PRICE-WAGE DYNAMICS IN A TRANSITION ECONOMY: THE CASE OF POLAND

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PRICE-WAGE DYNAMICS IN A TRANSITION ECONOMY:
The case of Poland[1]

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Abstract. In this paper we analyse the wage price relationship of an economy in transition characterized by important structural changes. It is known (see Perron 1989) that structural breaks in stationary time series can induce apparent unit roots. The stationarity analysis of the series employed in the present model is conducted jointly with the assumption that the breakpoint location is unknown. We follow a testing procedure recently proposed by Zivot and Andrews (1992). Cointegration analysis of wages and prices in presence of structural breaks permits to find empirical evidence in favour of two cointegrating vectors involving prices and wages. Our analysis faces on the different structural behaviour of price-wage dynamic relationship in the short and long run; we also enucleate the relative importance of import prices as a source of wage-price short run fluctuations.

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1. Introduction

In the applied economic literature wage-price relationships have been extensively studied for the western economies (see Monfort and Rabemanjara (1990), Blanchard (1986) among others) and, more recently, also for eastern economies (see Blangiewicz and Bolt (1992), Commander and Coricelli (1991), Welfe (1991) among others). A major problem encountered in these studies concerns the non-stationarity of wages and prices series, with the complication that, for eastern economies, this evolutionary course is frequently accompanied with one or more structural breaks. The most important of these breaks corresponds, roughly, to the structural economic changes, occurred towards the beginning of nineties, which characterize the social and economic evolution of the eastern economies.

On the other hand, it is well known that one of the main features which characterizes the central planned economies is the almost complete absence of inflation, due to the fact that prices were determined administratively. The transition towards a market based economy has translated repressed inflation into an open one, and the resulting increase in the inflation rate often has turned to be persistent.

The aims of this paper are to investigate the dynamic relationships of prices and wages in presence of such important structural changes, and to try to discover cointegration relationships in the presence of structural breaks. Testing for cointegration has become an important feature of the empirical analysis of economic time series. Various tests have been proposed and largely used in empirical analysis, but most of their theoretical properties rest on the assumption of unit root processes in the absence of structural breaks. It is well known that structural breaks in stationary time series induce apparent unit roots, as shown by Perron (1989), with the consequence that tests of unit roots have low power when used for series with structural breaks.

The economic time series of eastern economies, over a period which contains the transition phase moving from a centrally planned economy towards a market based economy, typically show peaks and changes that hint at the presence of regime shifts and structural breaks. Any empirical analysis which make use of such series must explicitly consider these changes which become a bench-mark for any model aiming to propose an explanation for a transition economy.

The simultaneous treatment of these two aspects, namely stationarity and structural changes, is rather complicate and only recently suitable testing procedures have been developed. We test the unit root hypothesis against an alternative of trend stationarity with a break-point that occurs somewhere in the period. The location of such break-point is not fixed exogenously as in Perron (1989), but rather is estimated according to the testing procedure proposed by Zivot and Andrews (1992).

The model we propose incorporates, where feasible, these important specific features providing a specification which consider separately short run dynamics from the long run. Our analysis pays attention to the price-wage spiral emphasizing the role of exchange rate and import prices in determining domestic inflation. In fact, price liberalization measures progressively reduced the relative weight of administered prices, with the consequence that an important linkage between domestic and foreign prices was established and imported inflation become progressively more important.

In section 2 we present the stationarity analysis which is performed jointly with the structural change analysis in order to avoid spurious unit roots. In section 3 the theoretical model is sketched while section 4 is devoted to cointegration analysis. At first we analyse the two subsets of prices and wages separately; subsequently a global model is presented and the cointegration analysis is extended to the whole system. Finally some concluding remarks can be found in section 5.

2. Stationarity and structural change analysis

Perron (1989) showed, both analytically and empirically, that structural breaks in stationary time series can induce apparent unit roots. We want to consider here the problem of testing for the presence of a unit root in a series, against the alternative of stationarity, in the presence of a structural change that occurs at some unknown point. In a recent paper, Perron (1990) proposed a procedure for testing the hypothesis of a unit root with drift jointly with an exogenous structural break that occurs at any time T_b, with $1 < T_b < T$, against the alternative that the series is stationary about a deterministic trend with an exogenous break in the trend function at time T_b.

We consider here three alternative parameterizations of the structural break, and consequently the unit-root null hypotheses are written in three different ways, *i.e.*:

Model (1): $y_t = \mu + aD(T_b)_t + y_{t-1} + e_t$

Model (2): $y_t = \mu_1 + y_{t-1} + (\mu_2 - \mu_1)DU_t + e_t$

Model (3): $y_t = \mu_1 + y_{t-1} + aD(T_b)_t + (\mu_2 - \mu_1)DU_t + e_t$

where:

$D(T_b) = 1$ if $t = T_b + 1$

$= 0$ otherwise

$DU_t = 1 \quad \text{if} \quad t > T_b$

$\quad = 0 \quad \text{otherwise}$

As we can see, Model 1 permits an exogenous change in the levels, Model 2 allows for an exogenous change in the rate of growth, while Model 3 admits both changes.

The alternative trend stationary hypotheses that are implicitly considered are:

Model (1): $y_t = \mu_1 + \beta t + (\mu_2 - \mu_1) DU_t + e_t$

Model (2): $y_t = \mu + \beta_1 t + (\beta_2 - \beta_1) DT_t^* + e_t$

Model (3): $y_t = \mu + \beta_1 t + (\mu_2 - \mu_1) DU_t + (\beta_2 - \beta_1) DT_t^* + e_t$

where:

$DT_t^* = t - T_b \quad \text{if} \quad t > T_b$

$\quad = 0 \quad \text{otherwise}$

Perron calls Model 1, which allows for a one-time change in the level, the *crash* model; the difference $(\mu_2 - \mu_1)$ measures the magnitude of the change in the intercept of the trend which occurs at time T_b. Similarly Perron calls Model 2 as *changing growth* model, where the difference $(\beta_2 - \beta_1)$ represents the importance of the change in the slope of the trend function occurring at time T_b. Finally Model 3 considers changes both in the level and in the slope.

The usual Dickey-Fuller test, in its augmented version (ADF), can be used to test the unit root hypothesis in models 1, 2, and 3, which involves the following augmented regression equations:

$$\textit{Model}\ (1):\quad y_t = \mu^{(1)} + \theta^{(1)} DU_t + \beta^{(1)} t + d^{(1)} D(T_b)_t + \alpha^{(1)} y_{t-1} + \sum_{j=1}^{k} c_j^{(1)} \Delta y_{t-j} + e_t \tag{1}$$

$$\textit{Model}\ (2):\quad y_t = \mu^{(2)} + \beta^{(2)} t + \gamma^{(2)} DT_t^* + \alpha^{(2)} y_{t-1} + \sum_{j=1}^{k} c_j^{(2)} \Delta y_{t-j} + e_t \tag{2}$$

$$\textit{Model}\ (3):\quad y_t = \mu^{(3)} + \theta^{(3)} DU_t + \beta^{(3)} t + \gamma^{(3)} DT_t^* + d^{(3)} D(T_b)_t + \alpha^{(3)} y_{t-1} + \sum_{j=1}^{k} c_j^{(3)} \Delta y_{t-j} + e_t \tag{3}$$

As usual, the number k of additional regressors is chosed on the basis of a test of the significance of the estimated coefficients $c_j^{(i)}$ for $i = 1, 2, 3$.

Testing for the presence of a unit root, according to Perron, is done considering the following statistics in models (1) - (3):

$$t_{\alpha^{(i)}}(\lambda) \qquad i = 1, 2, 3. \tag{4}$$

i.e. the standard t statistics for testing $\alpha^{(i)} = 1$ where $\lambda = T_b/T$ represents the location of the break point over the sample. Perron's null hypotheses take the break points exogenous and removes these periods from the noise functions of the series considered in the analysis.

In our analysis we treat the break points as endogenous and consequently the null hypothesis for the three models is represented by an integrated series without structural break *i.e.:*

$$y_t = \mu + y_{t-1} + e_t \tag{6}$$

Under the alternative hypothesis the series y_t will be represented by a trend-stationary process with a breakpoint in the trend which occurs at an unknown point in time. The value of λ will be selected according to the trend-stationary alternative which receives the most of the weight.

