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Modelling wage growth dynamics in Italy: 1960-1990

by

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Maggio 1995

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Abstract: In this work we analyse the dynamic properties of wage inflation, price inflation, unemployment and labour productivity using Italian annual data (1960-1990, source: Prometeia). Starting from a well-specified parsimonious VAR representation with Gaussian residuals, we test for the presence of long run comovements in the series. Tests of structural hypothesis on the cointegrating coefficients and tests on the adjustment coefficients allow us to isolate an error correction representation for wage inflation in which the rate of acceleration in wages depends on the contemporaneous rate of acceleration in prices and on the adjustment to long run disequilibrium as represented by a Phillips type relation. Recursive analysis suggests a reasonable constancy in the parameters of the model. Furthermore, there is evidence that wage inflation does not Granger-cause price inflation. The overall conclusion is that the wage inflation process is well modelled both in its long run and in its short run properties as an error correction mechanism conditioning upon price inflation, unemployment and for which labour productivity is excluded from the cointegrating vector, while analysis of channels other than wages is required to account for the inflation process.(J.E.L C32, J30)

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1. Introduction[1]

After the Phillips (1958) work about the existence of a stable negative relationship between wage inflation and unemployment rate a growing consensus arose during 1960s around it. During the 1970s, as it is well known, this view was challenged and the attention to the Phillips relation fell as a consequence of two related factors. The first one was simultaneous rise in inflation (in wages and prices) and unemployment experienced by main western economies, leading to the failure in the predictive performance of the econometric models based on that relation. The second factor was the theoretical arguments originally put forward by Friedman (1968) and Phelps (1967) and developed by the so called Rational Expectation approach according to a significant long run trade-off doesn't exist. During the 1980s those researchers who made use of the Phillips relation both as a medium term forecasting tool and as theoretical equation in the macroeconomic models relied on the existence of a short run trade-off, as in the case, for example, of the New Keynesian Economics.

The developments sketched above are deeply reflected in the empirical works on the Italian economy. For example, Gallaway and Koshal (1970) and Modigliani and Tarantelli (1973) find significant long run trade off in the 1960s, while Mohabbat and Arshanapalli (1985) analysing the period from 1970 to 1980 deny any role to the unemployment rate as an indicator of labour market unbalances in the wage equation. Onofri and Salituro (1985, 1987) studying the period from 1960 to 1984 give account for the collapse of a stable structural relation over the sample.

Following the new developments in the time series analysis concerning the unit root econometrics, a new attention has recently been payed to the modelling of dynamic

[1]This paper is an expanded and substantially revised version of D'Amato and Pistoresi (1994). The revision has benefited from helpful discussions with Mario Forni and Marco Lippi. The usual disclaimers apply. All the applied econometrics has been performed by using PC-FIML 8.0 by Jurgen Doornick and David Hendry.

relationships between wage, prices, unemployment and in some cases labour productivity in the spirit of the Phillips curve. Examples of studies that use cointegration analysis to investigate the nature of trends (stochastic versus deterministic) in these series and their long run comovements are: Alexander (1993), Favero (1988), Hall (1986, 1989), Juselius (1992), Mehra (1991, 1994), Nymoen (1992). A detailed analysis of the presence of a long and short run trade-off in the series of post-war U.S data on inflation and unemployment has recently been performed by King and Watson (1994) in the context of structural VAR analysis renewing the interest about the debate on the "Phillips relations and correlations" as shown by the thoughtful reply by Evans (1994) and McCallum (1994).

In the spirit of this revived debate, this paper analyses dynamic relationships between wage and price inflation, unemployment rate and productivity of labour using annual Italian data from 1960 to 1990 (source: Prometeia). Our effort will be mainly devoted at developing a model for an aggregate wage equation in the context of multivariate cointegration analysis. This approach (Johansen, 1991a) avoids the possible presence of simultaneity bias arising from the usual single equation estimation procedure and addresses the problem of non standard distributions in testing hypotheses about long run parameters. Given that some or all of the variables traditionally included in the wage equation exhibit features of non stationarity this appears to be the most appropriate context in which to recast issues of inference and testing of the relevance of the trade-off. Furthermore, as already noted, for example, by Prosperetti (1981) in his comment on Modigliani and Tarantelli (1977), the simultaneity bias arising from the possible presence, beyond the wage equation, of a second structural equation linking price inflation to wage inflation via unit costs, is particularly relevant in modelling long run equilibrium in this context. As it is well known, in the framework of cointegrated systems, there indeed exist circumstances in which weak exogeneity in some of the variables allows to consider partial system and even single equation estimation without affecting the efficiency of the estimates and these circumstances can be tested; (see, for example, Johansen, 1992a).

The modelling strategy pursued in this paper will follow these lines: after collecting evidence using univariate tests on the degree of integration of each of the series considered, the starting point of the modelling strategy is the estimation of a VAR tested for the lag length and general misspecification, including tests for stability. Once a satisfying parsimonious representation has been obtained, we test for cointegration under different restricting hypotheses about the deterministic component in the series as in Johansen-Juselius (1990) and in Johansen (1991a). It turns out that the presence of two cointegrating relationships cannot be rejected on the basis of the standard trace and λ-max tests. Furthermore, the analysis of recursive eigenvalues and of the cointegrating residuals supports the idea that the two long run relationships we find in the data are reasonably stable. We also use LR-tests for simultaneous restrictions on the long run coefficients to support the choice about the degree of integration of the variables (multivariate tests of integration) we made in the context of univariate analysis and to address the issue of identification of the two relationships. In particular, in order to achieve identification of the two long run relationships we will pursue different test strategies involving structural tests on the cointegrating coefficients and tests for weak exogeneity on the adjustment coefficients. There is also evidence that the Granger causality implicit in the result of cointegration runs from price inflation to wage inflation and not viceversa. These results imply that the mark-up view of the inflation seems not to be supported by Italian data, as found by Mehra (1994) for the US economy, and it suggests that the analysis of channels other than wages is required to account for the inflation process (as pursued by Juselius 1991 on Danish data and as claimed, for example, by Zenezini 1989 for the Italian economy). Furthermore, these tests allow us to estimate the short run error correction dynamics for the single equation representing the wage inflation process. Recursive analysis suggests reasonable empirical constancy over the sample.

