Exchange rate, inflation and unemployment in East European economies: the case of Poland and Hungary*

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Abstract

The aim of the paper is to model the impact of exchange rate on both inflation and unemployment variables in economies which are characterized by important structural changes, i.e. a transition phase moving from centralised economies towards market economies. This phenomenon, which is common to the east european countries, stressed different effects both for what concerns the behaviour of economic agents and for what concerns fiscal and monetary measures adopted by governments and aiming to keep under control the inflation-unemployment trade off. Time series relationships between these variables are investigated within an econometric model. Economic theory and the available data on the hypothetically relevant variables, along with the consideration of the main facts occurred in the period under study, characterize our information set. It is found that single equation analysis yelds inefficient inference relative to the whole system analysis, and important structural changes are detected which reflect possible breaks in the structure of the economic system along with a change in economic policy.

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1. Introduction

The aim of the paper is to model the impact of exchange rate on inflation and unemployment taking into account the peculiarities of economic systems under scrutiny, namely the transition from a centrally planned economy towards a market economy. This phenomena characterizes the economic behaviour of the main East European countries during the 90s, but stressed different effects both on economic agents and on the fiscal and monetary policies adopted by local Governments aiming to keep under control the inflation-unemployment trade off.

The model we use is an extension of the dynamic system proposed in Golinelli and Orsi (1994). After a general formulation of a linear dynamic system for both stationary and integrated data concerning Polish and Hungarian economies, we will adopt a general to simple modelling strategy, according to the so called LSE methodology (see Mizon, 1995 as a recent contribution on this argument). In the wide range of models used for economic analysis and in trying to characterize an appropriate way to develop and evaluate models, we think that the possibility of discovering economic structure using a progressive research strategy must be preferred to an approach which leads to models representing the pure time series properties of data.

In adopting a structural approach we implicitely recognize that the valuble information for the economic analysis of our model can come from numerous sources, among others the economic theory, the available observations of what we consider as the relevant variables, the knowledge of social, economic and historical facts of the period considered, and so on. According to this view, many aspects of the centrally planned economies, distinguishing them from standard market economies, will be investigated in order to try to understand and to model such disparities. This stylized facts include administrative and non-market based determination of prices and wages, artificially created full employment, external trading prices, monopolistic structure of goods markets and so on. During these periods, furthermore, important changes in relative prices took place and Governments adopted different policies with respect to the exchange rate and inflation rate in different countries.

The specific information provided by each phenomena mentioned above, will vary across the sources and the degree to which each source is related to the problem under study. Attention will be devoted to consideration of the extent to which available information has been exploited along with the relationship between this and the evaluation of models proposed.

The next section is devoted to present the main stylized facts characterizing the economies under study. Section 3 contains an econometric analysis of the time series relationshiph between exchange rate, prices, wages and labour productivity both in Poland and in Hungary, for the period 1971.2 -1995.3, from either a single equation or system perspective. The main findings confirm that single equation analysis can produce misleading inferences, with respect to system analysis, and that important changes in the structure of the economies appear significative. The final section provide some tentative conclusions.

2. Some Stylized facts for Poland and Hungary

Many significant features can be retained from the Polish and Hungarian experience of moving from an almost controlled price system towards a mixed regime where some price remain under control while others are market based prices. Empirical evidence based on statistical tests for structural breaks suggests that, the after reform period, must be considered as not simply characterized by an increase in the inflation rate, but rather as a new regime in prices behaviour.

When we consider the inflationary experiences for these two countries it is important to try to understand why similar measures had such a different impact and tended to generate almost divergent inflationary paths. A first difference is related to the conditions under which the reforms of the 80s were enacted. In Poland that period is marked by major external imbalance, persistent current account deficits and pressure on the exchange rate. Unlike in Hungary, monetary growth has been more important than price growth, creating significant imbalance which can be interpreted as the premium of the black market exchange rate over the official exchange rate. This premium persists all over the 80s, indicating a sort of repressed inflation regime in the system, and the frequent devaluations of the official exchange rate did not succed in reducing the size of the premium.

A second divergent feature relates to the wage setting process. In Hungary centrally imposed wage determintion rules led to moderate wage increases, while in Poland the wage bargaining was much more deregulated and caused a stronger wage and cost pessure, with the consequence of an inflation rate which is strongly related to changes in wages and exchange rate.

These factors of undoubted importance are not sufficient to explain, for example, the jump into high inflation rate which occurs in 1989 in Poland. Deterioration in fiscal accounts required an almost continuous depreciation of the exchange rate, the price liberalization of a large part of basic products along with a fastest adjustment of administred prices. The inflation deceleration phase, occurred after a period of high inflation and experienced in a different manner during the transition period, is rather different for the two countries under scrutiny and is partially related both to the measures of fiscal correction adopted and to the adjustments introduced for exchange rate and wages.

3. Modelling strategy: From single equation to system analysis

Some modelling strategies reveal to be more likely to get useful and satisfactory models than others, even if the route by which a model is discovered not necessairly determines its quality. In this light we will adopt a general to specific modelisation, and an empirical analysis approach which starts with a single equation analysis and leads to the whole system model. As a matter of facts, it is well known that single equation modelling has relevant effects on the system modelisation. So all what is related to the non-stationarity, structural break and cointegraton analysis is carried out more easily in a single equation context, and these empirical results are a good point of departure for the system analysis and a complete model specification.

Quarterly data from an updated data set for Poland and Hungary are used for the variables of interest, namely: consumer prices (PC), per capita wages (W), exchange rate against the US dollar (EX), labour productivity (ETA) and the employed-population ratio ($NPOP$). All these data are not adjusted for seasonality and have been available for the period 1971.2-1995.3. Since these variables are all non negative, the use of their logarithmic transformation is data admissible; in what follows their logarithmic values will be denoted, respectively, with LPC, LW, LEX, $LETA$ and $LNPOP$. The data appendix, at the end of the paper, provides additional information about data sources and definitions.

3.1. Poland

Visual inspection of graphs of these variables show that they have pronounced trends and consequently they need to be modelled either as stationary deviations from a linear trend or as variables with stochastic trends in the framework of an integrated-cointegrated system. Particularly, the variable measuring the productivity, i.e. $LETA$, shows also a linear trend. Deviations from it reveal wide fluctuations that is important to try to capture in order to model the process we thought as a possible generator of these movements. All these elements toghether are of practical importance in indicating the potential components we must consider in analysing the structure of the series along with the modelisation of their non-stationarity and possible regime shifts.

Subsequent autocorrelation analysis of these variables reveals high first order serial correlation coefficients with slowly decreasing higher order coefficient, indicating nonstationarity and a possible integratedness of order one. When we pass to consider quarterly variations, i.e. $DLPC$, DLW, $DLEX$, $DLETA$ and $DLNPOP$ their stochastic nonstationariry seems to be removed since the correlograms for these variable suggest I(0) patterns althought seasonal behaviours are revealed. Univariate test statistics for unit roots were computed in Golinelli, Orsi (1994) and the sample values compared with those provided by Dickey-Fuller in the augmented version. In computing these tests we choose a maximun length of 5 lags, and a certain evidence of structural breaks stands out. This information will be used later when we will try to model the system.

After this rough analysis on the univariate statistical properties of the series, in the spirit of a modelling strategy from specific to general (see Hendry (1995) and Mizon (1995) among others), we will start by analysing a single equation model. In a previous paper, Golinelli and Orsi (1994), we study the price-wage dynamic relationship for Poland within an integrated-cointegrated system when important structural changes in the economic policy occurs. The results we will present here however, are intended to describe the relationship between wages, prices, productivity and exchange rates, drawing attention to the special care that must be taken in the modelisation and forecasting of non-stationary time series with regime shifts.

Each of the series used reveals a nonstationarity and a partial integratedness of order 1. Univariate test statistics for the presence of unit roots using the augmented Dickey-Fuller testing procedure shows that the series are modelled as an I(1) process, i.e. they have a unit root in their autoregressive representation. All the variables considered, i.e. LPC, LW, LEX, $LETA$ and $LNPOP$, which measures the labour-population ratio, appear to be I(1), in the sense that the null hypothetisis of I(0) is rejected against I(1), while the null of I(1) is not rejecteed against I(2). Furthermore the real wage variable, i.e. $(LW - LPC)$, reveals to be I(1) with the consequence that the variables LW and LPC do not cointegrate each other.

In what follows, a model is used that allows a joint modelling of the relationships between all the five variables above, and of thier long run relationships.

The sample period on which our analysis is conducted, i.e. 1971.2-1995.3, contains important events as well as significant changes in the institutional setting of the country under scrutiny. In order to try to obtain a constant parameter model we put some specific consideration for these events by introducing dummy variables that take into account of that. For a complete list see the annex 1. The assumption of constant parameters does not preclude regime shifts, adaptation to changing shocks and so on; what it does require is that such factors are potentially open to being modelled as functions of more basic parameters that are constant in turn (cfr. Hendry (1995), pag. 355). For this reason, and in order to improve the ability of the model to interpret the above mentioned breaks, we added in the system, as exogenous, the variable *LPRIV*, the logarithm of the share of private employment in total employment. Among other things, this will allow a joint analysis of structural breaks and regime shifts, as they remain in the data under analysis.

The starting model we consider is a VAR that, for a generic vector x_t, shows the following

$$x_t = \sum_{j=1}^{k} A_j x_{t-j} + \gamma D_t + \varepsilon_t \tag{3.1}$$

containing the constatnt term, an unrestricted trend, seasonal dummies and other dummy variables and finally, as usual, ε_t is a vector of error terms zero means and constant covariance matrix. In the subsequent analysis we will assume that the roots of the $\det(I - \sum A_j L^j) = 0$ lie on or outside the unit circle in order to avoid explosive roots.

It is well known that the model (3.1) can be written in an observationally equivalent parametrisation, named the "vector equilibrium correction model" (VECM), having the following form:

$$\Delta x_t = \sum_{j=1}^{k-1} \Pi_j \Delta x_{t-j} + \Pi x_{t-1} + \gamma D_t + \varepsilon_t \tag{3.2}$$

where $\Pi = (\sum_{i=1}^{k} A_i - I)$ represents the static equilibrium response matrix and $\Pi_j = -\sum_{i=j+1}^{k} A_i$, for $j = 1, 2, ...(k-1)$.

Empirical results for the single equation anlysis of the unrestricted VAR (UVAR) (3.2), with 5 lags on each variable, for the whole period 1971.2 - 1994.3, are reported in table 1, while the last four observations are used for evaluating the parameter constancy forecast performance of the estimated model.

Table 1. Poland: UVAR diagnostic statistics (sample 1971.2-1994.3)*

	DLETA	DLW	DLPC	DLEX	DLNPOP
ser	3.56%	3.73%	2.67%	8.30%	0.51%
AR 1-5, F(5, 45)	0.545	2.194	2.427	0.418	1.853
	[.741]	[.072]	[.050]	[.833]	[.122]
ARCH 4, F(4, 42)	1.467	0.101	0.213	0.125	0.272
	[.229]	[.982]	[.930]	[.973]	[.894]
Normality, $\chi_2(2)$	7.683	2.245	1.198	1.274	5.996
	[.022]	[.325]	[.549]	[.529]	[.050]

Vector statistics: AR 1-5, F(125, 108) = 1.261 [.11]; Normality, $\chi_2(10)$ = 18.27 [.051]
Parameter constancy forecast test (1994.4-1995.3), F(20, 50) = 1.364 [.185]
(*) Each column corresponds to a single equation diagnostics. P-values are in brackets

The usual descriptive statistics such as serial correlation, conditional heteroscedasticity and normality test statistics are shown in the body of the table, while the statistics for vector autoregressive, vector normality and parameter constancy forecast are reported at the bottom of the same table[1]. The reported results, as a whole, are rather satisfactory, in the sense that the only apparent source of non-congruence is the AR test statistic for the $DLPC$ equation and normality test statistics for $DLETA$ and $DLNPOP$, all of which are significant at 5%. The inspection of the residual plots in figures 1a and 1b suggests the presence of some apparent outliers. The correlograms reported in the same figures suggests that a weak residual correlation may be in order for lags bigger than 5.

Finally model parameters constancy is checked through recursive estimation; even if the graphs reported in figure 2a for the 5 lags UVAR can only be considered as a sort of informal dignostics, we can observe that the plots of 1-step residuals with $\pm 2SE$ provide a strong evidence in favour of a reasonable constancy for all five equations. In figure 2b both individual equation and whole model break-point analysis using Chow F-testing procedures, scaled by their 5% significance level, are shown. Their values exceed critical values only for $DLNPOP$ equation, but almost no system break-point test values reveal to be significant.