The estimation procedure we choose consists in selecting the breakpoint that gives the least favourable result for the null (6) using the testing procedure (4).

Once the null is defined as in (6), we no longer need the dummy variable $D(T_b)$ in our model specification and therefore the testing for a unit root can be conducted using the following model specification:

$$\textit{Model}\quad (1'):\quad y_t = \mu^{(1)} + \theta^{(1)} DU_t(\lambda) + \beta^{(1)} t + \alpha^{(1)} y_{t-1} + \sum_{j=1}^{k} c_j^{(1)} \Delta y_{t-j} + e_t \tag{1'}$$

$$\textit{Model}\quad (2'):\quad y_t = \mu^{(2)} + \beta^{(2)} t + \gamma^{(2)} DT_t^{*}(\lambda) + \alpha^{(2)} y_{t-1} + \sum_{j=1}^{k} c_j^{(2)} \Delta y_{t-j} + e_t \tag{2'}$$

$$\textit{Model}\quad (3'):\quad y_t = \mu^{(3)} + \theta^{(3)} DU_t(\lambda) + \beta^{(3)} t + \gamma^{(3)} DT_t^{*}(\hat{\lambda}) + \alpha^{(3)} y_{t-1} + \sum_{j=1}^{k} c_j^{(3)} \Delta y_{t-j} + e_t \tag{3'}$$

where: $DU_t(\lambda) = 1 \quad \text{if} \quad t > T\lambda$

$= 0 \quad \text{otherwise}$

$DT_t^{*}(\lambda) = t - T\lambda \quad \text{if} \quad t > T\lambda$

$= 0 \quad \text{otherwise}$

2.1 Integration analysis

In this section we present the results, reported in Table 1, obtained by applying the univariate Dickey-Fuller analysis to our data about Polish economy in the period from 1970 1st quarter to 1991 4th quarter.

In our analysis we tested the null hypothesis, represented by model (6) above, of the presence of a unit root, by using the framework proposed by Dickey and Fuller (1979, 1981). In particular, we made the tests with reference to both models proposed in DF literature, namely a model of a random walk with drift under the null hypothesis, both in the augmented version of the model, *i.e.*:

$$y_t = \mu + \rho Y_{t-1} + \sum_{i=1}^{k} \delta_i \Delta y_{t-1} + \varepsilon_t \tag{7}$$

and a model of a random walk with drift and trend under the null, which is written as:

$$y_t = \mu + \rho Y_{t-1} + \beta t + \sum_{i=1}^{k} \delta_i \Delta y_{t-1} + \varepsilon_t \tag{8}$$

In the second column of Table 1 these two models are respectively labelled "notrend" and "trend". We also report, at the end of the Table, the 5% and 10% critical values of DF test for T=80. It is important to remember that the DF testing procedure is valid if residuals in estimated models are white noise; this is the reason why we accompany the DF statistics (reported in the third column of Table 1) with a complete set of mispecification test and their respective *p*-values. In particular we report Godfrey's 4th order autocorrelation test, labelled "LM(4)", White's heteroskedasticity test, labelled "White", and Engle's 4th order autoregressive conditional heteroskedasticity test, labelled "ARCH(4)". Where necessary, the models with or without trend were both adequately augmented with a number of lags (their order is indicated in column 4 of the Table) in order to prevent residuals autocorrelation.

The main problem in interpreting the results in Table 1 is represented by the possible presence of a break in our series. Among various problems, the break in Polish variables at the end of the 80s may induce heteroskedasticity in the residuals of the two previous models, and offer a non consistent OLS estimate of the variance-covariance matrix. In order to tackle the problem we present, in the latter two columns of Table 1, the DF statistics obtained by correcting (when necessary) the heteroskedasticity in the regression results. In particular the first correction, as proposed in Kim and Schmidt (1993)[3], was made by using the White (1980) heteroskedasicity consistent covariance matrix estimator (the corresponding DF test is labelled "ADFw" in Table 1); another correction was made using Newey and West (1987) generalization of White's estimators. This generalization requires an explicit setting of the size of the Bartlett window; by choosing the window size equal to zero one obtains White's estimator, in our study we set a truncation point equal to 15 (the corresponding DF test is labelled "ADFnw" in Table 1).

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3 More precisely in that paper Kim and Schmidt analyze the effects on DF test of the presence of residuals GARCH(1,1). Even though non formally demonstrated but on the basis of a Monte Carlo experiment, the authors say that "*the White standard error correction improves the accuracy of the test fairly dramatically, especially in the integrated and degenerate case. It generally does not solve the overrejection problem entirely, however*", Kim and Schmidt (1993, p. 293).

Tab. 1 - The results of Dickey-Fuller analysis

Variable (*):	model	ADF	lag	LM(4)	p-value	White	p-value	ARCH(4)	p-value	ADFw	ADFnw
Consumer prices	notrend	-3.07	8	4.31	0.37	1.23	0.27	24.5	0	-2.63	-3.1
	trend	-4.02	8	3.54	0.47	0.32	0.56	19.24	0	-3.92	-3.4
Average wages	notrend	-2.42	8	4.64	0.32	0.16	0.68	2.03	0.73		
	trend	-3.26	8	6.96	0.14	0.06	0.81	1.44	0.84		
Import prices	notrend	-9.29	0	0.33	0.99	0.02	0.89	0.08	1		
	trend	-9.62	0	0.54	0.97	0.02	0.89	0.06	1		
Output	notrend	-8.96	0	2.84	0.58	0.01	0.93	2.62	0.62		
	trend	-10	0	2.36	0.67	0.06	0.8	4.19	0.38		
Employment	notrend	-6.51	0	6.41	0.17	15.9	0	17.9	0	-3.06	-4.49
	trend	-8.74	0	0.67	0.96	3.53	0.06	4.35	0.36		
Labour productivity	notrend	-9.37	0	3.8	0.43	0.97	0.32	5.64	0.23		
	trend	-9.71	0	6.2	0.18	0.46	0.49	7.71	0.1		
Unit labour costs	notrend	-3.37	1	8.1	0.09	5.5	0.02	7.47	0.11	-1.98	-3.7
	trend	-4.13	1	7.7	0.11	6.8	0.01	7.34	0.12	-2.44	-4.69
Real average wage (**)	notrend	-2.97	4	6.08	0.19	0.09	0.76	2.15	0.71		
	trend	-2.95	4	5.63	0.23	0.35	0.55	1.93	0.75		
5% critical values	notrend	-2.89								-2.89	-2.89
10% critical values	trend	-3.46								-3.46	-3.46

(*) Test are based on the first difference of the logarithm of the series
(**) Logarithms of the levels

Table 1 only reports the results of the DF integration test applied to the first differenced series in logarithms, because in a preliminary stage of the analysis (not shown in the Table) both the previous models offered DF statistics that did not reject the null hypothesis of the presence of a unit root for the log-levels of all the variables examined.

Apart from the structural break problem, that will be analyzed in the next section, the results about import prices, output, employment and labour productivity are quite clear cut: all the variables seem to be integrated of the first order, given that their first differences seem to be all stationary. The first differenced average wage series does not seem to be stationary, given that, in absence of relevant mispecification problems, both models (with and without deterministic trend) produce DF statistic that do not reject the null hypothesis of unit root. The average wage series would seem to be second order integrated[4].

The analysis is more complicated for domestic prices and unit labour cost because the "traditional" Dickey-Fuller analysis would offer a DF statistic that reject the presence of unit roots, but the residuals mispecification tests show some heteroskedasticity. In particular, inference made on the first difference in consumer prices is incorrect because of a strong ARCH heteroskedasticity. Using White's heteroskedasticity consistent covariance matrix estimator for correcting the standard errors estimates, the model without trend does not reject the null hypothesis of unit root. On the other side if we use Newey-West's correction the null is not rejected for the model with trend. Unit labour costs estimates are affected by heterosckedasticity and the use of a consistent covariance matrix estimator gives results completely dependent on the method of correction adopted: no stationarity if we refer to both models with White's heteroskedasicity correction, rejection of the unit root hypothesis with Newey-West's correction.

At this stage we feel it is important to perform a deeper analysis of structural change in order to avoid the presence of spurious unit roots in the series.

2.2 Univariate break-endogenous analysis

The breakpoint analysis sketched at the beginning of this section has been applied to our variables; the main results are reported in Table 2 and deserve the following comments.