The paper is organised as follows: section 2 presents the method and the model employed to inspect the deterministic and stochastic components in the comovements of the variables, section 3 presents the empirical evidence about long run and short run dynamics. Our conclusions are presented in section 4.

2. The method and the model

In this section, the basic concepts of multivariate cointegration analysis, estimation and the testing of long run relationships are briefly reviewed. This is done by estimating the cointegration space as in Johansen (1991a,b) and in Johansen and Juselius (1990) and then by testing more specific hypotheses of economic interest within this space.

Multivariate cointegration approach implies that the long run relationships are jointly estimated with the short-run dynamics (by applying the M-L procedure) thus using all the information available in the data. This is the main difference with respect to Engle and Granger (1987) "two step procedure" and implies more efficient estimates. In the two step procedure, the cointegration relation is estimated by a static OLS regression and the cointegration residuals are tested for stationarity (first step). If stationarity is accepted, residuals are included as Error Correction Terms in the final ECM model (second step). The main problem with this procedure concerns the estimated long-run parameters. If the variables are cointegrated, static OLS regression yields superconsistent estimates of the long-run responses of the model, but the same cannot be said of the estimated standard errors of the regression. This is because the distribution of the OLS estimator is generally non-standard (Phillips 1988), in particular the OLS estimator depends on the nuisance parameters and it is (asymptotically) biased in mean and median.

Statistical inference about the static regression of Engle and Granger is also misleading if the regressors fail to be weakly exogenous for the cointegrating vectors, a condition that is usually assumed rather than tested. Hence, it is not possible to test the significance of the estimated coefficients and perform long-run restrictions of economic interest on the static regression.

Large part of the studies on Italian data, including papers referred above, were performed by regressing wage inflation on price inflation and unemployment can be considered as examples of first step regression in the Engle and Granger and may be affected by the short-coming referred above. Examples of declared two step procedure applied to modelling wage equations in the spirit of Phillips relations are: Favero (1988), Hall (1986) on UK data and Mehra, (1994) on US data.

The problem of efficient estimates and correct inference is dealt with applying the multivariate procedure. Johansen's full system approach yields: 1) maximum likelihood estimates of the cointegrating vectors and of the weights with which the deviations from the long run equilibrium enter each equation of the system; and 2) likelihood ratio (LR) tests of cointegration. Also within this framework one easily gets a LR test of weak exogeneity and a LR test of linear restrictions on the cointegration relationships.

In this work, the cointegration analysis is carried out for the observational variable vector $z'_t = (\Delta w_t, \Delta p_t, u_t, \Delta q_t)$ where Δw is the first difference of logged nominal wage, Δp is the first difference of logged consumer price, u is the logged unemployment rate[2] and Δq is logged labour productivity growth; we return later to the precise definition of these variables. We insert the wage and price inflation separately instead of real wage dynamics, because we do not want to impose a sort of *a priori* "no-

[2]The logarithmic specification for unemployment has been chosen on the basis of the traditional explanation given by Lipsey (1960) and of the formal demonstration provided by Nickell (1988).

money-illusion" condition as it is currently preferred - see for example Hall(1986) and Hall(1989), Alexander (1993) and Darby and Wren-Lewis (1993). Money illusion will be tested on the data as well as other hypotheses[3]. The information set of I(1) variables to model is similar to Mehra (1994)[4]. The theoretical model underlying the selection of these variables is given by an "augmented" Phillips curve model assuming that prices are set as a mark-up over productivity-adjusted labour costs. This model is quite common in the empirical studies of wage and price dynamics so we refer to Mehra (1994) for details. Notice that with respect to this author, we do not use a priori productivity-adjusted labor costs $(\Delta w_t - \Delta q_t)$ letting this free to be determined as a tested restriction on the cointegrating relationships.

Assuming z_t is I(1), a dynamic modelling of the comovements between selected series starts from the following VAR representation:

$$A(L)z_t = \mu + \varepsilon_t \qquad [1]$$

where ε_t is a vector of white noises such as $E(\varepsilon_t) = 0$, $E(\varepsilon_t \varepsilon_t') = \Sigma_\varepsilon$ and A(L) is a matrix lag polynomial of order k with the normalisation A(0)=I, μ is a vector of constants that, as it will be seen in moment, will require a different treatment, according to the different hypothesis about the deterministic component in the VAR representation. Under the null of cointegration, model [1] can be reparameterised as a vector error-correction model (VECM):

$$\Delta z_t = \mu + \sum_{i=1}^{k-1} \Gamma_i \Delta z_{t-i} + \Pi z_{t-k} + \varepsilon_t \qquad [2]$$

where $\Gamma_i = -(I + A_1 + \ldots + A_i)$ and $\Pi = -(I - A_1 .. - A_k)$ are 4×4 matrices of unknown parameters, Δz_t and Δz_{t-i} are 4×1 vectors of I(0) variables, while the z_{t-k} is a

[3]For a discussion addressing the issue of expectations underlying the VAR specification refer to Juselius (1991) p.4. The intuition is that a reparameterised VAR in the ECM form requires that endogenous variables always adjust to their long run target. This implies that if, for example, in the wage equation, a price inflation coefficient less than one is found, agents find too costly fully adjusting wage inflation to price inflation. This may be due either to the impossibility of continuous by recontracting wages, or to the high costs for the economy of a full indexation that is taken into account by bargaining agents.

[4]For different arguments underlying the choice of these variables and references on the related literature see D'Amato and Pistoresi (1994).