If $x_t \sim I(1)$ the rank of Π corresponds to the number of cointegrating vectors. Let us indicate the rank of Π with r, while we will use α and β as matrices of order $j \times r$, of rank r, j beeing the number of variables modelled, and such that

[1]This study was carried out by using the PcFiml econometric package. In Doornik and Hendry (1994) there is a full explanation about the particular implementation of the procedures we used.

$\Pi = \alpha\beta'$. The reduced rank VAR system can be rewritten as:

$$\Delta x_t = \sum_{j=1}^{k-1} \Pi_j \Delta x_{t-j} + \alpha\beta' x_{t-1} + \gamma D_t + \varepsilon_t \tag{3.3}$$

where $\beta' x_{t-1}$ indicates the r cointegrating vectors and α represents the adjustment (loading) coefficients. It is well known that the maximum likelihood procedure suggested by Johansen (1988, 1995), modified in order to allow for dummy variables, enables the inference on r, once the systems (3.1) and (3.2) above are correctly specified.

The empirical values of the cointegrating test statistics, considering the presence of a linear trend in the cointegrating space, are reported in table 2. Denoting with r the rank of the cointegration matrix β, the cointegrating vectors are defined by the eigenvectors corresponding to the first r largest eigenvalues, and are the columns of the matrix β. Since the other $(j - r)$ eigenvalues, i.e. $\mu_{r+1} \ldots \mu_j$, should be zero for the non-cointegrating combinations, a test statistic of the hypothesis that the number of cointegrating vectors is r , $0 \leq r \leq j$, is given by the likelihood ratio (LR) statistic:

$$LR = -T \sum_{i=r+1}^{j} \ln(1 - \mu_i)$$

known as the *trace test*. An alternative test for the dimension of the cointegrating space can be based on the largest eigenvalue μ_{r+1} and is defined by the statistic:

$$LR = -Tln(1 - \mu_{r+1})$$

called the *max eigenvalue test*.

Both these tests have non standard distributions and appropriate critical values have been tabulated by Johansen (1988) and Osterwald-Lenum (1992).

The values for the *max* and *trace* test statistics we report in the cointegration tests tables are referred both to the standard tests and to the test statistics adjusted for the degrees of fredoom, in order to avoid the possibility of un incorrect estimate the dimension of the cointegrating space. The empirical values of the tests, reported in table 2, give support to the existence of three ($r = 3$) cointegrating vectors. These results are obtained in the presence of intervention dummies as well as of broken drifts into the model used for the cointegration analysis, and consequently the choice of the dimension of the cointegration rank must be done with some caution (in the sense that the critical values used would not allow for dummy variables). Incidentally a dimension of the cointegrating space of two ($r = 2$) can be accepted too, but the result of $r = 3$ would be preferred also on economic theoretical basis. In fact, from an economic point of view, we expect three long run relationships among the variables of interest.

Table 2. Poland: Cointegration statistics (sample 1971.2-1994.3)

rank:	1	2	3	4	5
log-likelihood	1793.05	1832.37	1849.60	1858.45	1862.33
eigenvalue μ	0.71	0.57	0.31	0.17	0.08
Max eigenvalue test	116.5*	78.6*	34.5*	17.7	7.8
Max eigenvalue test[a]	85.5*	57.7*	25.3	13.0	5.7
95% critical values	37.5	31.5	25.5	19.0	12.2
Trace test	255.1*	138.6*	59.9*	25.5	7.8
Trace test[a]	187.2*	101.7*	44.0	18.7	5.7
95% critical values	87.3	63.0	42.4	25.3	12.2

[a] Osterwald-Lenum (1992) definition.
(*) 1% significant.

The analysis of the long run relationships between variables and the related parameter identifiction is conducted on the basis of the results reported in table 3.

Table 3. Poland: Identified long run relations (standard errors in brackets)*

		LETA	LW	LPC	LEX	LNPOP	LPRIV	Trend
ecmLETA	=	1.000	-	-	-	-1.123	-1.043	-0.014
		(-)				(0.110)	(0.055)	(0.0003)
ecmLW	=	-0.506	1.000	-1.000	-	-0.881	-	-
		(0.035)	(-)	(-)		(0.189)		
ecmLPC	=	0.108	-0.108	1.000	-0.892	-	-	-
		(-)	(-)	(-)	(0.049)			

(*) Overidentifying LR test: $\chi_2(9) = 14.568$ [.104]; including 3 zero restrictions on the three loadings in DLNPOP equation.

One long-run relationship entering the cointegrating vector $\beta' x_t$ is given by the equation:

$$ecmLETA_t = LETA_t - 1.12LNPOP_t - 1.04LNPRIV_t - .014t \qquad (3.4)$$

implying that productivity, in the long-run equilibrium, shows an almost unit (positive) elasticity with respect to both the employed-population ratio and the ratio of private on total employment. Furthermore the autonomous labour productivity trend (linear) elasticity is about 1.4% per quarter.

A second estimated long-run relatonship which can be obtained from the cointegration analysis above is represented by:

$$ecmLW_t = (LW - LPC)_t - 0.506LETA_t - 0.881LNPOP_t \qquad (3.5)$$

The results seem to be compatible with those found in a previous study (Golinelli and Orsi, 1994) where the real wage elasticity to labour productivity was estimated to be around 0.4. Here we have the additional explanatory variable $LNPOP_t$ which is supposed to measure the long run effect of labour market attritions. These effects are positively related to real wages showing an elasticity of about 0.88.

A long-run relationship which is deducible from the cointegration analysis we performed is the following:

$$ecmLPC_t = LPC_t - 0.11(LW - LETA)_t - 0.89LEX_t \qquad (3.6)$$

showing the relative predominance of the exchange rate effects in explaining the domestic price level.

Although the cointegrating vectors considered separtely are identified, we can test the overidentifying restrictions on them. At first we have many exclusion restrictions, namely the absence of wage, price and exchange rate effects on productivity, no exchange rate and private employment effects on wages and the absence of population and private employment effects on domestic prices. In addition there are other three more parameter restrictions: unit elasticity of consumer price in wage equation, the same parameter value (with inverted sign) for wages and productivity in consumer price equation, and, always in the same equation, the sum of labour cost and exchange rate parameters equal to unity. All these overidentifying restrictions considered jointly are not rejected at the 10% significance level (see the note at the bottom of table 3). Since the null assumptions are linear on an I(0) parametrization of the model, in the sense that if $x_t \sim I(1)$ and there is cointegration between the elements of x_t then $\Delta x_t \sim I(0)$ and $\beta' x_t \sim I(0)$, the test statistic useful for testing these restrictions is the likelihood ratio having a limiting χ^2 distribution under the null, with a number of degrees of freedom equal to the number of independent restrictions to be tested. The empirical evidence on the restrictions enables us to consider the whole system as congruent and consequently the full sample estimates can be viewed as consistent with the existence of long run equilibrium relationships of the form (3.4), (3.5) and (3.6).

Given the restrictions on the loading parameters, the $LNPOP_t$ variable can be classified as weakly exogenous. In the analysis of the structural model we did not marginalise with respect to this variable because of the rejection of the Granger-noncausality test (i.e. the past variations of the other variables in the system give a significant contribution in explaining the present variations in $LNPOP_t$, as it appears in the last equation of table 4).

Defining with ECM_t the three dimension vector of the the identified long run *ecm* terms, the whole system parametrization can be obtained by mapping from $I(1)$ to $I(0)$ space and using the general formulation:

$$\Delta x_t = \sum_{i=1}^{5} \Pi_i \Delta x_{t-i} + \alpha ECM_{t-1} + \gamma D_t + \varepsilon_t \tag{3.7}$$

where the trend variable has been removed from D_t. Estimation results of the system are reported in table 4; these are obtained by adopting a strong semplification on the whole system, namely by reducing the number of parameters to be estimated imposing 52 zero restrictions on the short run parameters, which are all accepted.

Diagnostic tests of the model are reported in table 5, while in figures 3a and 3b are presented the associated residual plots, correlograms and frequency distributions. This important reduction enables us to refer to such a model as a parsimonious VAR (PVAR); even if there is still room for simplifying further this parsimonious model since some of the estimated coefficients, as presented in table 4, appear not to be significantly different from zero, this model specification is retained as the basic framework within which to test important structural assumptions.

Table 4. Poland: Simplified model FIML estimates (sample 1971.1-1995.3)*

$$\mathrm{DLETA} = \underset{(.067)}{0.299}\ \mathrm{DLW}(-3) + \underset{(.043)}{0.167}\ \mathrm{DLPC}(-1) + \underset{(.051)}{0.127}\ \mathrm{DLPC}(-3) + \underset{(.038)}{0.08}\ \mathrm{DLPC}(-4) +$$

$$- \underset{(.034)}{0.11}\ \mathrm{DLEX}(-1) - \underset{(.041)}{0.11}\ \mathrm{DLEX}(-2) - \underset{(.03)}{0.08}\ \mathrm{DLEX}(-3) - \underset{(.075)}{0.56}\ \mathrm{ecmLETA}(-1) +$$

$$- \underset{(.039)}{0.139}\ \mathrm{ecmLPC}(-1) + \text{deterministic terms}^{**}$$

$$\mathrm{DLW} = \underset{(.091)}{0.195}\ \mathrm{DLETA}(-1) + \underset{(.081)}{0.274}\ \mathrm{DLETA}(-2) + \underset{(.058)}{0.238}\ \mathrm{DLW}(-2) + \underset{(.051)}{0.383}\ \mathrm{DLW}(-4) +$$

$$\underset{(.069)}{0.076}\ \mathrm{DLPC}(-3) + \underset{(.037)}{0.052}\ \mathrm{DLEX}(-2) + \underset{(.032)}{0.121}\ \mathrm{DLEX}(-3) + \underset{(.38)}{0.95}\ \mathrm{DLNPOP}(-2) +$$

$$- \underset{(.030)}{0.94}\ \mathrm{DLNPOP}(-3) - \underset{(.56)}{0.67}\ \mathrm{DLNPOP}(-4) - \underset{(.054)}{0.25}\ \mathrm{ecmLETA}(-1) +$$

$$- \underset{(.026)}{0.094}\ \mathrm{ecmLW}(-1) - \underset{(.033)}{0.068}\ \mathrm{ecmLPC}(-1) + \text{deterministic terms}^{**}$$

$$\mathrm{DLPC} = \underset{(.056)}{0.156}\ \mathrm{DLETA}(-1) + \underset{(.052)}{0.156}\ \mathrm{DLETA}(-2) + \underset{(.045)}{0.06}\ \mathrm{DLW}(-2) + \underset{(.045)}{0.292}\ \mathrm{DLPC}(-1) +$$

$$\underset{(.029)}{0.046}\ \mathrm{DLPC}(-2) + \underset{(.026)}{0.062}\ \mathrm{DLPC}(-4) + \underset{(.03)}{0.07}\ \mathrm{DLEX}(-1) + \underset{(.031)}{0.121}\ \mathrm{DLEX}(-2) +$$

$$- \underset{(.228)}{0.52}\ \mathrm{DLNPOP}(-2) + \underset{(.211)}{1.03}\ \mathrm{DLNPOP}(-4) - \underset{(.019)}{0.082}\ \mathrm{ecmLETA}(-1) +$$

$$\underset{(.031)}{0.28}\ \mathrm{ecmLW}(-1) - \underset{(.025)}{0.043}\ \mathrm{ecmLPC}(-1) + \text{deterministic terms}^{**}$$

$$\mathrm{DLEX} = - \underset{(.186)}{0.256}\ \mathrm{DLETA}(-1) + \underset{(.161)}{0.665}\ \mathrm{DLETA}(-2) + \underset{(.143)}{0.516}\ \mathrm{DLW}(-2) - \underset{(.11)}{0.465}\ \mathrm{DLPC}(-1) +$$

$$- \underset{(.079)}{0.182}\ \mathrm{DLPC}(-2) + \underset{(0.1)}{0.125}\ \mathrm{DLPC}(-3) - \underset{(.077)}{0.305}\ \mathrm{DLPC}(-4) + \underset{(.102)}{0.308}\ \mathrm{DLEX}(-1) +$$

$$\underset{(.106)}{0.473}\ \mathrm{DLEX}(-2) + \underset{(.081)}{0.354}\ \mathrm{DLEX}(-3) + \underset{(.116)}{0.78}\ \mathrm{ecmLETA}(-1) + \underset{(.102)}{0.76}\ \mathrm{ecmLW}(-1) +$$

$$\underset{(.112)}{0.74}\ \mathrm{ecmLPC}(-1) + \text{deterministic terms}^{**}$$

$$\mathrm{DLNPOP} = \underset{(.016)}{0.025}\ \mathrm{DLETA}(-1) + \underset{(.016)}{0.032}\ \mathrm{DLETA}(-2) + \underset{(.013)}{0.033}\ \mathrm{DLETA}(-3) +$$

$$- \underset{(.011)}{0.048}\ \mathrm{DLW}(-3) - \underset{(.012)}{0.056}\ \mathrm{DLW}(-4) - \underset{(.005)}{0.025}\ \mathrm{DLPC}(-1) + \underset{(.009)}{0.015}\ \mathrm{DLPC}(-2) +$$

$$- \underset{(.009)}{0.02}\ \mathrm{DLPC}(-3) + \underset{(.009)}{0.029}\ \mathrm{DLPC}(-4) - \underset{(.006)}{0.013}\ \mathrm{DLEX}(-1) - \underset{(.006)}{0.008}\ \mathrm{DLEX}(-2) +$$

$$- \underset{(.005)}{0.01}\ \mathrm{DLEX}(-3) + \underset{(.059)}{0.089}\ \mathrm{DLNPOP}(-2) + \underset{(.064)}{0.248}\ \mathrm{DLNPOP}(-3) +$$

$$\underset{(.064)}{0.336}\ \mathrm{DLNPOP}(-4) + \text{deterministic terms}^{**}$$

(*) Heteroschedastic consistent standard errors in brackets.