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4 This result is confirmed by the stationarity of second order differenced wage series, not reported in the Table, and is in agreement with the findings in Blangiewicz and Bolt (1992), based on a different sample period.

Prices. The first part in Table 2 presents the results about prices and inflation (approximated by the first difference in the prices logarithm). Prices are measured by the index of consumer prices (base 1980=1) and do not present any seasonality. Given a visual inspection of prices data, we choose model 3 (that assumes a break both in the intercept term and in the slope) as the most appropriate model to test for the break. The null hypothesis that prices are a random walk with a drift is not rejected at a 5% significance level (our minimum t statistic is -3.58 in 1986 2nd quarter against a 5% critical value of -5.08). One can also notice that both models 1 and 2 show a 5% significance break in 1989 3rd quarter: the t statistics are respectively -4.89 and -4.75 against 5% critical values of -4.80 and -4.42.

Tab. 2 - Results of the estimated breakpoint test statistic

	Model 1		Model 2		Model 3	
Variable:	t stat	period	t stat	period	t stat	period
Price levels	-4.89	1989.3	-4.75	1989.3	-3.58	1986.2
Inflation rate	-5.85	1989.2	-5.64	1989.2	-7.05	1989.3
Average wage	-5.53	1989.3	-5.3	1989.3	-4.68	1987.1
Wage growth	-4.81	1988.3	-4.61	1988.3	-7.67	1989.3
import prices	-5.2	1988.4	-5.02	1988.4	-3.71	1985.3
Import inflation	-4.78	1990.3	-4.78	1990.3	-12.55	1989.1
Official exch. rate	-5.98	1989.1	-5.42	1989.1	-3.32	1989.1
Output	-2.27	1990.1	-2.28	1990.1	-2.27	1988.1
Output growth	-4.97	1989.3	-4.94	1989.3	-4.93	1982.2
Employment	-2.98	1990.1	-3.09	1990.1	-3.27	1990.1
Employment growth	-4.67	1990.1	-4.71	1990.1	-3.97	1989.4
Labour productivity	-2.33	1989.3	-2.38	1989.3	-2.22	1988.1
Productivity growth	-5.35	1982.2	-5.14	1989.3	-5.36	1982.2
Unit labour costs	-5.6	1989.3	-5.38	1989.3	-3.98	1987.1
ULC growth	-4.74	1989.2	-4.45	1989.2	-8.18	1989.3
Real average wage	-5.08	1981.4	-5.7	1981.4	-5.53	1981.4
5% critical values	-4.8		-4.42		-5.08	
10% critical values	-4.58		-4.11		-4.82	

As far as inflation is concerned, the results about the presence of significative breaks in the analyzed period is quite clear. In fact models 1, 2 and 3 are all characterized by a minimum t statistic that rejects the null hypothesis in 1989 2nd quarter (models 1 and 2) and in 1989 3rd quarter (model 3).

Nominal and real wages. The average monthly wages in State sector behaviour is quite similar to the price one. Before applying the testing procedure for detecting the presence of a significant breakpoint, wages data were seasonally adjusted by using the X11 method because of their seasonality. Average wage logarithms are not significant at the 5% level in model 3, that we consider the most appropriate for this variable, after graphical inspection of the data.

The significance of the test is not rejected at the 5% level for the first differences in wage logarithms (that approximate the wage growth). Even though models 1, 2 and 3 show different break points, it is important to note that wage growth models 1 and 2 present values of the test slightly over the critical values in 1989 3rd quarter while model 3 shows a unique break in 1989 3rd quarter. For this reason we interpret the results as suggesting a wage growth break in 1989 3rd quarter.

Import prices and official exchange rate. The import prices in national currency were obtained by multiplying the price index of manufactured goods in the world trade (in US dollars) times the series of official exchange rate of the national currency, against the US dollar. The log-levels of import prices show a break in 1988 4th period if we use models 1 or 2, while a major break in 1989 1st quarter is detected by applying model 3 to the first differenced variable.

As far as the exchange rate is concerned, the tests about this variable have a behaviour quite similar to that shown by prices in levels: a single break in 1989 1st quarter is present in models 1 and 2 for the exchange rate levels, while model 3 does not present any break in the period analyzed (the minimum t statistic is a not significative -3.32 in 1989 1st quarter). We can also notice that the information about an exchange rate break in 1989 1st quarter tends to confirm the break period evidenced by using the model 3 specification combined with first differenced import prices series.

Output, employment and productivity. Output is measured by the index of industrial production (basis 1980=1), data were seasonally adjusted by X11 method. The output logarithm is not trend stationary in any model, data are not significative at even the 50% level. Output series differentiated once (that approximate the output growth rate) are always trend stationary with a break in 1989 3rd quarter (for models 1 and 2) and in 1982 2nd quarter (model 3, but in this case at only the 10% level of significance), in addition we note that model 1 shows another significative break at the 5% level in 1982 2nd quarter (-4.88 against a critical value of -4.8).

Employment is measured by the average employment in the State sector; like the output ones, employment data were seasonally X11 adjusted. The results of testing for unit roots in the logarithms of the levels are quite clear cut: in all the models analyzed the null hypothesis of random walk with drift is not rejected. When inspecting the results from the regression in differences, we note that the only case where the null is rejected at the 5% significance level is model 2 (the value of the t statistics is -4.71 against a critical value of -4.42). The break in this latter model is shown in 1990 2nd quarter.

We approximate the logarithms of the labour productivity by using the difference between the logarithm of output and the logarithm of employment. The non rejection of the null hypothesis analysing the levels of output and employment previously mentioned can, in absence of cointegration between output and employment, help us to explain the non rejection of the null for the logarithms of the productivity levels (in all the three models examined). On the other side, for all the models (1, 2 and 3) the hypothesis of the segmented trend stationarity is accepted at a 5% level of significance; periods of break vary from 1982 2nd quarter (models 1 and 3) and 1989 3rd quarter (model 2). It is also interesting to note that models 1 and 3 show another break (t statistics are slightly over the minimum t statistic, respectively -5.18 in model 1 and -5.12 in model 3) in 1989, 3rd quarter.

Unit labour costs and real wages. Both these variables are obtained by combinations of previously analysed variables. The difference between the logarithms of the average wage and productivity is used to approximate the logarithm of unit labour costs. In this case the choice of the model (1, 2 or 3) is crucial in order to establish if this variable reveals broken trend stationary or a random walk with drift. In fact, on the basis of models 1 or 2, we have an evidence for break-stationarity of the variable in levels; otherwise, using model 3, the evidence is in favour of a random walk with drift. In any case is very important to stress that the break period is always the same: 1989 3rd quarter. In all the complementary cases the null hypothesis is rejected at the 10% significance level.

The real wage is obtained by the ratio average wage on consumer prices. The real wage is broken trend stationary in the three models analyzed (1, 2 and 3). The break period in this case is unique, 1981 4th quarter.

After all the integration analysis results reveal to be very sensitive to the model chosen under the alternative. The only robust result seem to be that the data are broken-trend stationary (in levels or in differences) while no evidence was found supporting the hypothesis of I(2) series. This result apparently contraddicts what we get in the preliminary staionarity analysis, performed without taking into account the break points of the series, for wages, unit labour costs and prices.

3. Theoretical hints for model specification

The theoretical model is based on a simple mark-up pricing system where price movements, either through controls by public authorities or through the behaviour of monopolistic firms, are linked to cost movements. At the same time we rule out any form of market constraints on prices or wages, except for the case where we implicitly assume that price or wage controls must be compatible with the stability of the whole system.

3.1 Price equation

Prices are assumed to be a function of money, exchange rate, foreign prices and real factors; moreover a choice variable should be included summarizing the preference for their desired level. Furthermore, in countries where administered prices are important, subsidies arising from price control would tend to repress inflation in the short run, and the elimination of such subsidies would result in an increase of inflation. Moreover, with regard to the exogenous shift factors affecting the inflation rate, a specific weight has to be attached to the discrete adjustments to administered prices. Such adjustments can be viewed as a lagged response by the planner to macroeconomic imbalances (particularly excess demand in goods market) and as an attempt to modify the structure of relative prices.

A mixed system of controlled and market prices leads to two sources of non synchronization in price setting. The first one is related to the non synchronization between the change of controlled prices and the change of market prices, while the second one arises from the lack of coordination of price settings for the market goods. The greater the proportion of controlled prices on total prices, the weaker the link between price changes and capacity levels.