4×1 vector of I(1) variables. The Π matrix, which conveys the information about the long-run behaviour of the system, is the main object of our investigation. In fact the hypothesis about the stationarity or non stationarity of the z_t vector, the presence of cointegration and the number of long run relationships are formulated as restrictions on the rank of this matrix.

Then we need to test the hypothesis of reduced rank of the Π matrix:

$$H_o\!: \Pi = \alpha\beta',$$

where the adjustment coefficient matrix α is $(4 \times r)$ and β is the $(r \times 4)$ matrix of cointegrating vectors, r is the rank of Π and determines the number of linearly independent stationary relations between the variables, i.e. the number of cointegrating vectors in the data. In other words, if the rank is zero there is no cointegration, while if Π is of full rank, then z_t is stationary (the null is rejected). Clearly if the rank is r<4, we can interpret the r relations: $\beta' z_t$ as the stationary relations among four nonstationary variables, i.e. as cointegrating relationships.

As it is well known, the asymptotic distribution of the rank test statistics crucially depends on the nature of the deterministic component in the variables included in model [2]. In our context two possibilities may arise: (i) the intercept μ either enters as an autonomous growth factor in the VECM representation of the system, (ii) the intercept only enters the ECM terms. In the former case, μ can be partitioned into $\mu = \alpha\beta_0 + \alpha_{\perp}\gamma$, where β_0 is an $(r \times 1)$ vector of intercepts in the cointegrating relationships, $\alpha_{\perp}$ is a $(4 \times (4 - r))$ matrix of full rank orthogonal to the columns of α, and γ is $((4 - r) \times 1)$ vector of linear trend slopes. Hence, to avoid singularity problem, model [2] has to be estimated without imposing restrictions on μ and it implies that there is a linear trend in some of the I(1) variables entering the z_t vector. Instead, when (ii) turns out to be the relevant case, μ has to be estimated restricted as follows:

$$\Delta z_t = \sum_{i=1}^{k-1} \Gamma_i \Delta z_{t-i} + \alpha(\beta' z_{t-k} + \beta_o) + \varepsilon_t \qquad [3]$$

Model [3] implies that there is no linear trend in the I(1) variables entering the z_t vector, in other words this model implies that: $\alpha_{\perp}\gamma = 0$. Since we deal with first differences in the variables, except for the unemployment rate, the hypothesis of no linear trend is likely to hold, however as we have doubts about the nature of the deterministic component in the unemployment rate and productivity growth, both hypothesis will be tested[5].

The hypothesis that there are no linear deterministic trends can be tested by applying the likelihood test procedure described in Johansen and Juselius (1990) and Johansen (1991a,b). In particular, it is possible to check the hypothesis of absence of linear trend versus the hypothesis of linear trends in some of the variables via the following LR test:

$$-2\ln Q\{H^*(r)/H(r)\} = -T\sum_{i=r+1}^{4}\{(1-\lambda_i^*)/(1-\lambda_i)\} \qquad [4]$$

where H* represents the model [3] and H represents the model [2] with the intercept unrestricted; λ indicate the eigenvalue(s) associated to the long run matrix (see Johansen 1991b, pag.10). This test statistics, under the null: $\gamma = 0$, is distributed as a χ^2 with $(4 - r)$ degrees of freedom. We will apply test [4] in order to select between model [2] and model [3], furthermore LR tests will also be performed to check some restrictions of economic interests on the selected cointegrating vectors, to test weak exogeneity and to perform multivariate integration analysis following Johansen and Juselius (1992a). In the latter case, to test that a given variable, z_{1t}, is stationary it has to be checked that the unit vector: (1, 0, 0, 0) is contained in the β-space. If the hypothesis is rejected z_{1t} can be treated as an I(1) variable in the case of the restricted model and as an I(1) variable plus a trend in the case of the unrestricted one.

[5] For a similar argument see Alexander (1993), p.89.

3. Empirical evidence

This section presents empirical results. The data used are annual and cover the period from 1960 to 1990. This data consist of a statistical reconstruction performed by Prometeia that is consistent with the major revision of the national accounts recently undertaken by the ISTAT[6]. This reconstruction is partly based on the reworking of the national accounts by Golinelli and Monterastelli (1990).

We analyse the interactions between rate of growth of nominal wage and prices, unemployment rate and rate of growth of productivity in a four-variable system consisting of: the difference of logged gross nominal wage expressed in per capita terms (Δw), the difference of logged consumer price deflator (Δp), basis 1980=1, the logged total unemployment rate (u), the difference of logged productivity of labour (Δq). We approximate the logarithms of the labour productivity by using the difference between the logarithm of real GDP and the logarithm of total employment. In the figures the following notation is used: Δw=Dlwr; Δp=Dlpc; u=lunr; Δq=Dlprod.

3.1 Univariate integration analysis

The necessary condition to perform the Johansen procedure, concerning model [2] and [3], is that all the variables in the VAR are I(1)[7]. Consequently, the first step in the Johansen procedure is to test the order of integration of the variables in the system.

Figure 1 plots the changes of the series and their correlograms. From the inspection of the plots it is possible to argue that nominal wage and price inflation are not stationary while the change in unemployment rate seems to be stationary.

[6]The major revisions were performed in 1966, 1969, 1975, 1987. See ISTAT (1990).
[7]Interpretation of the Johansen procedure is based on the premise that the variables are integrated of order one, I(1), if some (or all) the variables in the system are integrated of higher order than one , e.g. I(2), then a different procedure is required, see Johansen (1992b). Instead, if some variables are trend-stationary another model must be used: model with a restricted trend in the cointegrating vector (Johansen 1991b).

Furthermore, the stationarity of the first difference of labour productivity exhibits a negative trend. Thus, from a simple inspection of the plots, the nominal variables look like I(2), while the unemployment rate seems to follow an I(1) process with or without, it is not so clear, linear deterministic trend in the levels. The labour productivity could be I(1) with trend too, but the evidence of one (versus two) unit root is not so clear. The plots of the correlograms confirm the arguments put forward above; in particular the doubt about the stationarity of the first difference of labour productivity.