(**) Details about the deterministic components of the model are listed in Annex 1.

Table 5. Poland: Simplified model diagnostic statistics*

	DLETA	DLW	DLPC	DLEX	DLNPOP
ser	3.65%	3.52%	2.59%	7.36%	0.54%
AR 1-5, F(5, 58)	2.504	3.138	3.335	1.792	3.300
	[.041]	[.014]	[.011]	[.129]	[.011]
ARCH 4, F(4, 55)	0.630	0.167	0.739	0.315	0.124
	[.643]	[.955]	[.570]	[.867]	[.973]
Normality, $\chi_2(2)$	5.131	1.864	0.713	4.443	1.043
	[.077]	[.394]	[.700]	[.109]	[.594]

Vector statistics: AR 1-5, F(125, 226) = 1.252 [.0731]; Normality, $\chi_2(10)$ = 12.31 [.265].
LR test of overidentifying restrictions: $\chi_2(52)$ =52.01 [.473]
(*) Each column corresponds to a single equation diagnostics. P-values are in brackets

In order to analyse the structural constancy of the system parametrisation adopted, we perform the recursive FIML model estimation. The analysis of the residuals of such estimation is reported in figure 4. The recursive residual plots show that they are always included in the $\pm 2\ SE$ bands. Chow parameter constancy tests for the period 1993-1995 accept the assumption of no structural breaks.

Finally, always on the basis of the whole model estimation results reported in table 4, we do an impulse-response exercise by imposing a shock of one on the $DLEX_t$ equaton. The results produced by such a shock are presented in form of graphs in figure 5. As we can see, a unit shock on the exchange rate produces, in the long run, a permanent rise on both wages and domestic prices. If we compare the relative importance of such increases, we note that the rise in prices is relatively bigger that the rise in wages, with the consequence that, in the long run, this shock produces a reduction in real wages. On the other side, we can see that both labour productivity and employment-population ratio seem to be negatively affected, but not in a significative way.

3.2. Hungary

A similar model for inflation, exchange rate and employment is used for analysisng the Hungarian economy, always for the period 1971.1 - 1995.3. The main difference with the Polish case is due to the fact that in the data set we used for Hungary we included an additional (exogenous) indicator of international manufacturing price level ($LPM59$), measured by the international price in US dollars for manufactured goods (source OECD, SITC codes 5-9). All the variables considered exhibit prounounced trends as well as cyclical movements, so they need to be modelled as variables with stochastic trends within a cointegrated-integrated framework where changes in economic policies and other autonomous events may play an important role in producing cyclical features.

The integration analysis was conducted, as in Golinelli and Orsi (1994), by using the testing procedure proposed by Zivot and Andrews (1992). In the light of the results obtained, all the series analysed seem be generated by an I(1) process. The periods in which significant structural breaks occurred are identified as the 1987.4 (mainly for wages) and 1990.1 (mainly for prices). This relevant information has been used in the dynamic specification of the whole system by introducing two dummies taking specific account of these structural breaks.

The subsequent step of our analysis is the estimation of an UVAR model of order 5. This VECM formulation reveals to be congruent, though there is some evidence of residual serial correlation of 5th order in $DLEX_t$. Subsequent analysis reveal that a possible explanation of such problem is not an inadequate lag length, but rather the system inadequacy to catch the many changes in Government and economic structure of the country. However, we decide not to introduce more specific dummies for such events in order to avoid the risk that the system is too finely tuned to the sample data. Other test statistics used for detecting the presence of eteroscedasticity, normality and parameter constancy did not reject the null. Since the joint exclusion of both the linear trend and $LPRIV_t$ variables from the system was not rejected at a significance level less than 10%, another UVAR system, where the trend and $LPRIV_t$ variables are omitted, is estimated the corresponding diagnostic statistics are reported in table 6.

Table 6. Hungary: UVAR diagnostic statistics*

	DLETA	DLW	DLPC	DLEX	DLNPOP
ser	2.57%	2.78%	1.33%	7.34%	0.58%
AR 1-5, F(5, 45)	2.093	1.795	0.472	2.014	2.565
	[.081]	[.130]	[.795]	[.092]	[.038]
ARCH 4, F(4, 42)	0.128	0.199	2.375	0.379	0.230
	[.972]	[.938]	[.066]	[.822]	[.920]
Normality, $\chi_2(2)$	1.440	3.579	1.027	0.590	2.725
	[.487]	[.167]	[.599]	[.745]	[.256]

Vector statistics: AR 1-5, F(125, 142) = 1.473 [.013]; Normality, $\chi_2(10)$ = 12.91 [.229]
(*) Each column corresponds to a single equation diagnostics. P-values are in brackets

In addition, the set of plots presented in figures 6 and 7 provides empirical evidence in favour of an UVAR system with normal white noise residuals and constant parameters.

The cointegration analysis of the system is reported in table 7. The results are not clear-cut: while the empirical evidence provided by the *max* test suggests an $r = 1$ (or $r = 2$ if we use a larger significance level), the *trace* test indicates a full rank matrix, i.e. $r = 5$.

Table 7. Hungary: Cointegration statistics

rank:	1	2	3	4	5
log-likelihood	2015.34	2027.42	2036.14	2043.38	2048.46
eigenvalue μ	0.40	0.22	0.16	0.13	0.098
Max eigenvalue test	50.2*	24.2	17.4	14.5	10.2*
Max eigenvalue test[a]	37.4	18.0	13.0	10.8	7.6*
95% critical values	33.5	27.1	21.0	14.1	3.8
Trace test	116.4*	66.3*	42.1*	24.7*	10.2*
Trace test[a]	86.7*	49.4	31.4	18.4	7.6*
95% critical values	68.5	47.2	29.7	15.4	3.8

[a] Osterwald-Lenum (1992) definition.
(*) 1% significant.

With the aim of enlarging the available information useful for identifying the cointegration relationships, the plots of the fitted and actual values, for both the cointegrating vectors and the remaining non stationary components, are reported in figures 8a and 8b. More precisely, figure 8a plot the estimated disequilibria $\beta' x_t$ for cointegrating vectors and non-stationary components. Disequilibria in wages seem not too important until 1990, while after that period start to become more pronounced. A rater similar behaviour is revealed by domestic prices, where large disequilibria are observed starting from 1989. On the contrary, for both productivity ($LETA_t$) and $LNPOP_t$ variables the estimated long run relationships clerly show a non stationary behaviour. Finally, important disequilibria are observed for the exchange rate vector, specially after the important changes occurred at the end of the 80s, but the existence of a cointegrating relationship seems uncertain. In figure 8b, we can see that actual and fitted values for both LW_t and LPC_t follow very closely, while in the case of $LETA_t$ and $LNPOP_t$ they tend to diverge, confirming the non stationarity of these two cointegration combinations. More uncertain is the case of LEX_t since the plot of actual and fitted values for LEX_t does not provide a right view of things. The empirical values of the cointegrating test statistics, along with the visual inspection of plots in figures 8a and 8b, suggest a dimension of the cointegration space equal to 2 (r=2),(see table 7). On the basis of that, we expect two identified long run relationships which are reproduced, in an implicit form, in table 8, namely $ecmLW_t$ and $ecmLPC_t$. They correspond to the identified cointegration vectors which reveal to reproduce a stationary path, while the remaining three cointegrating combinatons are those that reveal to be non stationary. In a previous attempt, not reported here, we discover that the inclusion of $LNPOP_t$ variable in the LW_t long-run specification was not significant, since at the parameter estimate of -.58 was associated a standard error of .36.

Table 8. Hungary: Identified long run relations (standard errors in brackets)*

		LETA	LW	LPC	LEX	LNPOP	LPM59
ecmLW	=	-0.920	1.000	-0.658	-	-	-
		(0.092)	(-)	(0.054)			
ecmLPC	=	0.382	-0.382	1.000	-0.618	-	-0.618
		(-)	(-)	(-)	(-)		(0.062)

(*) Overidentifying LR test: $\chi_2(8) = 17.215$ [.028]; including 2 zero restrictions on the two loadings in DLNPOP equation.

The overidentifying long-run restrictions reveal to be significant at a level of 5%. Important restrictions are represented by the omission of both $ecmLW_t$ and $ecmLPC_t$ terms from the $DLNPOP_t$ equation; without the first restriction on the loading parameter the overidentifying LR test is given by $\chi^2(7) = 13.634$ [.058]. Given both restrictions on the loading parameters, the $LNPOP_t$ variable can be classified as weakly exogenous also for Hungary . The two equilibrium correction terms, representing the long run relationships, are reported in table 8.

According to the simplification phase adopted for the cointegrated VAR model, we exclude from the system 7 explanatory variables, corresponding to 35 parameters, on the basis of an F significance test $F(35, 242) = 0.81[.77]$. The obtained PVAR system appear to be congruent and, after a further reduction in the number of parameters, the whole simplified FIML estimates are reported in table 9. As before, the list of deterministic components of the model for Hungary are listed in the annex 2. The additional information for the model given by diagnostic tests, as reported in table 10, provide further evidence that the simplified model performs in a satisfactory way, as well as the congruent system. Figures 9 and 10 tend to confirm this view.

Table 9. Hungary: Simplified model FIML estimates*

$$\text{DLETA} = \underset{(.083)}{-0.406}\,\text{DLETA(-1)} \underset{(.08)}{-0.3}\,\text{DLETA(-2)} \underset{(.067)}{+0.3}\,\text{DLETA(-4)} \underset{(.053)}{-0.243}\,\text{DLW(-3)} +$$

$$\underset{(.054)}{-0.209}\,\text{DLW(-4)} \underset{(.029)}{-0.063}\,\text{DLEX(-1)} \underset{(.016)}{+0.076}\,\text{ecmLPC(-1)} + \text{determ. terms**}$$

$$\text{DLW} = \underset{(.01)}{-0.56}\,\text{DLW(-1)} \underset{(.118)}{-0.347}\,\text{DLW(-2)} \underset{(.115)}{-0.352}\,\text{DLW(-3)} \underset{(.084)}{+0.205}\,\text{DLW(-4)} +$$

$$\underset{(.153)}{-0.616}\,\text{DLPC(-1)} \underset{(.174)}{-0.283}\,\text{DLPC(-2)} \underset{(.032)}{-0.098}\,\text{DLEX(-1)} \underset{(.177)}{-0.703}\,\text{DLNPOP(-4)} +$$

$$\underset{(.063)}{0.103}\,\text{DLPM59} + \underset{(.061)}{-0.229}\,\text{ecmLW(-1)} \underset{(.027)}{+0.1}\,\text{ecmLPC(-1)} + \text{determ. terms**}$$

$$\text{DLPC} = \underset{(.05)}{-0.118}\,\text{DLETA(-1)} \underset{(.046)}{-0.189}\,\text{DLETA(-2)} \underset{(.034)}{-0.105}\,\text{DLW(-2)} \underset{(.040)}{-0.091}\,\text{DLW(-2)} +$$

$$\underset{(.037)}{-0.131}\,\text{DLW(-4)} \underset{(.088)}{-0.212}\,\text{DLPC(-1)} \underset{(.089)}{-0.146}\,\text{DLPC(-2)} \underset{(.095)}{-0.166}\,\text{DLPC(-3)} +$$

$$\underset{(.021)}{-0.047}\,\text{DLEX(-1)} \underset{(.073)}{+0.228}\,\text{DLNPOP(-1)} \underset{(.083)}{+0.145}\,\text{DLNPOP(-4)} +$$

$$\underset{(.031)}{-0.078}\,\text{DLPM59} \underset{(.014)}{-0.053}\,\text{ecmLPC(-1)} + \text{deterministic terms**}$$

$$\text{DLEX} = \underset{(.146)}{-0.257}\,\text{DLETA(-2)} \underset{(.126)}{-0.734}\,\text{DLETA(-3)} \underset{(.127)}{+0.325}\,\text{DLW(-1)} \underset{(.091)}{-0.224}\,\text{DLEX(-1)} +$$

$$\underset{(.41)}{1.83}\,\text{DLNPOP(-1)} \underset{(.182)}{-0.783}\,\text{DLPM59} +$$

$$\underset{(.112)}{-0.311}\,\text{ecmLW(-1)} \underset{(.074)}{+0.245}\,\text{ecmLPC(-1)} + \text{deterministic terms**}$$

$$\text{DLNPOP} = \underset{(.027)}{0.04}\,\text{DLETA(-1)} \underset{(.024)}{+0.09}\,\text{DLETA(-2)} \underset{(.019)}{+0.05}\,\text{DLETA(-3)} \underset{(.017)}{+0.02}\,\text{DLETA(-4)}$$

$$\underset{(.017)}{+0.038}\,\text{DLW(-1)} \underset{(.018)}{+0.027}\,\text{DLW(-2)} \underset{(.018)}{+0.04}\,\text{DLW(-3)} \underset{(.019)}{+0.075}\,\text{DLW(-4)} +$$

$$\underset{(.031)}{-0.061}\,\text{DLPC(-2)} \underset{(.035)}{-0.07}\,\text{DLPC(-3)} \underset{(.051)}{+0.25}\,\text{DLNPOP(-4)} \underset{(.018)}{-0.024}\,\text{DLPM59} +$$

$$\text{deterministic terms**}$$

(*) Heteroschedastic consistent standard errors in brackets.