The dynamics of this equation would also be affected by the timing of price decisions, as administrative rules generally impose long lags between adjustments of controlled prices while the timing of changes for market prices is determined by price setters. We assume that the planner takes care of macro disequilibria and periodically fixes controlled price levels with the aim of reducing the perceived imbalances. The dynamic form of the price equation would include an output gap, along with a variable related to the purchasing power in consumer markets, to try to keep particular excess demand features under control, when controlled and market prices co-exist and when the planner monitors their relative prices. To the extent that supply remains insensitive to price changes, one would expect a strongly damped equilibrating relationship between shortage and inflation.

3.2 Wage equation

Wages are treated as largely exogenous and structured by centrally determined norms. It is well known that in partially reformed economies, productivity related wages introduce a crude association of wages to the average productivity rather than to the marginal one. We assume that wage growth adjusts to the centrally determined rate of annual expansion of the economy. A wage expansion could be expected to result from a centralized decision related to expected inflation and a trend capturing the productivity changes, offset by tax increases.

The dynamics of the wage equation may be affected by two elements. First of all the official prices do not reflect the actual inflation rate in an economy where important black goods markets exist. Second, some mechanism would capture the sensitivity of wage demands to the availability of goods in the economy, because in systems characterized by chronic shortages, availability of goods may significantly affect wage demands. We expect that greater availability should induce demand for higher wages since they can be translated in higher real consumption. A variable measuring the deviation from trend of inventories of consumer goods at the retail level may approximate availability, since inventories of consumer goods in centrally planned economies are kept to a minimum.

Both price staggering and wage indexation could be expected to translate shocks at the price level into a persistent increase in the rate of inflation.

4. Cointegration analysis in the presence of structural breaks

4.1 The general design of the multivariate exercises

In this part of the paper we are interested in obtaining quantitative answers by applying to Polish data the theoretical model discussed in section 2. The study of the multivariate case will be conducted by making a number of steps that can be summarized as follows:

1st step. We go from the theoretical structure (presented in section 2) to a two equation estimable model. In other terms, in this part we try to link the economic theory with the identification of a set of specific (observable) variables of interest:

$$LP = f_p(ULC, LPM)$$

$$LW = f_w(LP, ETA)$$

where (all the variables are expressed in logarithms):

LP = consumer price index (proxy variable for the domestic, home country, prices);

LPM = import price index (international manufactured prices in US dollars times Polish official exchange rate, this is a proxy variable for the prices outside the home country);

LW = average wage in the State sector (proxy variable for labour income in the home country);

ETA = labour productivity in the home country (whose proxy is the ratio of the home country industrial production and total the employment in the State sector);

ULC = unit labour costs (obtained, by definition, from the difference *LW-ETA*);

Obviously, this step implies a number of specific assumptions about the behaviour of the Polish economic agents and we have to stress that, unfortunately, many of the assumptions made in this paper were driven by data availability. The basic hypothesis is that the behaviour of the State and the private sector are consistent and that, for example, we can use mixed information (extracted from both private and State sectors) in order to quantify the aggregate behaviour of the Polish economy.

At the end of this first step we analyze the relationships among the variables that are included in each equation in order to clarify their basic internal connections. This exercise will lead to a two (partial) sets of multivariate analysis, respectively *LP, ULC* and *LPM* in one side, and *LW, LP* and *ETA* on the other side.

2nd step. Then, we put together all the variables of interest, irrespective of whether they explain the price equation or the wage equation, in order to investigate the overall coherence of the results obtained in the first step. More precisely the only multivariate analysis in this step involves all the four variables that we want to analyze: *LP, LPM, LW,* and *ETA*. Of course, we will ignore the behaviour of *ULC* because it is a simple linear combination of other two variables (*LW* and *ETA*).

4.2 The analysis of the two multivariate subsets

In this section we describe the results obtained in the previously outlined step 1. The statistical techniques used are based on Johansen's multivariate cointegration analysis.[5]

Price equation. The cointegration analysis is based on the assumption, that we consider acceptable in the light of the results of the univariate analysis, that all the three variables of interest (*LP, ULC* and *LPM*) are all integrated of order one.

A VAR of order 2 with drift and trend passed the standard residual whiteness tests.

The first problem we face is to identify the cointegration rank, r, of the system. Given that we are analyzing three variables integrated of order one, the range of possibilities is $0 \leq r \leq 2$. In particular, if the cointegration rank is equal to zero, the three variables are not cointegrated in the long run and consequently a stable long run relationship among the three variables does not exist. On the other side, if the cointegration rank is two (its maximum value), the price equation variables are linked together by more than one long term relationship. The results of the rank cointegration analysis is reported in Table 3, that summarizes the cointegration analysis for both the price (in Table 3a) and the wage (in Table 3b) systems.

The results of the cointegration tests are quite unambiguous: the null hypothesis of no cointegration is rejected and the null of cointegration rank lesser than (or equal to) one is not rejected, so we can conclude that consumer prices, unit labour costs and import prices are linked together by a unique cointegrated relationship, presented (in normalized form) in Table 3a. Given the estimates of the price equation cointegration vector, we tested the homogeneity of labour costs and import prices effects by imposing that the two parameters estimates add to one. The test (of likelihood ratio type, LR) did not reject the restriction; Table 3a shows both the estimation results of the restricted long term cointegration vector and the LR test of the parameters restriction.

At this stage we note the two following stylized facts:

- The *LP, ULC* and *LPM* variables show a unique long term cointegration relationship that we identify as the homogeneous domestic price long term equation.

5 See *e.g.* Johansen (1988, 1989)

Tab. 3 - Rank test, estimation and restrictions in the 1st step

a) Price equation variables subset

86 observations from 1970Q3 to 1991Q4. Maximum lag in VAR = 2.
List of variables included in the cointegrating vector:
LP ULC LPM

Cointegration LR Test Based on Max Eigenvalue of the Stochastic Matr.:

Null	Alternative	Statistic	95% Critical Value	90% Critical Val
r = 0	r = 1	69.6698	20.9670	18.5980
r<= 1	r = 2	10.2822	14.0690	12.0710
r<= 2	r = 3	.013121	3.7620	2.6870

Cointegration LR Test Based on Trace of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Val.
r = 0	r>= 1	79.9651	29.6800	26.7850
r<= 1	r>= 2	10.2953	15.4100	13.3250
r<= 2	r = 3	.013121	3.7620	2.6870

Normalized Estimated Cointegrated Vector (r=1):

LP	ULC	LPM
(-1.0000	0.1033	0.8642)

Restricted Cointegrated Vector:

LP	ULC	LPM
(-1.0000	0.2137	0.7863)

LR Test of Restriction: CHI-SQ(1)= 1.2225[.269]

b) Wage equation variables subset

86 observations from 1970Q3 to 1991Q4. Maximum lag in VAR = 2.
List of variables included in the cointegrating vector:
LW LP ETA
List of additional I(0) variables included in the VAR:
CONST82 BREAK82

Cointegration LR Test Based on Max Eigenvalue of the Stochastic Matr.:

Null	Alternative	Statistic	95% Critical Value	90% Critical Val.
r = 0	r = 1	31.4666	20.9670	18.5980
r<= 1	r = 2	6.9231	14.0690	12.0710
r<= 2	r = 3	5.0388	3.7620	2.6870

Cointegration LR Test Based on Trace of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Val.
r = 0	r>= 1	43.4285	29.6800	26.7850
r<= 1	r>= 2	11.9619	15.4100	13.3250
r<= 2	r = 3	5.0388	3.7620	2.6870

Normalized Estimated Cointegrated Vector (r=1):

LW	LP	ETA
(-1.0000	1.0425	0.7653)

Restricted Cointegrated Vector:

LW	LP	ETA
(-1.0000	1.0000	0.7474)

LR Test of Restriction: CHI-SQ(1)= 2.1187[.146]

- From an economic point of view, we note that the import prices effect on domestic prices (with an elasticity equal to 0.786) seems to be bigger than the unit labour cost effect (with an elasticity equal to 0.214).