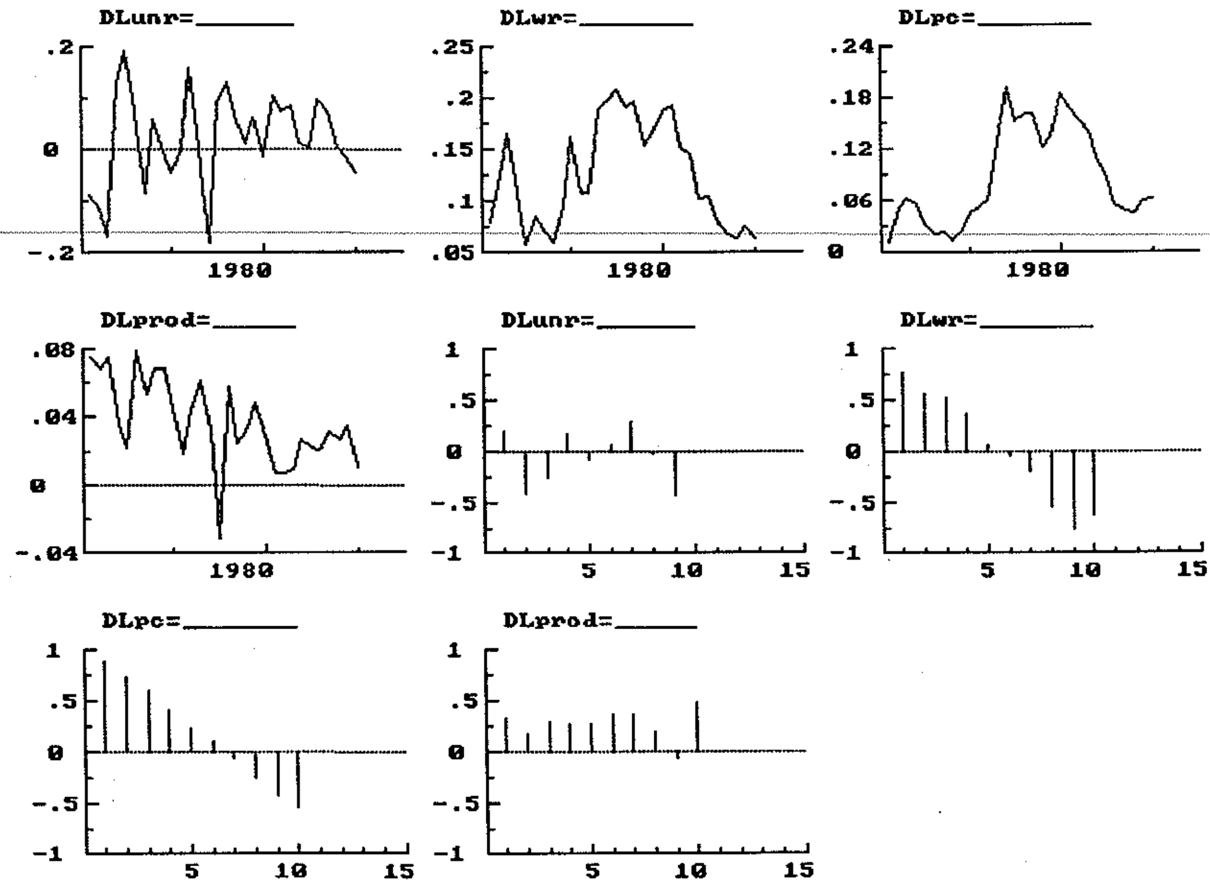

Figure 1 First differences of the time series data and their correlograms

We test the null hypothesis of the presence of unit root, by using the framework proposed by Dickey e Fuller (1979, 1980), represented either by a model of a random walk without drift:

$$y_t = \alpha y_{t-1} + \varepsilon_t \qquad [5]$$

or by a model of a random walk with drift:

$$y_t = \mu + \alpha y_{t-1} + \varepsilon_t, \qquad [6]$$

or by a model of a random walk with drift and trend:

$$y_t = \mu + \beta t + \alpha y_{t-1} + \varepsilon_t \qquad [7]$$

When necessary, the models with or without trend are augmented with a number of lags in order to prevent autocorrelated errors. The augmented Dickey-Fuller models are the following:

$$y_t = \alpha y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + \varepsilon_t \qquad [8]$$

$$y_t = \mu + \alpha y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + \varepsilon_t, \qquad [9]$$

$$y_t = \mu + \beta t + \alpha y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + \varepsilon_t \qquad [10]$$

Table 1 reports the results of integration tests for the basic set of variables. The number of augmentation used in the ADF tests (Augmented Dickey-Fuller tests) to avoid the autocorrelation of the residuals are also reported; in general, one lag is enough given the low dynamics of annual data. The testing procedure is valid if the residuals in estimated models are white noises. This is the reason why we accompany the statistics (DF and ADF) with a complete set of standard mis-specification tests (we reported only the probability-value of the tests): tests for autocorrelation, heteroscedasticity and normality. In particular, in the latter two columns of Table 1 are reported the adjusted

DF or ADF statistics obtained by correcting (when necessary) the heteroscedasticity in the regression results[8].