(**) Details about the deterministic components of the model are listed in Annex 2.

Table 10. Hungary: Simplified model diagnostic statistics*

	DLETA	DLW	DLPC	DLEX	DLNPOP
ser	2.44%	2.57%	1.32%	7.20%	0.62%
AR 1-5, F(5, 62)	2.548	2.577	1.389	3.319	2.958
	[.037]	[.035]	[.024]	[.011]	[.019]
ARCH 4, F(4, 59)	0.116	0.742	0.258	1.036	0.455
	[.976]	[.567]	[.904]	[.397]	[.768]
Normality, $\chi_2(2)$	6.551	4.769	0.530	0.155	3.218
	[.038]	[.092]	[.767]	[.925]	[.200]

Vector statistics: AR 1-5, F(125, 241) = 1.149 [.181]; Normality, $\chi_2(10)$ = 11.96 [.288].

LR test of overidentifying restrictions: $\chi_2(48)$ = 45.89 [.56]

(*) Each column corresponds to a single equation diagnostics. P-values are in brackets

Finally, the impulse response exercise reported in figure 11 does not show any significant long run effect for the variables considered, except the case of the exchange rate whereas the shock persists all over the period considered.

4. Some concluding remarks

In this paper we presented a model where the impact of exchange rate on inflation and employment for Poland and Hungary is analysed. The model, based on Johansen's maximum likelihood approach, permits to identify long-run behaviour in presence of important structural breaks.

The study about the two countries has been conducted on the basis of a common set of variables of interest. On the whole, in the short-run the behaviour of the two economies under scrutiny reveal to be rather similar. Both the models are almost overparametrized, and consequently the information set on which the models are based appears to be quite fragmented (too many short run parameters). If, on one hand, this is of importance in order to capture the structurl changes (statistical tests lead to the acceptance of parameter constancy), on the other hand such specification risks to be too finely tuned to sample data.

The empirical results, for the long-run, appear to be more interesting and give support to some important structural differences in prices and wages setting for the two countries. In the Polish model we have three significant long-run relationships. Besides the traditional price-wage spiral, we endogenize into this system the labour productivity, that appears to be significantly affected by the so-called labour market attrition. On the other side, no significant effect was played by international prices of manufactured goods.

Unlike for Poland, we did not succed in explaining the long run labour productivity in Hungary. For this country we just obtained a simple wage-price model where, in the long period, wages are more weakly linked to prices. In particular, in the Hungarian model, nominal wages tend to increase less than prices but, on the other side, they appear to be more strongly related to the labour productivity than in the Polish model; in fact, the productivity coefficient in wage equation is .92 for Hungary and .51 for Poland.

Finally, the exchange rate variable is significant in explainnig the domestic price behaviour for both countries, perhaps more in Poland than in Hungary, while the manufacturing international price variable reveals to be significant only for Hungary. These results may cause several comments; among others, the fact that Hungary seems to come to a market-economy oriented system a little bit early than Poland, even if this kind of evaluations deserves a deeper investigation.

Figure 1a. Poland:
Residuals and mispecification graphics for UVAR wage price and exchange rate equations

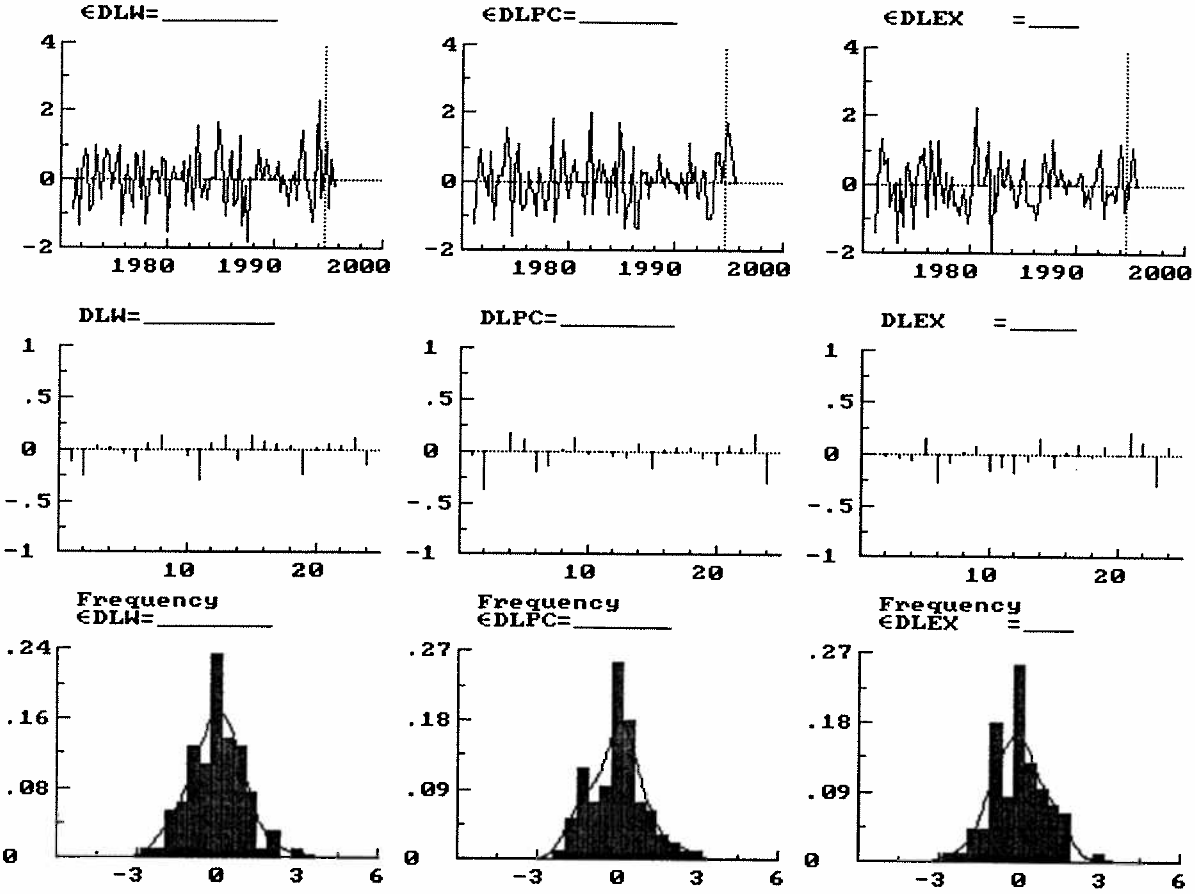


Figure 1b. Poland:
Residuals and mispecification graphics for UVAR productivity and employed-population equations