Wage equation. The wage equation analysis was very similar to that performed for the price equation. Also in this case we use a VAR(2) general model (involving *LW, LP* and *ETA* variables), including both drift and trend. With respect to this equation we decide to include two particular deterministic components in the VAR(2) system: a 1982 2nd quarter break in both drift (*CONST82*) and trend (*BREAK82*) terms. This choice was suggested by the results of the univariate breakpoints analysis that showed the heterogeneity of the breakpoints affecting the three variables here studied; for this reason we decided to include the hypothesis of a 1982 2nd quarter break in the VAR(2) model for the wage system, such a break characterizes in particular the productivity (*ETA*) variable.[6]

The rank estimation and test phases are summarized in Table 3b. As for price system analysis, the cointegration rank is clearly 1 and, given the long term estimates of the unrestricted model, we tested that the price effect on wage is 1; the LR test did not reject the null (the specification test and the restricted model estimates are both reported in the lower part of Table 3b). The test of unit wage elasticity to labour productivity was rejected.

The two relevant facts that we learned in this stage are:

- The existence of a single long term cointegration relationship among *LW, LP* and *ETA* variables, that we identify as the real wage long term equation.

- From an economic point of view, the main result is the unit elasticity between wage and consumer prices that allows us to specify a real wage long term equation where the only explanatory variable is the labour productivity (characterized by an elasticity significantly lower than one).

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6 In a part of the research not presented here, we tested for cointegration by omitting the two deterministic components *CONST82* and *BREAK82* (breaks in drift and trend in 1982 2nd quarter); in this case we were not able to detect any cointegrated combination for the three variables analyzed. A possible explanation of this fact is that, apparently, the cointegrated combinations are not heavily affected by a break if it is common to all the variables included in the multivariate analysis. For example, this was the case we experienced in testing for cointegration among *LP, ULC* and *LPM* variables, that are characterized by "almost common" periods of break (see the results of univariate breakpoint analysis in section 2.2). However, in this case only two variables, *LW* and *LP*, have breaks almost in the same period (1989 3rd period), while *ETA* is heavily broken in a preceding period (1982 2nd period). This fact can explain the absence of cointegration among the three variables if we do not allow for a specific break in 1982 2nd period.

4.3 The cointegration analysis for the whole set of variables

In this section the cointegration analysis is developed by merging all the variables (namely, *LP, LPM, LW,* and *ETA*) and by including in the VAR(2) the two deterministic components (*CONST82* and *BREAK82*) used in the wage subsystem.

In testing for cointegration it is possible to include, among the regressors in differences, any variable (that must be stationary) useful to improve the explanation of short term movements of the variables studied. In this specific case we decide to add the first difference of the series measuring the distances between black market and official exchange rate variables (*DPREM*), to the list of the regressors.[7] The inclusion of this regressor should help to capture some short term effects that may influence the behaviour of both price setters and workers.

On the basis of the cointegration rank analysis presented in Table 4 (that includes also the estimates of the cointegrating vectors), it is possible to detect two cointegrating relationships among the four variables studied. These relationships, as we will see, seem to be similar to the two cointegration vectors are connected to the two single relationships shown in the preceding section.

The main problem we must face now is how to identify the parameters of the two cointegrating vectors, given the hypothesis that they represent the two long term theoretical relationships for, respectively, domestic price and wage models. The strategy we decide to follow to answer to this question is the following. First we try to investigate the compatibility of domestic prices and wage determination subsystems, with the complete framework we are employing in this section. In other terms, we impose a number of restrictions on one cointegrating vector in order to identify, at a time, just one of the two vectors, leaving the other one free to adjust. The results of this double procedure are summarized in Table 5; the testing procedures show a clear non-rejection of the null (meaning that the restriction is compatible with the general framework). If we look carefully how the second vector adjusts when we constrain the parameters of the price equation (wage effect equal to 0.2, productivity effect equal to -0.2 and import price effect equal to 0.8), we see that in the second vector (which we identify as the wage equation) the sum of domestic price and import price effects on wages (the normalized parameters with respect to wages are in brackets) is almost equal to one. The productivity effect on wage has about the same level it had in the cointegration vector identified in the 1st step (0.614 here and 0.747 there).

For this reason we impose a set of complete restrictions that allowed us to identify the whole system. The results are presented in Table 5c; on the basis of a LR test, the restriction were not rejected.

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7 The variable *PREM* is equal to the log-difference between black market and official exchange rates and proxies the underground exchange rate market premium. Obviously,
$DPREM_t = PREM_t - PREM_{t-1}$.

Tab. 4 - Rank test and estimation of the whole system (2nd step)

85 observations from 1970Q4 to 1991Q4. Maximum lag in VAR = 2

List of variables included in the cointegrating vector:

LW LP ETA LPM

List of additional I(0) variables included in the VAR:

CONST82 BREAK82 DPREM

Cointegration LR Test Based on Max Eigenvalue of the Stochastic Matr.:

Null	Alternative	Statistic	95% Critical Value	90% Critical Val.
r = 0	r = 1	81.8119	27.0670	24.7340
r<= 1	r = 2	28.4195	20.9670	18.5980
r<= 2	r = 3	12.5942	14.0690	12.0710
r<= 3	r = 4	2.6853	3.7620	2.6870

Cointegration LR Test Based on Trace of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Val.
r = 0	r>= 1	125.5108	47.2100	43.9490
r<= 1	r>= 2	43.6990	29.6800	26.7850
r<= 2	r>= 3	15.2795	15.4100	13.3250
r<= 3	r = 4	2.6853	3.7620	2.6870

Estimated Cointegrated Vectors in Johansen Estimation
(Normalized in Brackets), r=2:

	Vector 1	Vector 2
LW	.098318	1.4244
	(-1.0000)	(-1.0000)
LP	-.44618	-1.3353
	(4.5381)	(.93748)
ETA	-.22802	-.95421
	(2.3192)	(.66990)
LPM	.35981	-.15247
	(-3.6597)	(.10704)

The first cointegration vector identifies the price equation:

$$LP = 0.2(LW - \mathrm{H}) + 0.8LPM = 0.2ULC + 0.8LPM$$

the second cointegration vector identifies the wage equation:

$$LW = 0.3LP + 0.7LPM + 0.6ETA$$

The latter equation is changed with respect to the equation identified in the 1st step, because there the price effect was captured by the domestic price effect (with coefficient equal to 1). The price effect on wage is splitted in two parts: a domestic price effect (with coefficient equal to 0.3) and a much more important effect represented by import prices (with effect equal to 0.7). However one can note that in any case the sum of the two price effects is still equal to 1.

At the end of this long run analysis of the interactions linking prices and wages in the Polish economy, we can list the following stylized facts:

- Import prices are the main source of wage-prices fluctuations in the Polish economy. The identified two-equation structural model highlights a strong effect of import prices on both domestic prices and wages.

- Passing from a simplified subsystem of variables (*LW, LP* and *ETA*) to a wider one (*LP, LPM, LW,* and *ETA*) we note the necessity to allow for a specific effect of import prices on wages, while previously (in the simplified VAR at three variables) such effect was included in the domestic price effect.

4.4 Analysis of some short term interrelations among variables

After the identification of the long term relationships between wages and prices in Poland, we try to discuss the short term behaviour of the variables studied.

The first step in the short term analysis is represented by the identification of the two short term disturbances (named *ECMP* and *ECMW*). This is given by the difference between the actual value of prices *LP* and wage *LW* variables and their respective long term explanations (the combination of cointegration parameters and regressors in the two equation system presented in the previous section):

More precisely these are given by the two long run combinations:

Tab. 5 - Testing for long term relationships

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a) Imposing the 1st step price equation parameters estimates to the whole system (Normalized in Brackets):

	Vector 1	Vector 2
LW	.20000	1.2278
	(-1.0000)	(-1.0000)
LP	-1.0000	-.35374
	(5.0000)	(.28812)
ETA	-.20000	-.75449
	(1.0000)	(.61453)
LPM	.80000	-.93774
	(-4.0000)	(.76378)

LR Test of Restrictions CHI-SQ(2)= 3.3163[.190]

b) Imposing the 1st step wage equation parameters estimates to the whole system (Normalized in Brackets):

	Vector 1	Vector 2
LW	-1.0000	-.18892
	(-1.0000)	(-1.0000)
LP	1.0000	-.17278
	(1.0000)	(-.91457)
ETA	.75000	-.021518
	(.75000)	(-.11390)
LPM	0.00	.37796
	(0.00)	(2.0007)

LR Test of Restrictions: CHI-SQ(2)= 1.6590[.436]

c) Testing a particular parametrization (Normalized in Brackets):

	Vector 1	Vector 2
LW	.20000	-1.0000
	(-1.0000)	(-1.0000)
LP	-1.0000	.30000
	(5.0000)	(.30000)
ETA	-.20000	.60000
	(1.0000)	(.60000)
LPM	.80000	.70000
	(-4.0000)	(.70000)

LR Test of Restrictions: CHI-SQ(4)= 5.3564[.253]

$$ECMP = LP - (0.2ULC + 0.8LPM)$$

$$ECMW = LW - (0.3LP + 0.7LPM + 0.6ETA)$$

It is interesting to note that the two long term combinations above include the relevant information we are looking for in this section.[8]

In order to study the short term dynamics, we must specify dynamically the short term relationships between the actual changes of the two variables we want to to explain (*DLP* and *DLW*)[9] and an error correction mechanism (in short ECM) that involves both short term effects (like current and lagged variations of some variables, *e.g.* productivity or import prices) and the distances from equilibrium (*ECMP* and *ECMW*).