Table 1 Degree of integration: Dickey Fuller and Augmented Dickey Fuller tests

Variab.	Model	DF	ADF	lag	χ^2_{aut} p-value	F_{het} p-value	χ^2_{nor} p-value	White	Newey-West
w	no-trend		-1.28	1	0.50	0.64	0.99		
w	trend		-1.67	1	0.55	0.57	0.73		
p	no-trend		-1.37	1	0.15	0.36	0.023		
p	trend		-2.50	1	0.18	0.14	0.20		
u	no-drift	1.37			0.31	0.01	0.71	1.53	1.81
u	no-trend	-0.12			0.14	0.009	0.56	-0.13	-0.16
u	trend	-3.60			0.11	0.35	0.12		
q	no-trend		-3.37		0.55	0.71	0.001		
q	trend		-1.86		0.58	0.82	0.027		
Δw	no-drift	-0.63			0.74	0.50	0.80		
Δw	no-trend	-2.03			0.37	0.012	0.67	-2.70	-2.72
Δp	no-drift	-0.57			0.18	0.49	0.061		
Δp	no-trend	-1.67			0.12	0.076	0.06	-1.95	-3.57
Δu	no-drift	-4.13			0.31	0.31	0.78		
Δu	no-trend		-6.00	1	0.93	0.32	0.36		
Δq	no-drift	-2.27			0.062	0.31	0.055		
Δq	no-trend		-2.97	1	0.43	0.75	0.43		
$\Delta\Delta$w	no-drift	-5.13			0.38	0.79	0.76		
$\Delta\Delta$p	no-drift	-4.53			0.27	0.032	0.55	-3.23	-5.36

Note: Critical values (sample size=50): in model with trend ([7] e [10]): -3.80 (2.5%), -3.50 (5%), -3.18 (10%); in model without trend ([6] e [9]): -3.22 (2.5%), -2.93 (5%), -2.60 (10%), in model without drift ([5] e 8]): -2.25 (2.5%), -1.95 (5%), -1.61 (10%). χ^2_{aut}: Godfrey's 4-th order autocorrelation test (Godfrey 1978), F_{het}: heteroscedasticity test (Koenker 1981), χ^2_{nor}: Jarque-Bera test for normality (Bera and Jarque 1981), White and Newey-West: DF and ADF statistics adjusted (when necessary) the heteroscedasticity in the regression (Kim and Schmidt 1993, Newey and West 1987). For these references see Pesaran and Pesaran (1991).

[8]The first correction is performed using the White (1980) heteroscedasticity consistent covariance matrix estimator, the second correction is made using the Newey -West (1987) generalisation of White's estimator, fixing the size of Bartlett window. We choose the window size equal to 5. For these references see Pesaran and Pesaran (1991).

Performing the unit root tests on logged nominal wage (w) and logged consumer price (p) it is not possible to reject the hypothesis that these variables are I(2)[9]. In particular, there is a clear-cut result of unit root in the levels and in the first difference of these variables, while the second differences appear to be stationary.

Logged unemployment rate (u) yields a less clear result[10]. It is non stationary at 2.5% critical value and stationary at 5% critical value if the model with trend is considered (model [7]), while the unit root for the series in the model without trend (model [6]) and in the model without drift (model [5]) is out of doubt. Drawing inference on the nature of unemployment process on the basis of these statistical evidence is a bit entangled. The main point to discuss is the nature of the trend (if any) in the series, when we refer to the more general model [7] the test statistics -3.6 implies a trend-stationary representation at the 5% but a unit root representation with a quadratic trend[11] at 2.5%, which is clearly unlikely. On this ground we rely on representation [5] and [6] in which the unit root is accepted. The choice between [5] and [6], that is the choice between the presence or not of a linear trend in the MA representation of the unemployment rate, will be addressed in the context of the multivariate analysis.

Logged labour productivity (q) is clearly non stationary in the more appropriate model with trend, while the first differences of the variable appear to be non stationary for both models with no-trend and no-drift. Also in this case the analysis of the degree of integration of the variable will be further inspected in a multivariate framework.

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[9]For the same result of I(2) nominal variables see Hall (1986, 1989) and Alogouskufis- Smith (1991), Granger (1993).

[10]A review of similar studies has shown similar non clear result about the degree of integration of unemployment rate, suggesting that the degree of integration of a time series depends on the characteristics of the single economies and on the sample. For example, Blanchard e Quah (1989) consider the US unemployment rate as an I(0) variable; Nymoen (1989), testing integration on Norwegian quarterly data, concludes that the null of non stationary is rejected, although not very comfortably. Instead, Alexander (1993), on quarterly UK data and Darby and Lewis (1993) on annual UK data, accept the unit root in the series of unemployment rate. Evidence of rejections of the unit root with historical data from many countries can be found in Raj (1992). For the Italian case Fachin and Cicchetti (1994) achieve the same ambiguous result.

[11]See Banerjee et al. (1993, p. 100).

Furthermore, as it is well known, the occurrence of structural breaks in stationary time series may induce apparent unit root as shown by Perron (1989). Hence tests for unit root have low power when used for series with change in the slope or/and in the intercept of the time trend, while such series mimic series with unit root. For this analysis see Fachin and Cicchetti (1994), who analyse the structure of the main Italian macroeconomic series using the same data set utilised here. In particular, they perform bootstrap Dickey-Fuller tests, test for the null of trend stationary and endogenous break analysis for nominal wage, consumer price deflator, unemployment rate, real GDP and employment[12]. The non stationarity hypothesis appears to be the most likely in all cases but the unemployment rate for which no definitive conclusion can be drawn (however, the hypothesis of I(1) variables can not be rejected).

The univariate statistical analysis performed above suggests to select a specification for the system in inflation of wages and prices whose co-movements will be analysed together with the growth of labour productivity and unemployment rate. In the context of cointegration analysis applied to labour market, this specification using wage and price inflation rather than levels has been used, for example, by Favero (1988) and Mehra (1994), whereas Hall (1986, 1989) and Alexander (1993) focus on levels[13].

[12]The test for the null of trend stationary is proposed by Kwiatkowski *et al.*(1992), the analysis about the checking of "exogenous" structural breaks is that proposed by Perron (1989), while the analysis to check the "endogenous" structural breaks is proposed, for example, by Christiano (1992). For these references see Fachin and Cicchetti (1994).