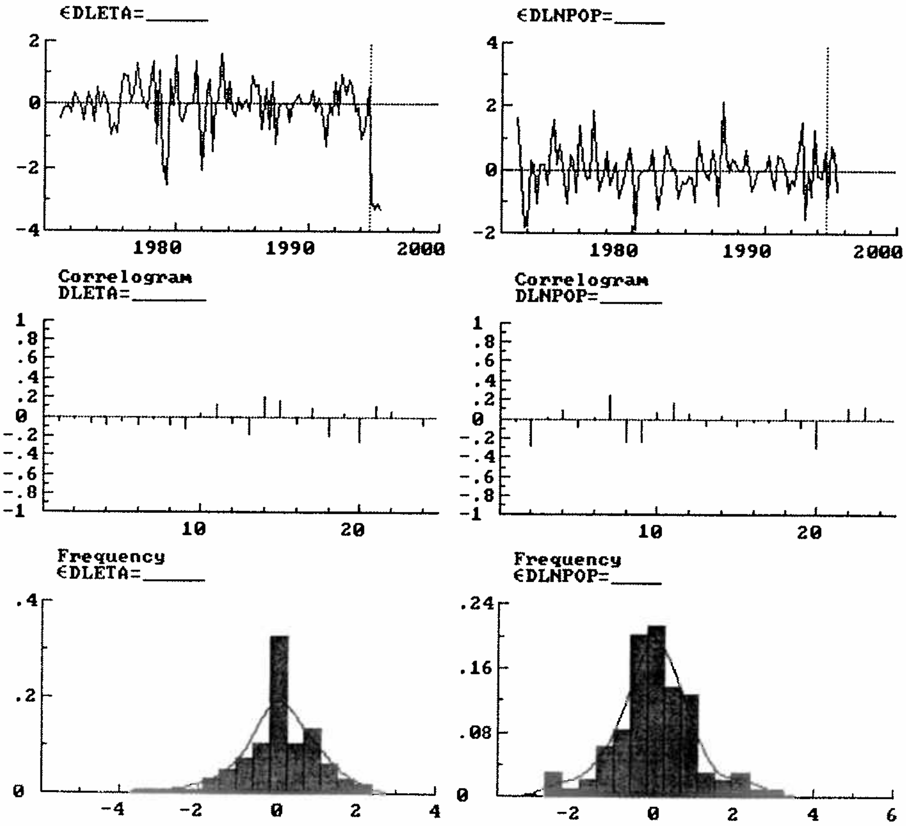


Figure 2a. Poland: UVAR 1-step residuals

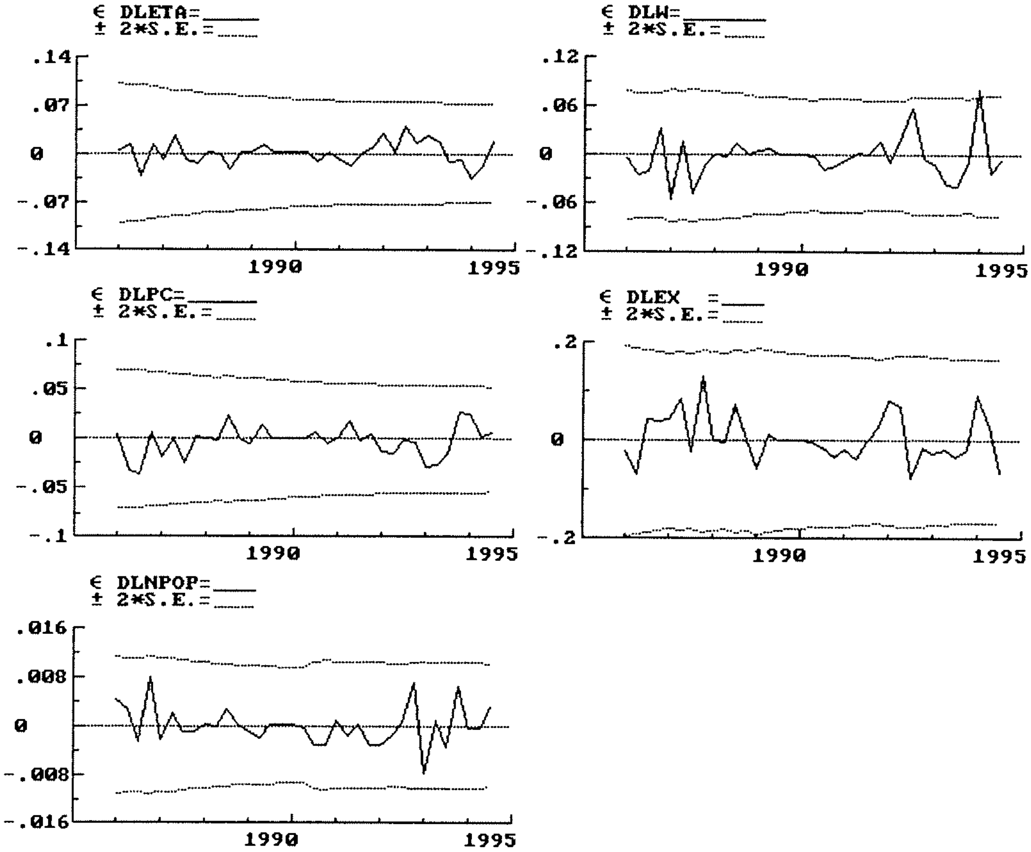


Figure 2b. Poland: UVAR structural stability tests

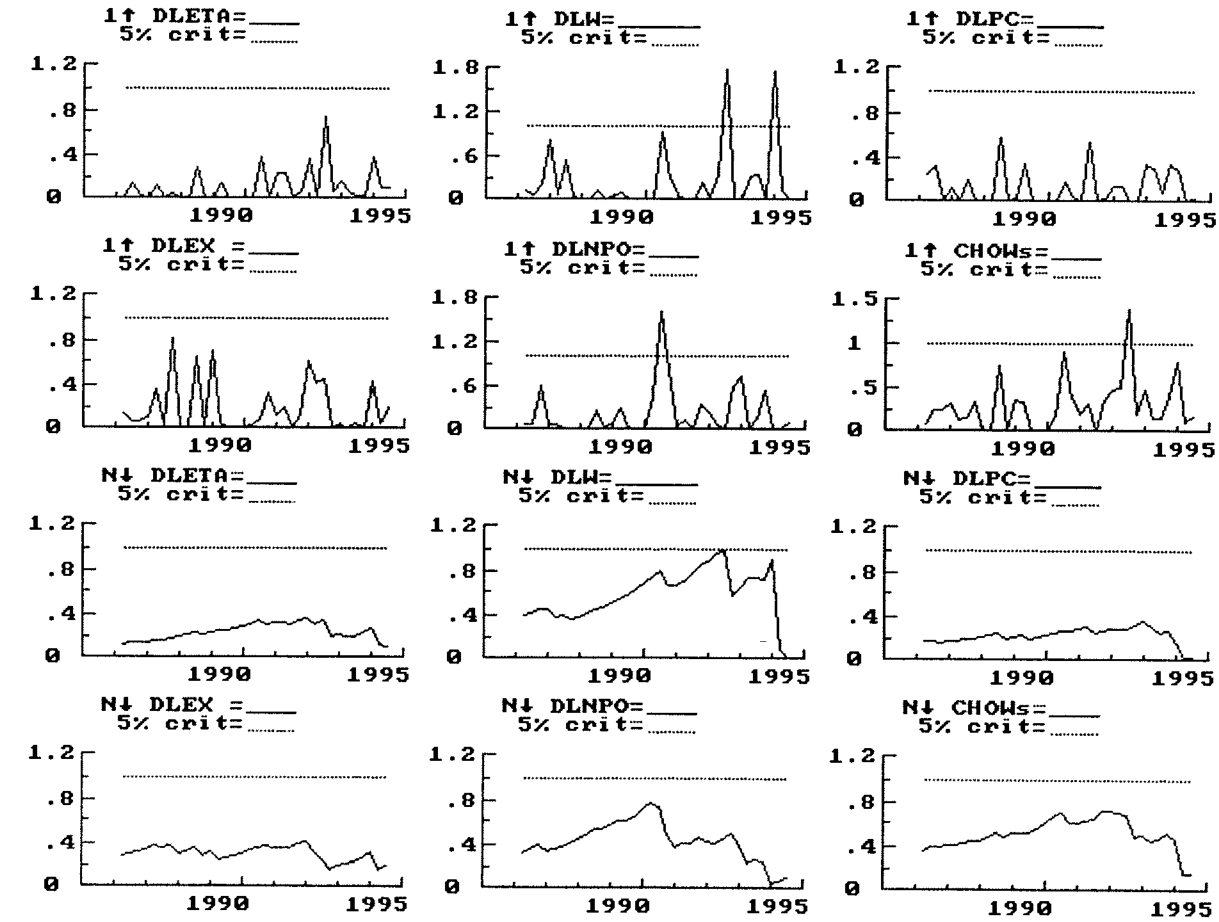


Figure 3a. Poland:
FIML residuals and mispecification graphics for wage, price and exchange rate equations

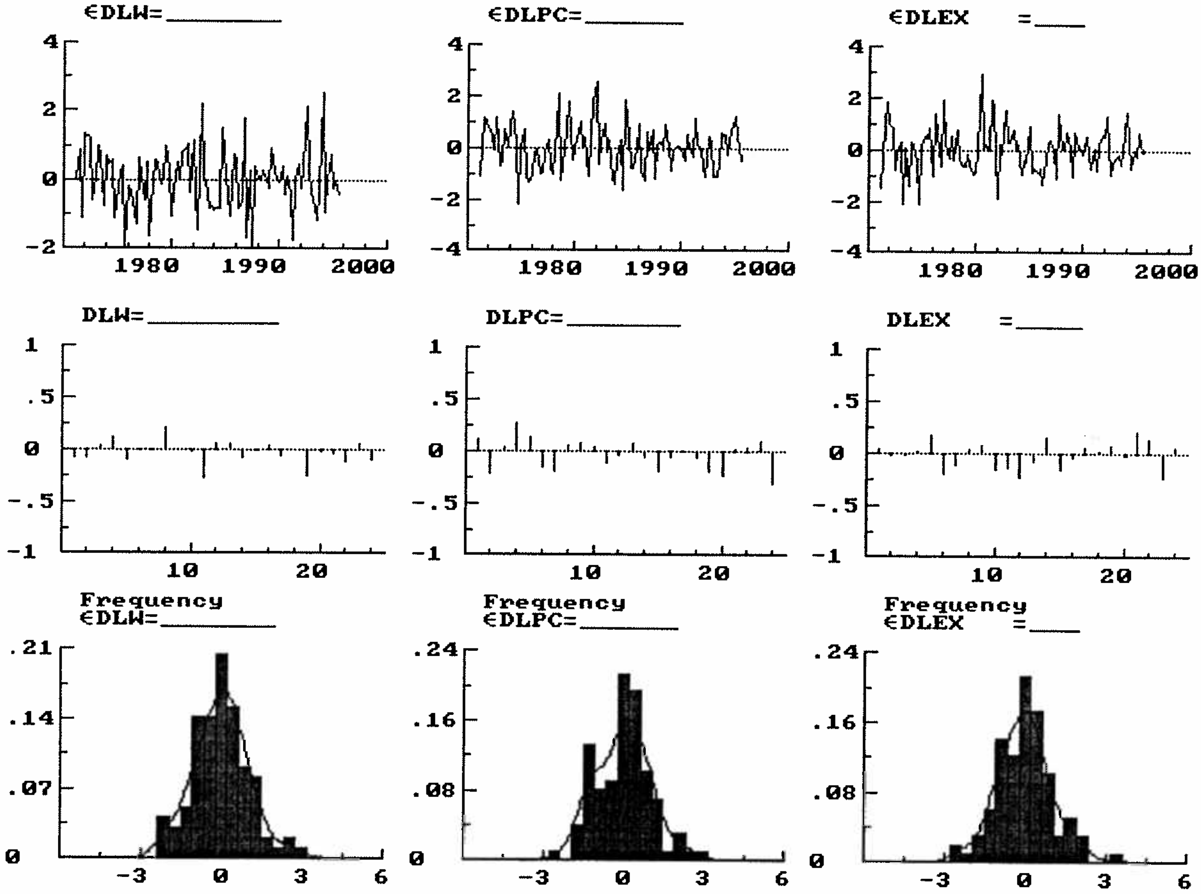


Figure 3b. Poland:
FIML residuals and mispecification graphics for productivity and employed-population equations

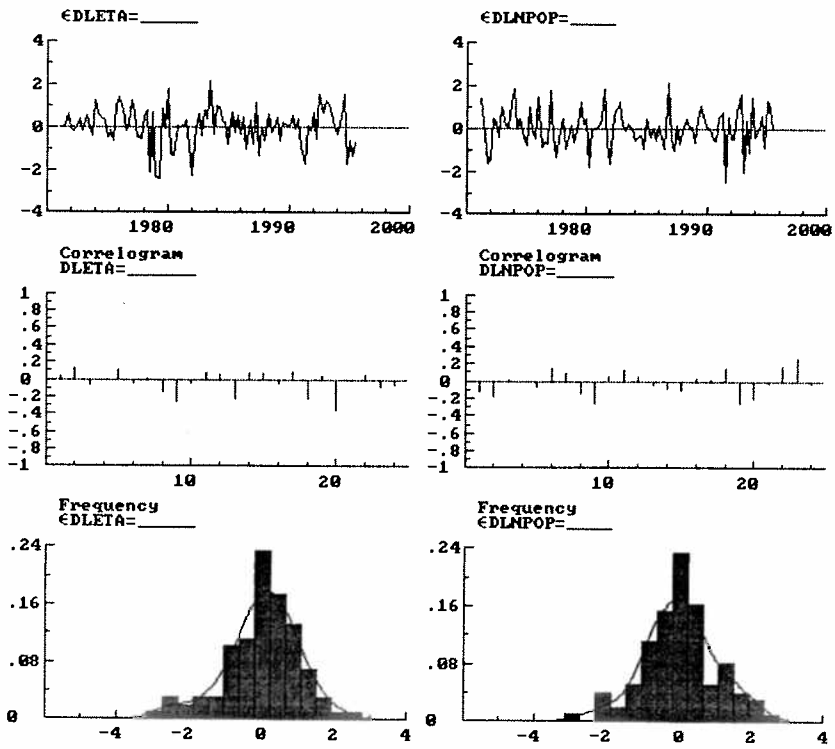


Figure 4. Poland: Stuctural stability of the simplified model

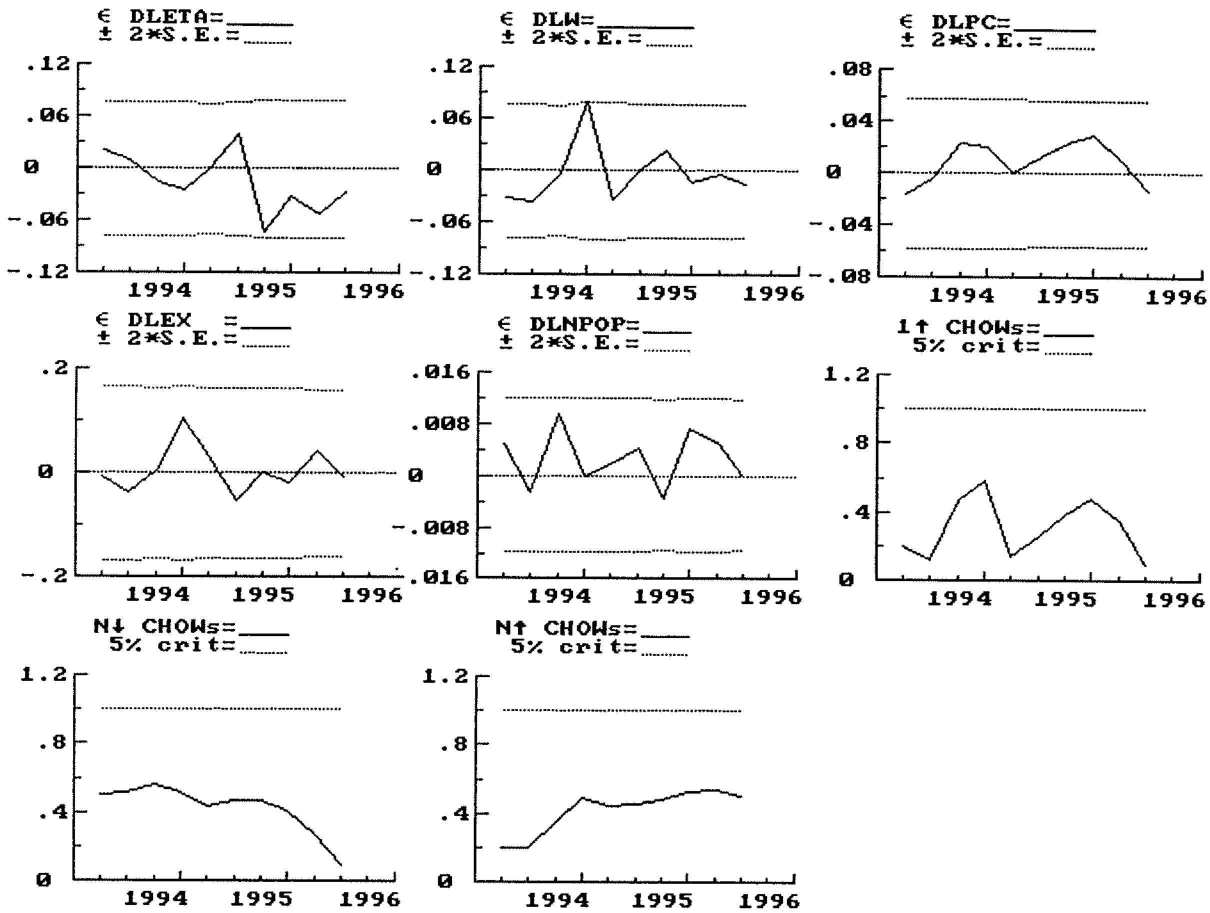


Figure 5.
Poland: Cumulated impulse-response function of a DLEX unit shock in the simplified model