Table 6 presents the results of the analysis of the two previous ECM specifications. Particular emphasis is given to a number of misspecification tests (in order to assess the significance of the behavioural models chosen) and to the parameter that explains the adjustment speed to long term equilibria. In the first column of the Table, the dynamic specification of price equation is labelled with "ecmp", and that of wages with "ecmw". From the second to the seventh column we show, respectively, Godfrey's residuals autocorrelation tests of 1st, 4th and 8th order [LM(1), LM(4) and LM(8)], White's heterosckedasticity test (W) and, finally, Engle's autoregressive conditional heterosckedasticity of 1st and 4th order (ARCH(1), ARCH(4)); in brackets, under the values of the tests, the respective p-values for the F distribution are reported. The "ECM term" in column 8 stands for the estimate of the parameter measured by short term disequilibria (standard errors in brackets); the last column shows the R^2 statistics in order to assess the capability of the short term model to explain the prices and wages short term fluctuations.

8 Obviously this claim is correct only if our implicit assumption of long term invariance of the parameters of the identified prices-wage model is correct (or, in other terms, if the structural breaks that affected the Polish economy during the 80s did not modify the long term structural parameters). Otherwise it would not be possible to accomplish the steps done in the previous phase of the research.

9 Where $DLP_t = LP_t - LP_{t-1}$ approximates the rate of change in consumer prices. The same for wage variable.

Tab. 6 - Estimation results of ECM models

	LM(1)	LM(4)	LM(8)	W	ARCH(1)	ARCH(4)	ECM term	R-SQR
ecmp	0.68	0.44	1.04	9.03	3.06	0.76	-0.104	0.74
	(0.41)	(0.78)	(0.42)	(0.004)	(0.09)	(0.56)	0.044*	
ecmw	0.02	1.48	1.54	0.207	0.052	0.333	-0.217	0.84
	(0.90)	(0.18)	(0.20)	(0.65)	(0.80)	(0.85)	0.029	

(*) White's standard error consistent estimate

As far as the price equation is concerned, we can exclude both autocorrelation and autoregressive heterosckedasticity, but we have to note a strong general White's heterosckedasticity that affects this specification. For this reason, we compare the estimate of the ECM parameter (equal to -0.104, that means slow adjustment to long term equilibrium) with the White heterosckedasticity consistent standard error. The analyzed specification is not only affected by heteroskedasticity: the estimated specification suffers also from structural instability, as it is clearly shown in figure 1, where the residuals of this estimate are compared with the corresponding recursive residuals, and in figure 3, that shows the recursive estimation of the ECM parameter. We can also note that the two period of bigger instability are at the beginning and at the end of the '80s, while in the last 8 quarters in the 90s the estimate seem to stabilize around -0.10. In a recent study, Charemza (1993) has found that, for the analysis of eastern economies, estimates in windows perform better. For this reason we show in figure 4 the plot of the rolling estimate for the ECM parameter (we set the window size equal to 40 periods). As far as the stability of the ECM parameter estimate is concerned, the results are quite similar to the recursive ones, but even more instable in the big shock period. In any case these further results tend to confirm the short term price equation structural instability.

The same facts seem to characterize the ECM specification for wages: in absence of relevant autocorrelation and heterosckedasticity problems, the structural instability of the specification is shown in figures 2, 5 and 6. The residuals and the recursive residuals, in figure 2, show very different tracks, especially at the beginning of the 80s. Both the recursive and the window estimates of the ECM parameter (respectively in figures 5 and 6) seem to stabilize just in the 90s, while they change before with important fluctuations. Another important feature is that the ECM parameter estimate in wage equation starts to be significantly different from zero just at the and of the 80s. Finally, given that the ECM parameter OLS estimate[10] for the period 1970-1991 is -0.217, we can assume also in this case slow adjustments towards the long run equilibrium, even though the detected parameter instability forces us to be particularly cautious in drawing this kind of conclusions.

4. Some concluding remarks

In this paper the price-wage spiral is analysed, for the Polish economy, for a period (1970-1991) that includes the transition to a marked-based economy.

Our empirical analysis deals mainly with the problems of stationarity and cointegration, in a situation where structural breaks are important. Structural breaks complicate stationarity and cointegration analysis on the one hand, and produce possible spurious results on the other.

Our findings reveal that all the series tested are broken-trend stationary in levels or in first differences. The analysis performed under the assumption of endogenous break-points reveal two major periods of significant structural changes, 1981.4-1982.2 and 1988.3-1989.4 .

Cointegration analysis of the whole system permits to detect two long-term significant relationships, which have been identified with wages and prices equations. Import prices seem to have an important explanatory role.

Short run dynamics also seem to be important for both equations, but seem to be characterized by coefficient instability which casts serious doubts on the forecasting ability of the model.

We consider these results as preliminary. Further investigation is in order.

10 Given the wage-price simultaneity, there is a serious possibility that the regressors used in the two ECM specifications are not weakly exogenous. In such a case it would be correct to apply instrumental variables estimates instead of ordinary least squares (OLS). We reported the OLS estimation results because the IV estimates where not significantly different from OLS ones.

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Fig 1 - Residuals and recursive residuals in price ECM equation