[13]A possible rationale for the choice of levels (I(2)) of nominal variables could be subset cointegration: as stressed in Hylleberg and Mizon (1990 p. 116), when the dependent variables is I(1), "the regressor must include some I(1) variables or a combination of variables of higher order of integration which are cointegrated (to I(1)). If in such a case some of the regressor are I(0), these variables will affect the short run behaviour of the dependent variable only. Furthermore there can be no deterministic trend variables as a regressor in such cases". This argument gives some rationale for the specification adopted by Hall (1986), where the final selected long run equation has an I(2) wage level as a dependent variable that cointegrates with another I(2) variable, price levels, among regressors, all other regressors being I(1). But the way in which Hall finds (subset) cointegration between nominal wage rates and prices is not the usual one: he computes the difference in logs and apply the DF and ADF to the resulting real wage variable which ends to be I(1). Italian annual data accept the same test, but strongly reject the cointegration computed in the proper way, i.e. by testing subset cointegration on a multivariate VAR, between wages and prices inflation. For this reason we prefer not to follow Hall's procedure to specify the Phillips relation on the sample at hand.

3.2 *VAR specification*

As cointegration tests are valid under the hypothesis of Gaussian residuals of each equation of the system, we have to select the lag length of [1] that ensures residuals approximately white noise normal. The initial general system is a VAR with 3 lags in all the stochastic variables. Simplification tests on the initial system lag length suggested that 1 lag sufficed, so this reduction was implemented using likelihood ratio tests: the reduction for eliminating the lag length 3 and 2 is acceptable on overall F-tests (Table 3) and reduces the "costs" as measured by the Schwarz information criterion of model selection (Table 2) that has a minimum for VAR(1). These results induce to accept the most parsimonious specification (1-lag), holding in mind that cointegration results showed below are robust with respect to the lag length in the VAR.

Table 2 Information criteria

VAR	log-likelihood	Schwarz	Hannan-Quinn
VAR(1)	405.22968	-27.58	-28.25
VAR(2)	425.62831	-27.13	-28.35
VAR(3)	444.43071	-26.57	-28.33

Note: Schwarz criterion is defined as: $\log|\hat{\Omega}| + p\frac{\log(T)}{T}$

Hannan-Quinn criterion is defined as: $\log|\hat{\Omega}| + 2p\frac{\log(\log(T))}{T}$, where p is the number of coefficients, T is the sample period and $\log\hat{\Omega}$ is the log-likelihood of the model.

Table 3 Tests of model reduction

VAR(3)--------> VAR(2): F(16, 32)=1.23 (0.292)
VAR(3)--------> VAR(1): F(32, 42)=1.57 (0.082)
VAR(2)--------> VAR(1): F(16, 46)=1.85 (0.052)

Note: (.) p-value

Once chosen the lag length, we perform a series of diagnostic tests on the simplified VAR specification (Table 4). Diagnostic tests for autocorrelation, normality, heteroscedasticity and autoregressive conditional heteroscedasticity show that one lag is enough to whiten the residuals in each of the equations.

Table 4 Goodness of fit and evaluation

Statistic	**Single Equation Tests**				**Vector tests**
	Δw	Δp	u	Δq	
$F_{aut}(2,22)$ $F^{V}_{aut}(32,49)$	2.69 [0.09]	3.47 [0.05]	1.55 [0.23]	0.87 [0.42]	1.39 [0.14]
$F_{arch}(1,22)$ $F^{V}_{arch}(140,17)$	0.039 [0.84]	0.65 [0.42]	6.51 [0.018]	0.21 [0.64]	0.55 [0.06]
$F_{het}(8,14)$ $F^{V}_{het}(80,46)$	0.90 [0.53]	1.03 [0.45]	5.52 [0.002]	1.65 [0.18]	1.1 [0.36]
$\chi^2_{nor}(2)$ $\chi^{2\ V}_{nor}(8)$	2.48 [0.28]	5.23 [0.07]	3.14 [0.20]	3.74 [0.15]	14.88 [0.06]

Note: no serial correlation (F_{aut} against 2th-order autoregression, Godfrey 1988); no autoregressive conditional heteroscedasticity (F_{arch} against 2th-order autoregression: see Engle 1982); no heteroscedasticity (F_{het}: see White 1980); chi-square test for normality (χ^2_{nor}: see Doornick and Hansen 1993); analogous system (vector) tests are denoted by ^{V} (see Doornick and Hansen 1993). p-value are given in brackets [.]. For these references see Hendry and Doornick (1993).

For a first insight in the relationships between variables in the system, we consider the residual correlations which show that there is a large positive correlation between wage inflation and price inflation, there are negative correlations between unemployment rate, price inflation, wage inflation and productivity growth which merit modelling (Table 5).

Table 5 Residual Correlations

	Δw	Δp	u	Δq
Δw	1			
Δp	0.7664	1		
u	-0.4672	-0.3822	1	
Δq	0.4377	0.4506	-0.4551	1

The next step is to check the congruency of the autoregressive model. This involves testing for cointegration and for constancy. Since cointegration is only well defined if the long run relations are constant, we use recursive estimation as an indication of constancy for the I(1) system (Hendry and Doornick 1993, p.18). In Figure 2 are presented 14 plots: 1-step residuals, 1-step and N decreasing Chows (the hypothesis of stability tested for every possible sample split) for each of the four equations plus 1-step and N decreasing Chows for the system as a whole (indicated as Chows in the plots). Most of the plots suggest reasonably constancy for all the four equations. The individual equation break-point Chow F-tests are scaled by their 5% significance levels and none of their values exceed unity excepting for the productivity equation, however almost no system break-point test are significant.

Figure 3 reports the 1-step forecast statistics for 1985-1990. The constancy is easily accepted: every forecast lying inside the individual 95% confidence interval, shown by the vertical error bars of (±2S.Errors), based on the 1-step ahead forecast error variances, and the system constancy test is not significant at the 5% level at any horizon.