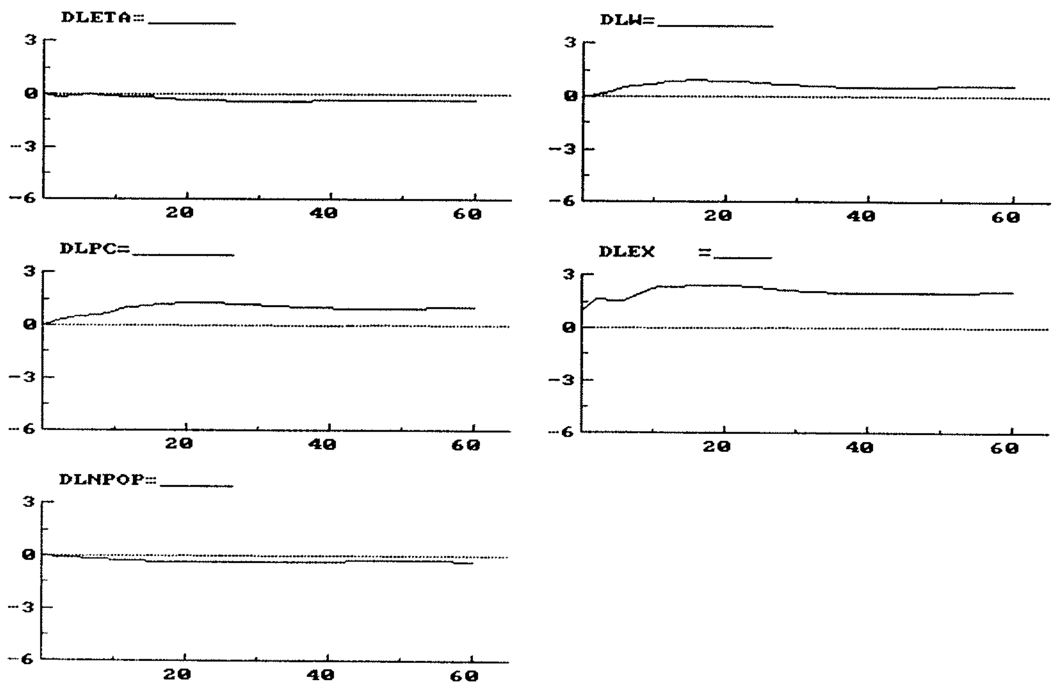


Figure 6a. Hungary:
Residuals and mispecification graphics for UVAR wage price and exchange rate equations

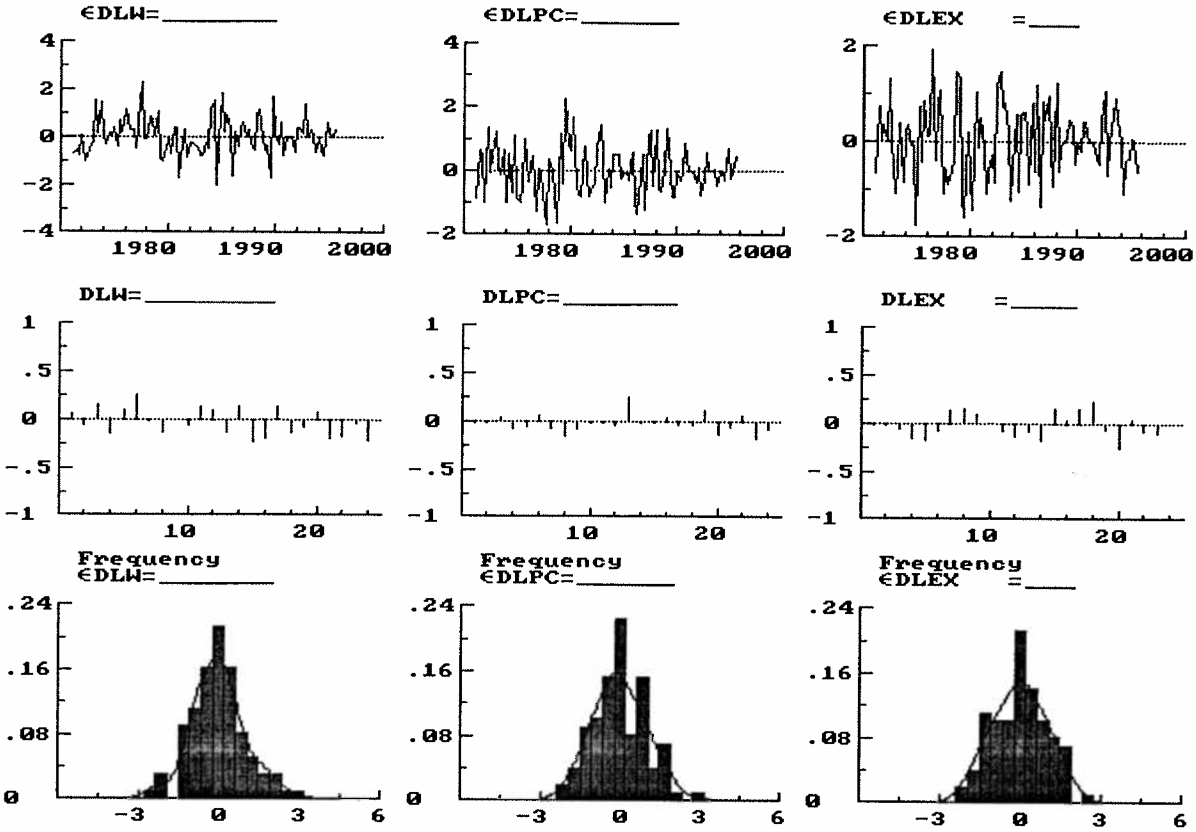


Figure 6b. Hungary:
Residuals and mispecification graphics for UVAR productivity and employed-population equations