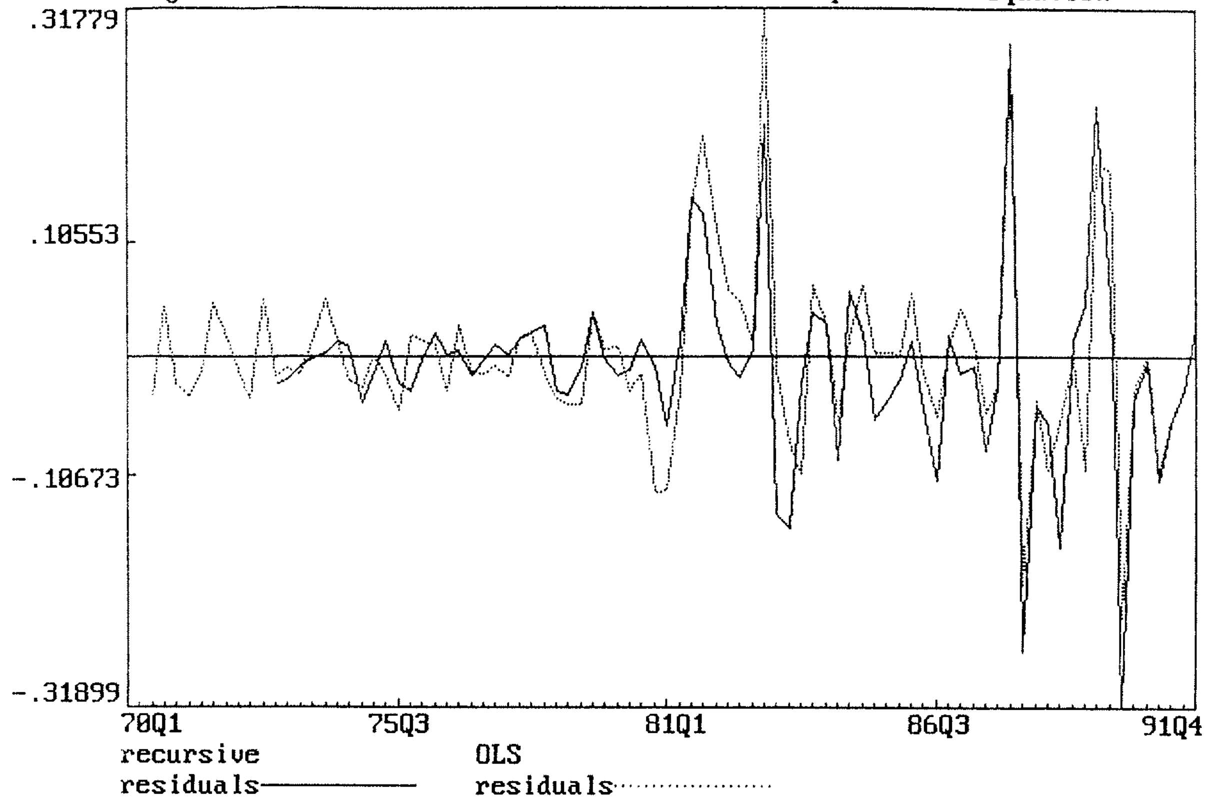


Fig 2 - Residuals and recursive residuals in wage ECM equation

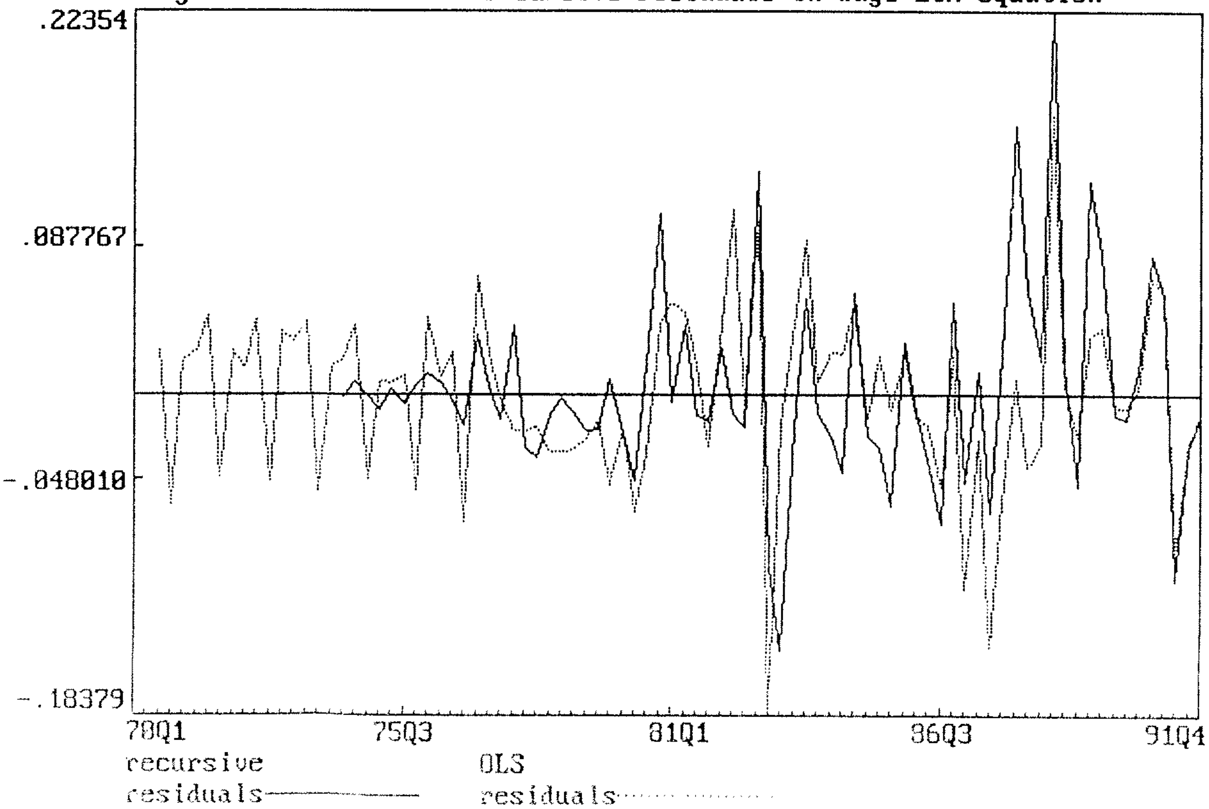


Fig 3 - ECM parameter recursive estimate in ECM price equation

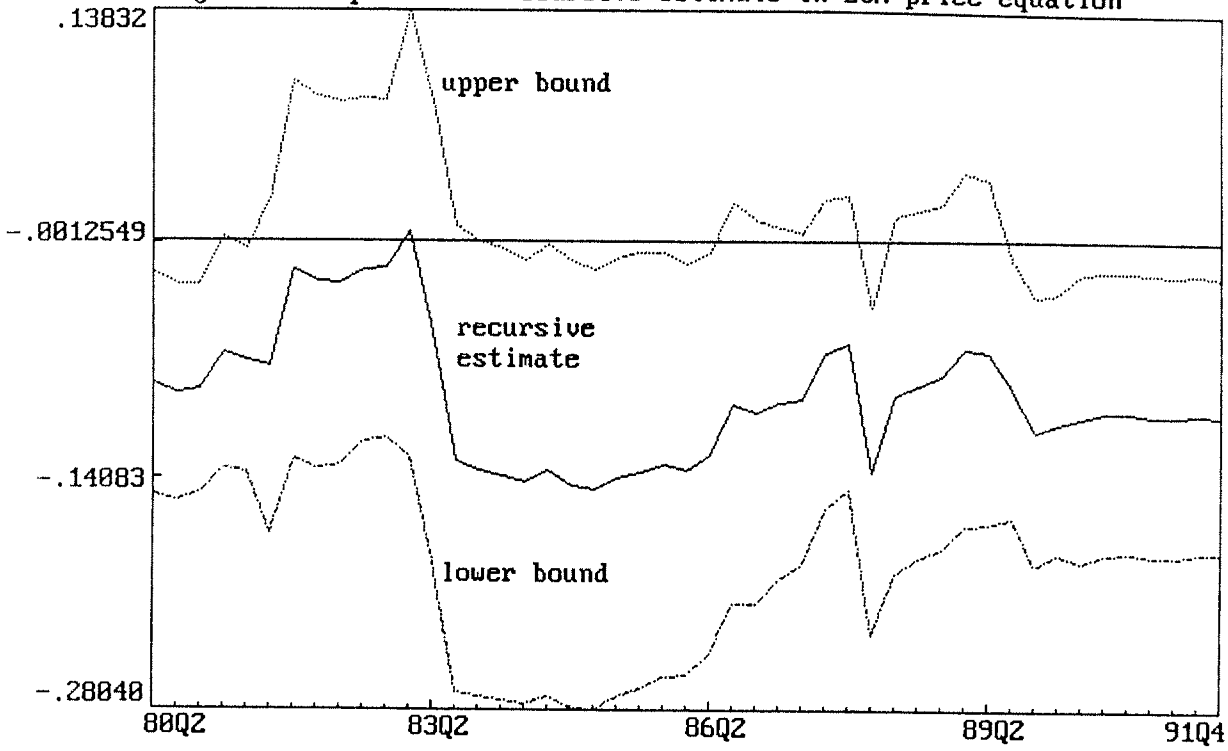


Fig 4 - ECM parameter rolling estimate in ECM price equation