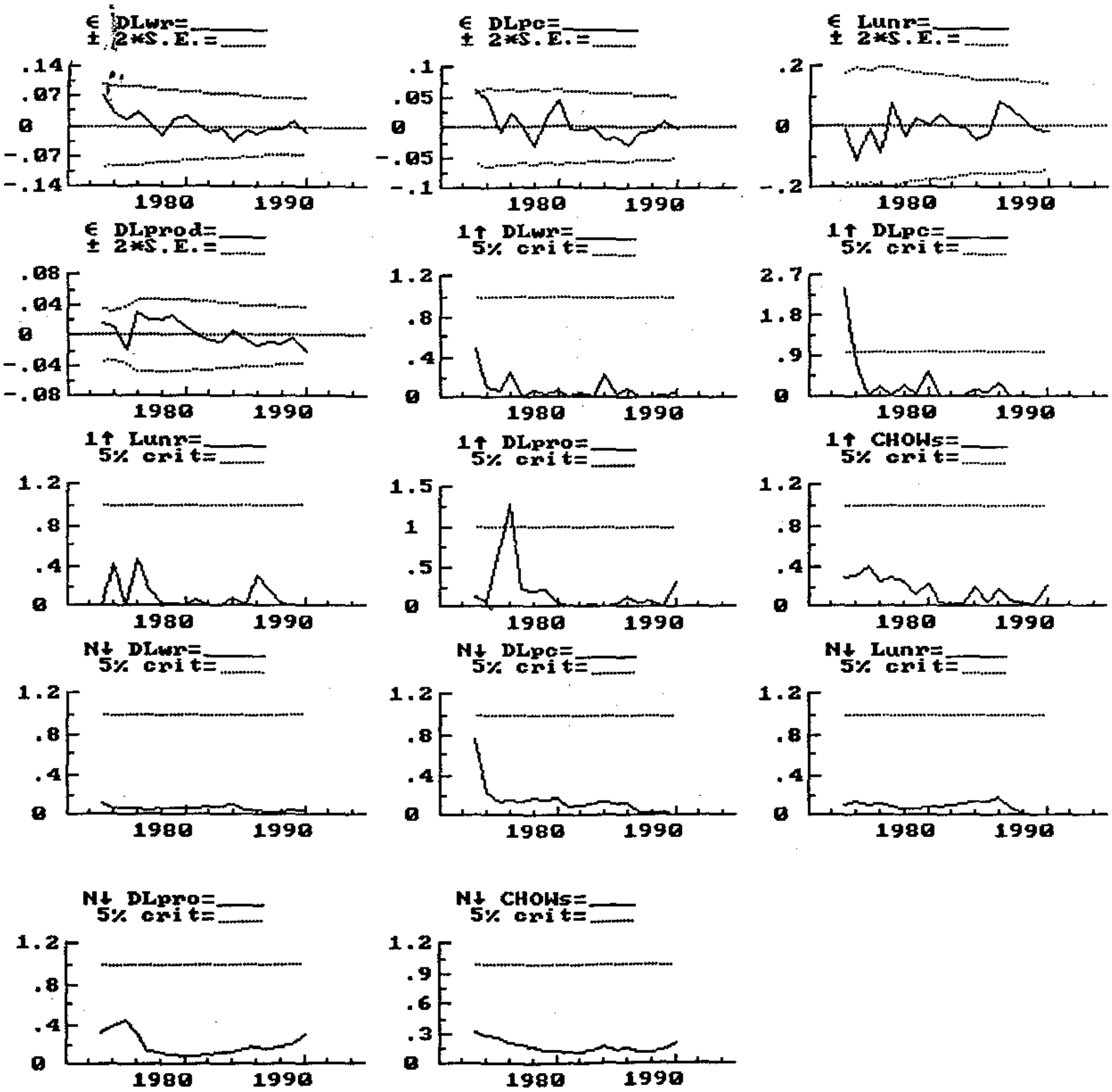

Figure 2 System recursive evaluation statistics

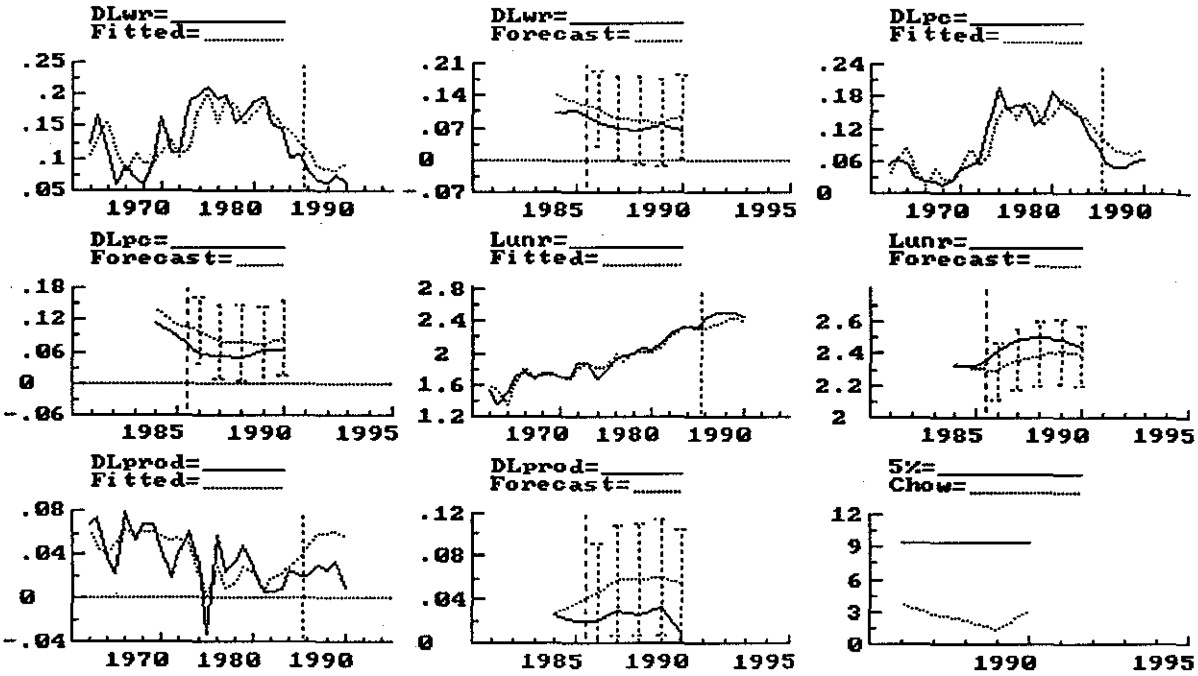

Figure 3 Forecast analysis: constancy statistics

3.3 Cointegration analysis

Given that the statistical analysis of parameter constancy in the VAR specification does not show significant structural breaks in the system, we are now able to test for cointegration. This is performed both in model [2] with unrestricted intercept and in model [3] with restricted intercept. The standard tests of the hypothesis of reduced rank of Π matrix enable us to accept r=2 in both cases. In fact, both the maximal eigenvalue and trace tests are consistent, at 95% significance level, with the presence of two cointegrating vectors for both the models (Table 6). Instead, the tests adjusted for the degrees of freedom, proposed by Reimers (1992), are consistent with just one cointegrating vector for both models. The estimated cointegrating vectors and the adjustment coefficients to the long run (dis)-equilibria are presented in Table 7, where unstarred coefficients indicate estimates for model [2] and starred coefficients indicate estimates for model [3].

Before addressing the issue of the selection of the number of the cointegrating vectors, we have to test model [2] against model [3] that is to test for the nature of the deterministic components in the system. To this aim, we consider testing H(2) in H*(2) via the LR test as in [4]. Using the trace values in Table 6 the test is given by:

$$-2\ln Q\{H^*(2)|H(2)\} = -T\sum_{i=3}^{4}\ln\{(1-\lambda_i^*)/(1-\lambda_i)\} = 11.26\text{-}5.83\text{=}5.43,$$

the asymptotic distribution of the test for the hypothesis $\gamma = 0$ is $\chi^2(2)$ and thus it unables us to accept the null, so that the restricted model [3] about the absence of deterministic trend is supported by data and consistent with the idea that the variables in difference and unemployment rate (levels) are not characterised by a linear deterministic trend[14].

Table 6 Johansen Maximun Likelihood Procedure

H_0	H_1	Statistic	Adj-Stat	95% CV	Statistic*	Adj-Stat*	95% CV*
Cointegration LR test based on maximal eigenvalue of the stochastic matrix							
$r=0$	$r=1$	51.1	44.05	27.1	51.2	44.15	28.1
$r\leq 1$	$r=2$	24.93	21.49	21	24.93	21.49	22
$r\leq 2$	$r=3$	3.87	3.33	14.1	8.26	7.12	15.7
$r\leq 3$	$r=4$	1.96	1.69	3.8	2.99	2.58	9.2
Cointegration LR test based on trace of the stochastic matrix							
