection of variables would suggests two distinct breaks, the first at the beginning and the second at the end of the 198Os. A possible solution is foreseen according to the results provided by Rybinsky (1997), where the structural breaks are supposed to depict a temporary shift of the mean of the series: one at the beginning and one at the end of the transition process.

[6]This analysis has been performed on raw data, since, as we noted before, seasonal adjustment entails smoothing of data. In particular the widely used Census X-11 seasonal adjustment programs have undesirable effects on procedures involving structural breaks and variable level shifts.

[7]The parameter point estimates were respectively 0.506 and 0.881. The *legenda* is in Appendix A1.

[8]The *legenda* of the variables is in Appendix A1.

[9]Namely: prices, per capita real wage, labour productivity, employment–labour ratio and exchange rate.

[10]The *legenda* of the variables is in Appendix A1.

[11]The *legenda* of the variables is in Appendix A1.

[12]OECD, *Short-term indicators for Central and Eastern Europe*, in Statistical Compendium.

[13]The old data consist in a fist set of variables that were made available at the end of 1993.

[14]A description of the LAM model is presented in Charemza (1994).

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