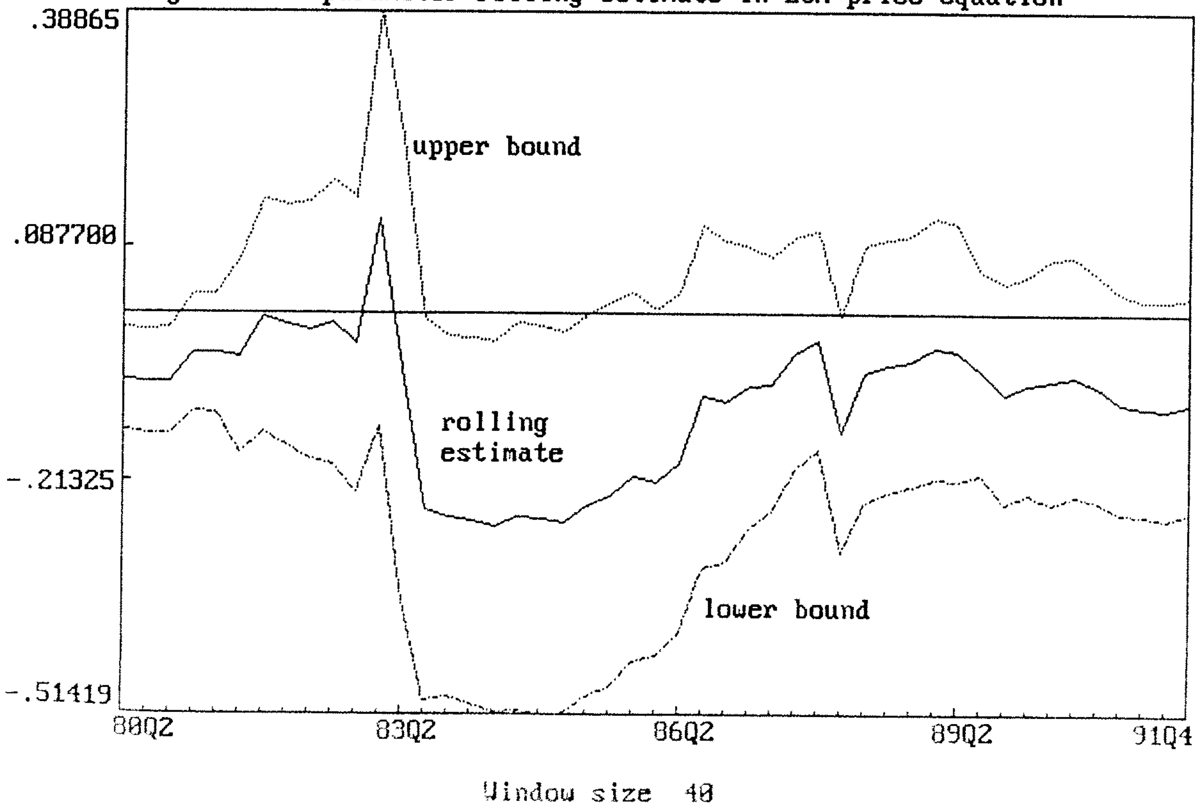


Fig 5 - ECM parameter recursive estimate in ECM wage equation

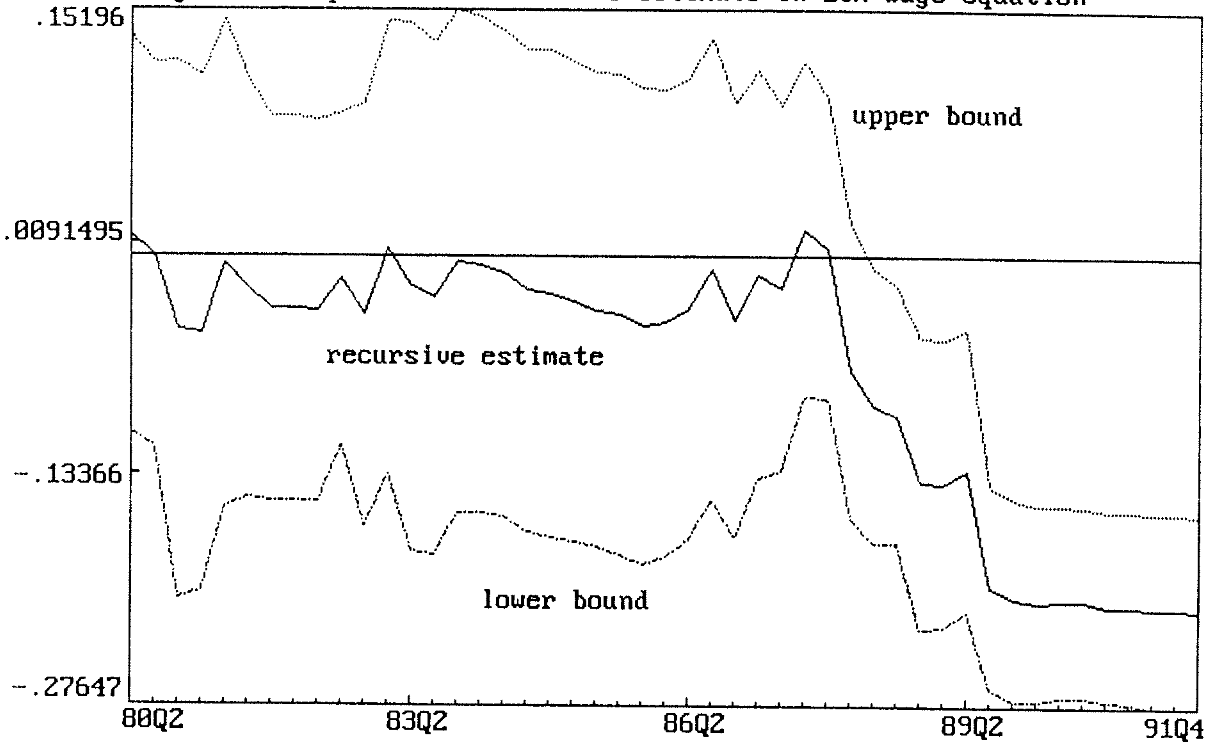
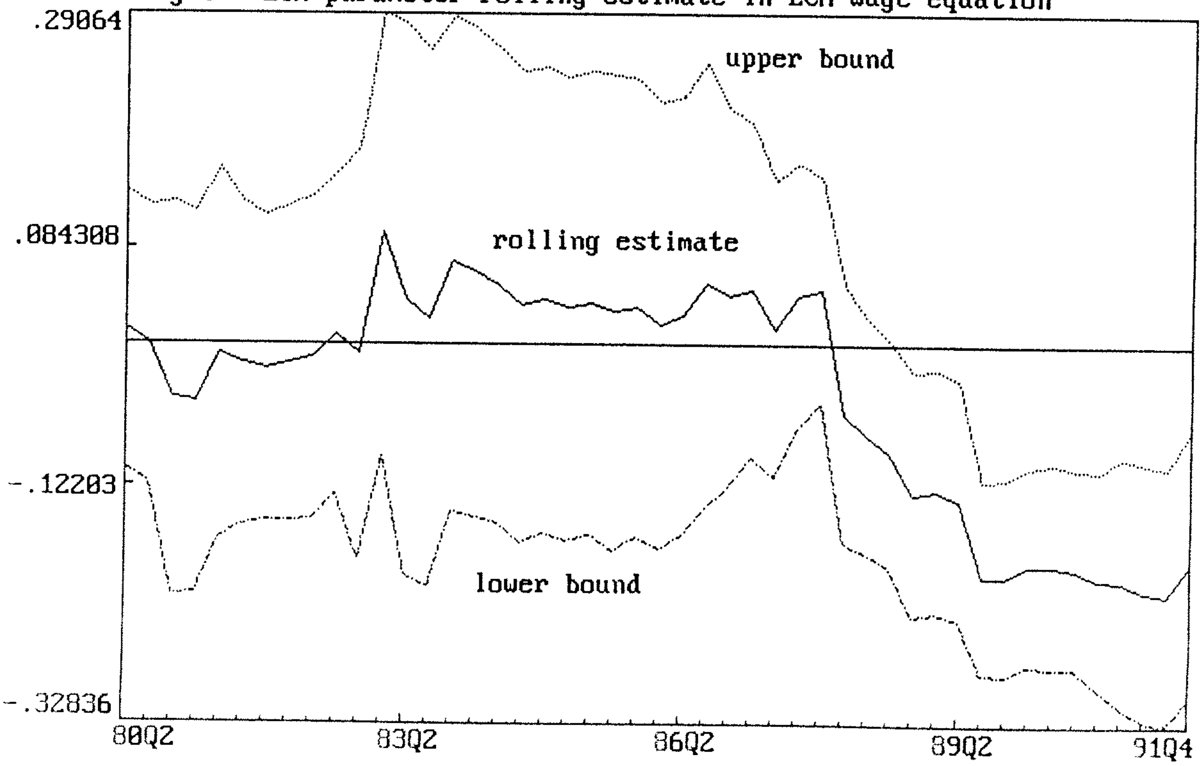


Fig 6 - ECM parameter rolling estimate in ECM wage equation



Window size 40