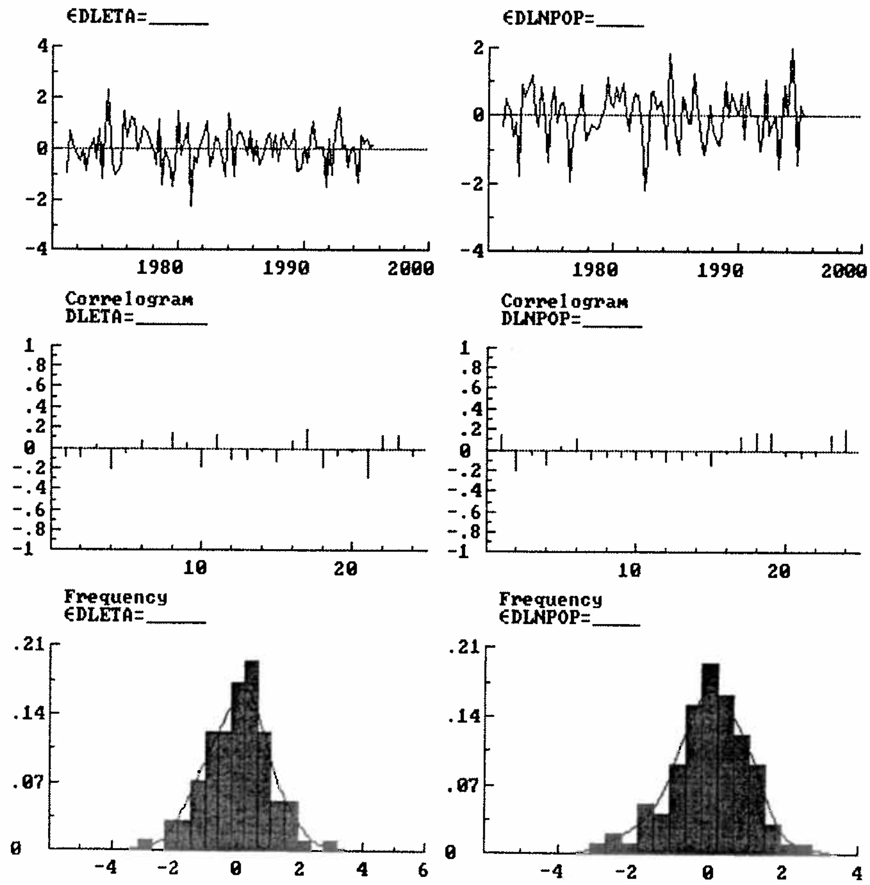


Figure 7a. Hungary: UVAR 1-step residuals

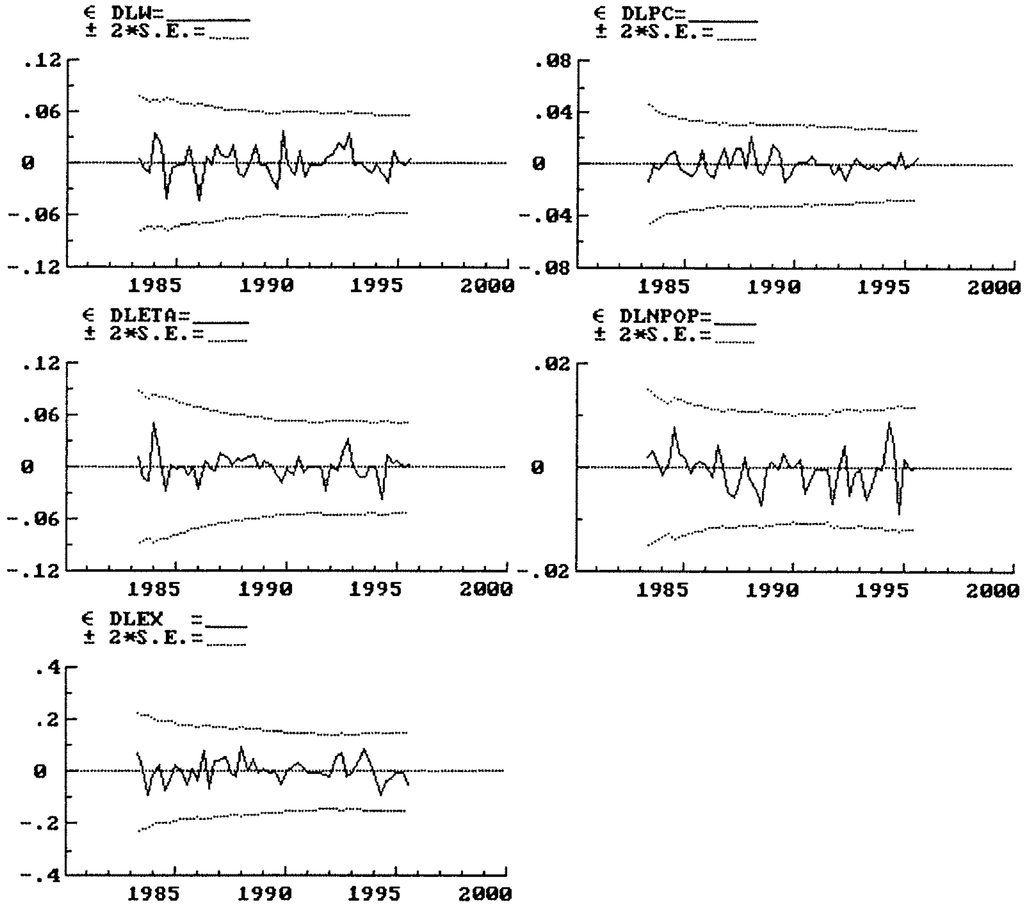


Figure 7b. Hungary: UVAR structural stability tests

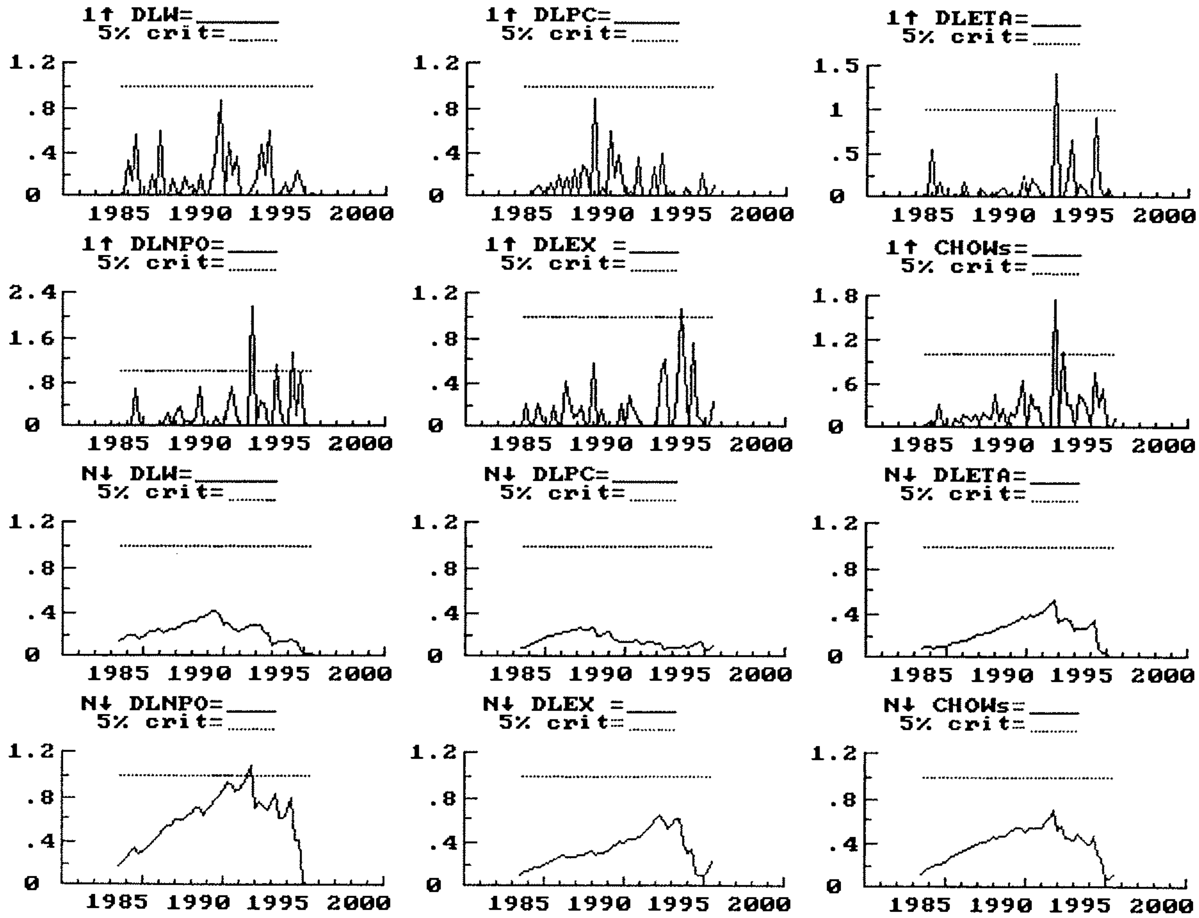


Figure 8a. Hungary: Estimated long-run (cointegrating) relationships

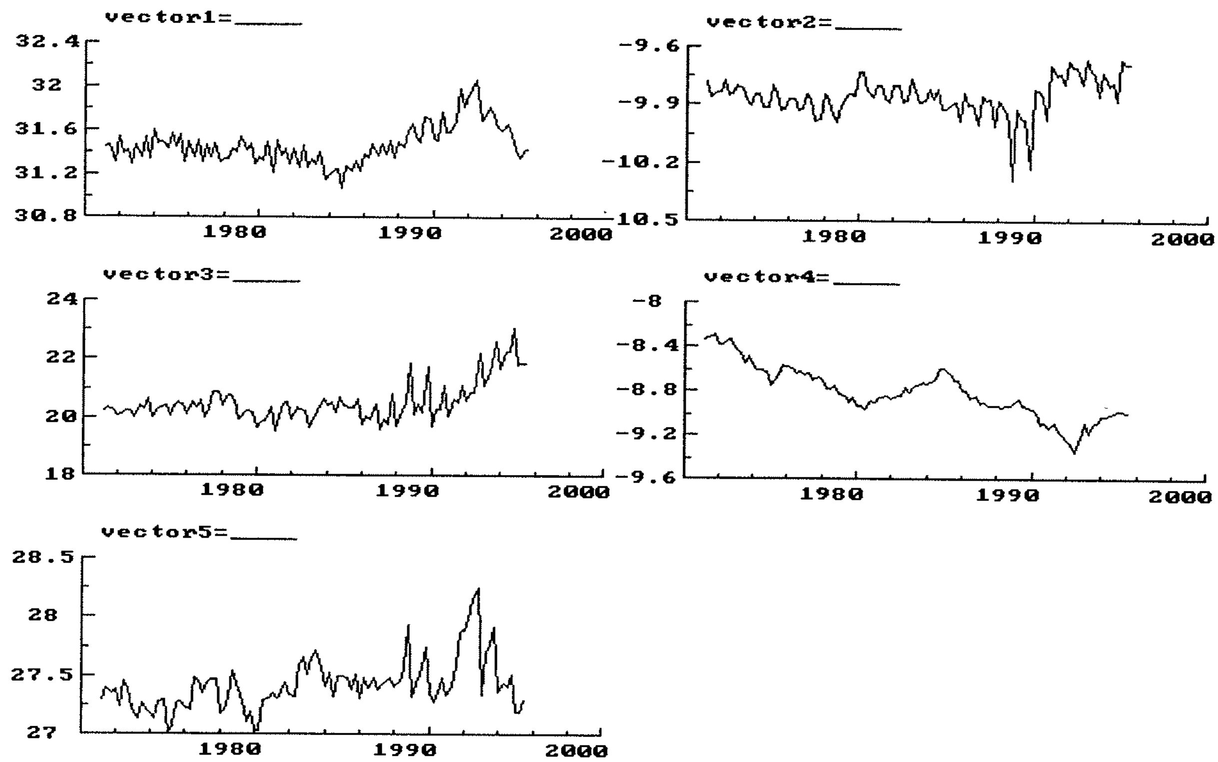


Figure 8b. Hungary: Actual and fitted values of cointegrating vectors

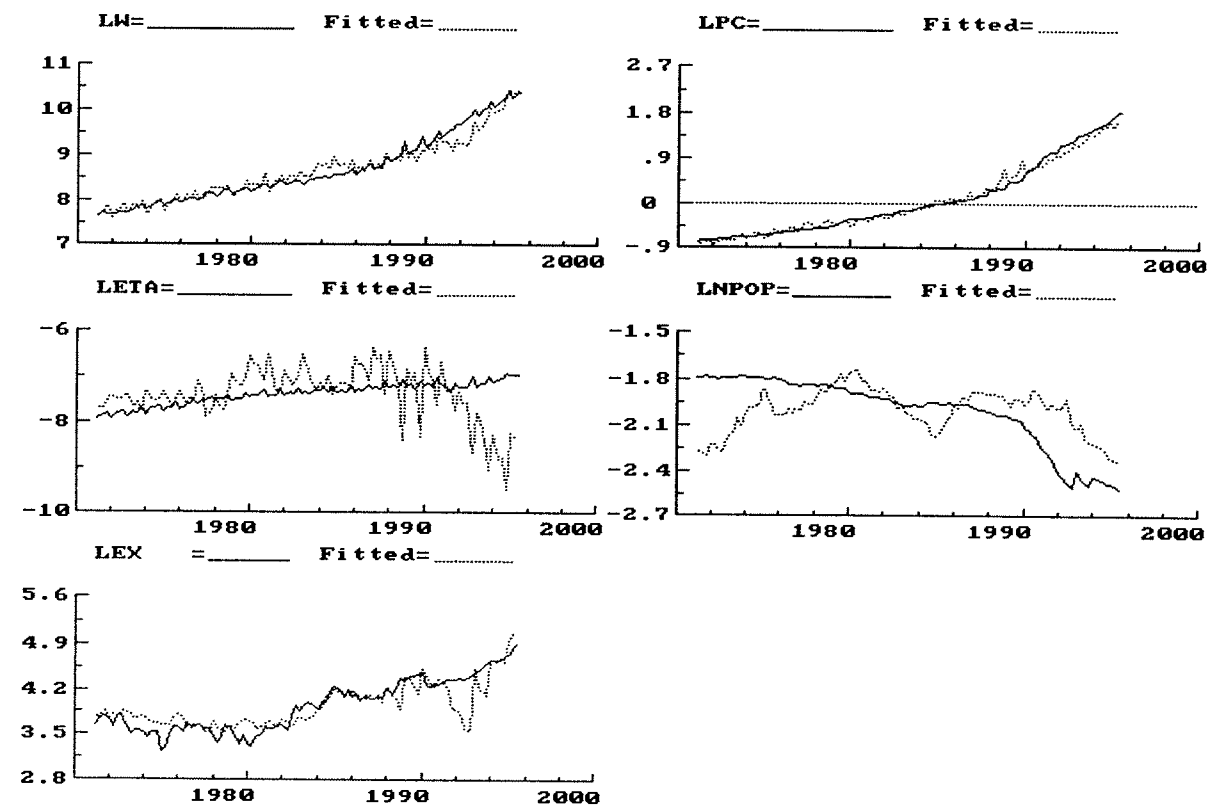


Figure 9a. Hungary:
FIML residuals and mispecification graphics for wage, price and exchange rate equations

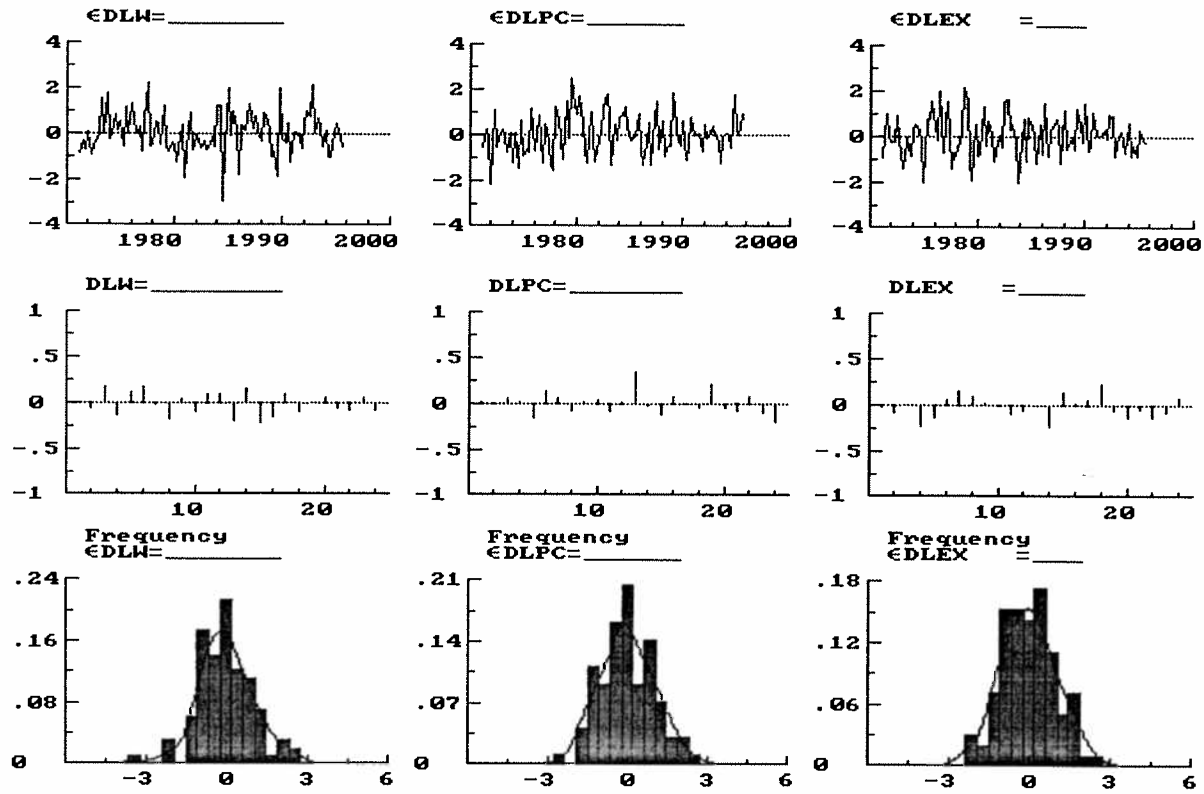


Figure 9b. Hungary:
FIML residuals and mispecification graphics for productivity and employed-population equations

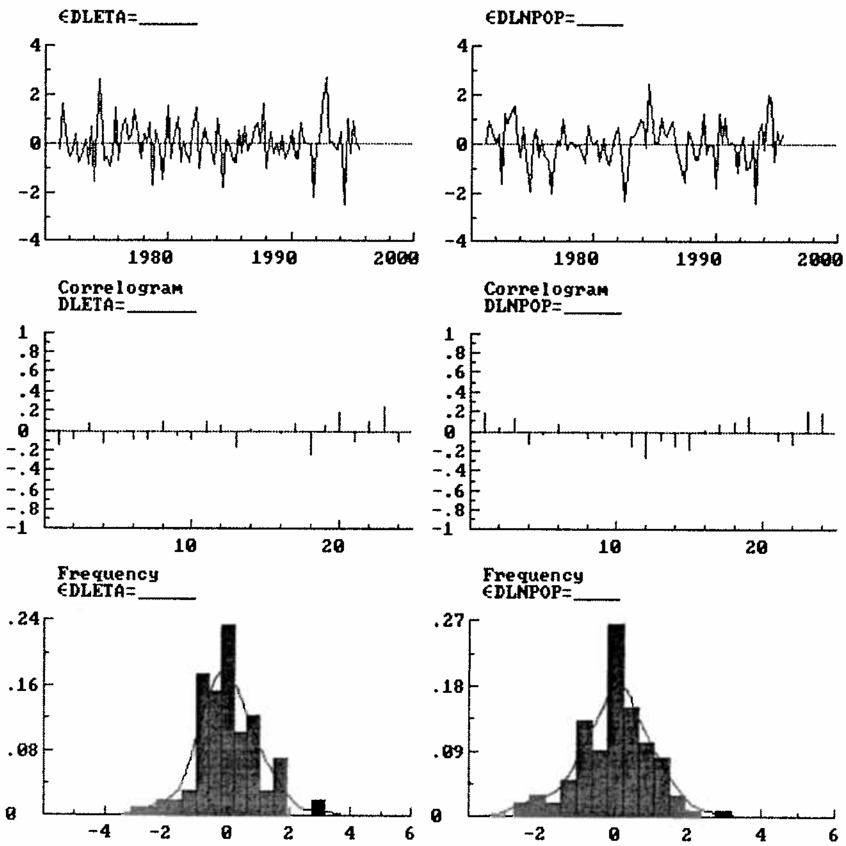


Figure 10. Hungary: Strucural stability of the simplified model

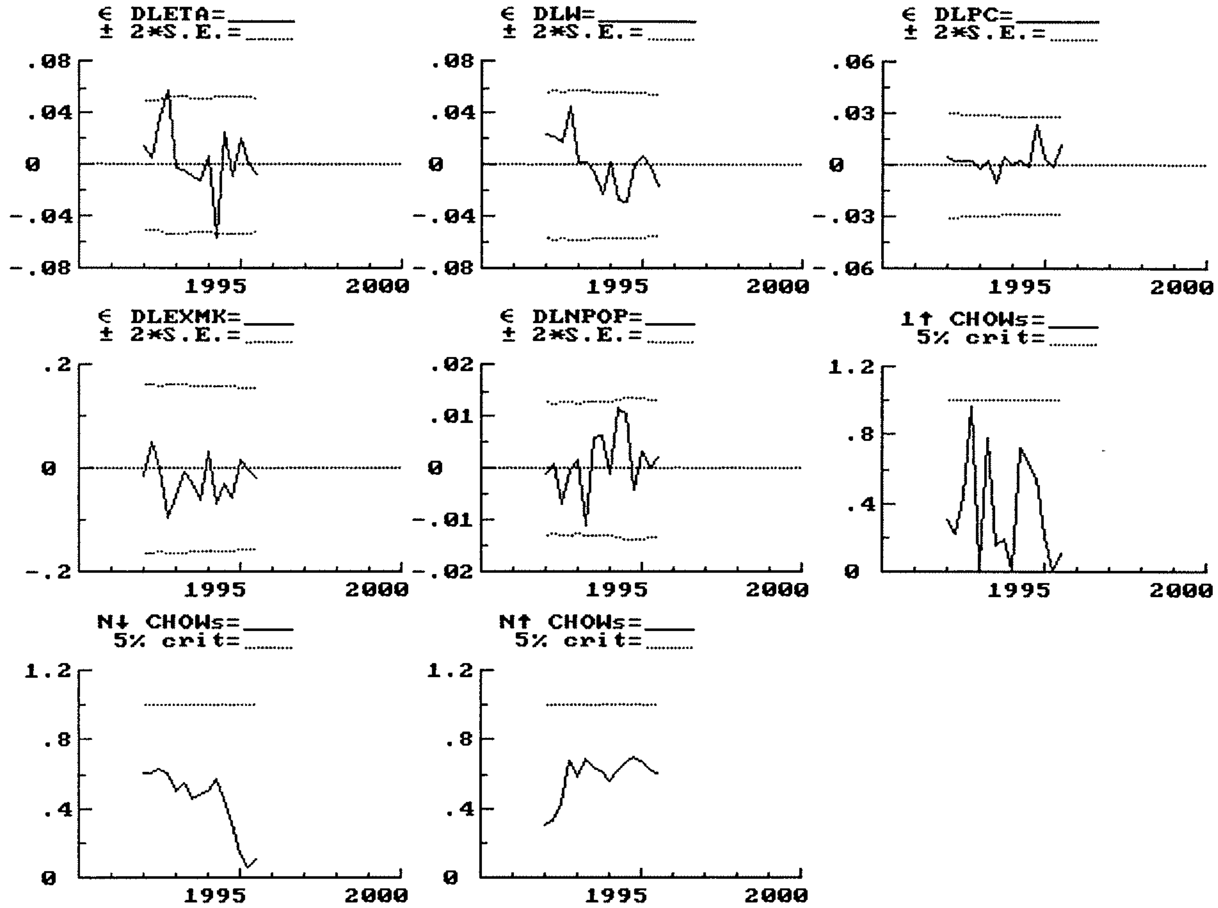
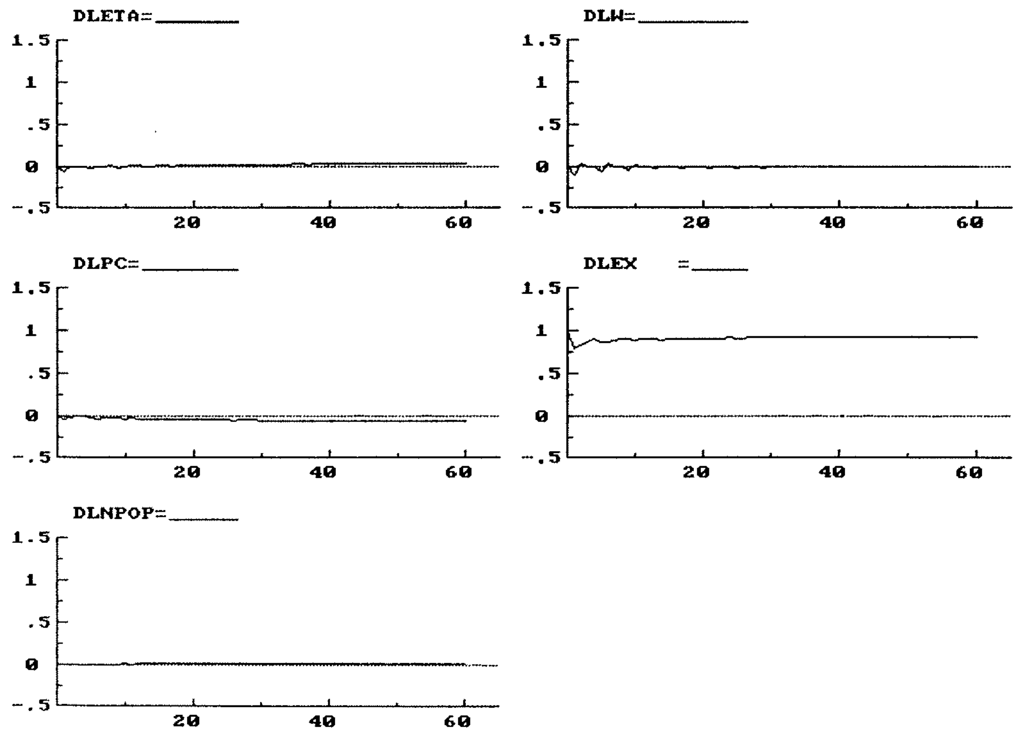


Figure 11.
Hungary: Cumulated impulse-response function of a DLEX unit shock in the simplified model



Annex 1. Poland: simplified model deterministic components

variable*	DLETA	DLW	DLPC	DLEX	DLNPOP
Constant	-3.2121	-	-5.0606	-3.1387	0.1828
Seasonal	-0.0543	-0.0445	0.0168	-0.0609	-0.0244
Seasonal(-1)	-0.0347	-0.0523	-	-0.1617	-0.0141
Seasonal(-2)	-0.0869	-	-	-	-0.0125
s1981p1	-0.1956	-0.0817	-	0.4447	0.0166
s1989p4	-1.7382	-0.7863	-0.2597	1.3340	0.1555
i1980p4	-0.1276	0.0967	-0.0553	-0.1378	0.0110
i1981p1	0.1004	-	-0.1584	-0.4165	0.0073
i1983p1	-	-0.2118	0.2526	-	0.0097
i1988p1	-	0.2435	0.3628	0.1908	0.0080
i1988p2	-	-0.0818	-0.2067	0.1177	-
i1988p4	-0.1397	0.1987	-	0.2633	0.0192
i1989p3	-0.2009	0.5190	0.2782	0.5606	-
i1989p4	0.8693	0.9671	0.6544	-0.8471	-
i1990p1	0.2439	0.4961	0.4745	-	-0.0528

* In each column we present the deterministic components of the corresponding equation in the model. With the letter *i* we refer to an impulse dummy (equal to 1 in the mentioned period, zero otherwise), with the letter *s* to a step dummy (equal to 1 from the mentioned period to the end of the sample, zero otherwise).

Annex 2. Hungary: simplified model deterministic components

variable*	DLETA	DLW	DLPC	DLEX	DLNPOP
Constant	0.9479	4.7344	-0.5924	7.5705	-
Seasonal	-0.1204	-0.0773	0.0477	-	-0.0071
Seasonal(-1)	-0.0475	-0.0440	0.0383	-	-0.0138
Seasonal(-2)	-0.0948	-0.0296	0.0301	0.0932	-
s1987p4	0.0147	0.1048	0.0395	0.0708	-0.0063
s1990p1	0.0367	0.0245	0.0485	-0.0451	-0.0119
i1981p1	-0.0628	-	-	-	-
i1988p4	-	0.1605	-	0.1333	-
i1991p1	-0.0497	0.0173	0.0244	-0.0737	-0.0292
i1991p2	-0.0826	0.0744	0.0527	0.0778	-0.0144
i1991p3	-0.1067	0.0303	0.0265	0.0510	-0.0115
i1993p1	-0.1006	-0.0827	0.0324	-	0.1169
i1994p1	-	0.0621	-0.0432	-	0.0529
i1995p2	-0.0547	-0.0590	0.0809	0.1187	0.0061

* The same as in the previous table.

Data appendix - Legenda of variables for Polish and Hungarian economies

LEX	national currency agaist dollar exchange rate, end of period
LY	industrial production index, 1985=1
LN[1]	employment, national economy, thousands
LETA	labour productivity, LY - LN
LPOP	total population, thousands
LNPOP	employment-population ratio, LN - LPOP
LPC	consumer price index, 1985=1
LW[2]	percapita monthly wages, national economy, net
LPM59	international manufactured goods index price (in dollars), 1980=100
LPRIV	private sector share in employment

[1] For Hungary, employment in industrial sector
[2] For Hungary, wages in industry from 1970 to 1993

5. References

Doornik, J.A. and Hendry, D.F. (1994a) *PcFiml* 8 : *An Interactive Program for Modelling Econometric Systems* . London: International Thomson Publishing.

Doornik, J.A. and Hendry, D.F. (1994b) *PcGive 8: An Interactive Econometric Modelling System.* London: International Thomson Publishing.

Golinelli, R. and Orsi, R. (1994) Price-wage dynamics in a transition economy: The case of Poland, *Economics of Planning,* 27, 293-313.

Hendry, D.F. (1995) *Dynamic Econometrics.* Oxford: Oxford University Press.

Johansen, S. (1988) Statistical analysis of cointegration vectors, *Journal of Economic Dynamics and Control,* 12, 231-254.

Johansen, S. (1995) *Likelihood Based Inference in Cointegrated Vector Autoregressive Models,* Oxford: Oxford University Press.

Mizon, G.E. (1995) Progressive Modelling of Macroeconomic Time Series: The LSE Methodology, in Hoover, K.D. (ed), *Macroeconometrics: Developments, Tensions and Prospects,* Dordrecht: Kluwer Academic Press, 107-169.

Osterwald-Lenum, M. (1992) A note with quantiles of the asymptotic distribution of the ML cointegration rank test statistics, *Oxford Bulletin of Economics and Statistics*, 54, 461-472.

Zivot, E. and Andrews, D.W.K. (1992) Further evidence on the great crash, the oil price shock and unit root hypothesis, *Journal of Business & Economic Statistics,* 3, 251-270.