$r=0$	$r\geq 1$	81.86	70.57	47.2	87.39	75.33	53.1
$r\leq 1$	$r\geq 2$	30.76	26.52	29.7	36.19	31.2	34.9
$r\leq 2$	$r\geq 3$	5.83	5.03	15.4	11.26	9.7	20
$r\leq 3$	$r=4$	1.96	1.69	3.8	2.99	2.58	9.2

eigenvalues: $(\lambda_1 = 0.8282; \lambda_2 = 0.5766; \lambda_3 = 0.1250; \lambda_4 = 0.06551)$
$(\lambda_1^* = 0.8288; \lambda_2^* = 0.5766; \lambda_3^* = 0.2480; \lambda_4^* = 0.0981)$

Note: CV= Critical Values are from Osterwald-Lenum (1992), * indicates the eigenvalues, statistics and critical values associated to the restricted model [3]; Adj-Stat = the test statistics are adjusted for degrees of freedom following Reimers (1992).

[14]It is worth to remember that even if only one variable in the system is characterised by a linear trend the right specification for the model is the one given in [2].

Table 7 Estimated values of cointegrating vectors and related adjustment coefficient

Variables	β_1	β_2	α_1	α_2	β_1^*	β_2^*	α_1^*	α_2^*
Δw	- 1	- 1	-0.082	0.74	-1	-1	-0.082	0.74
Δp	0.29	0.90	- 0.13	- 0.33	0.29	0.90	-0.13	-0.33
u	- 0.14	- 0.049	0.75	0.22	-0.15	-0.050	0.78	0.21
Δq	- 3.49	0.33	0.30	-0.25	-3.5	0.32	0.30	-0.25
$\mu = \alpha\beta_0$*					0.52	0.13		

Note: the estimated cointegrating vectors: β_i, β_i^* and the relative adjustment coefficient vectors: α_i, α_i^* presented above are normalised with respect to the nominal wage inflation, * indicates the results related to the case of the restricted model [3].

Figure 4 contains the plots of the estimated disequilibria (or error corrections) for the two cointegrating vectors of the selected model, the remaining two non-stationary components together with their eigenvalues estimated recursively.

The first and second eigenvalues are constant and they are also different to zero supporting the result of two cointegrating relations, their constancy suggests that the cointegrating relationships are stable over the sample. Furthermore, it is worth noting that even if there may be possible evidence of a structural break in the drift in some of the (univariate) series enclosed in the system, the cointegrated combinations are not affected by them, possibly because all the series included share the breaks. The remaining two components (vector 3 and 4 in the plots) are distinctly non-stationary with the associated eigenvalues are close to zero. Nevertheless, the difference in the result between adjusted and unadjusted cointegration tests may suggest to further inspect this issue. We chose to work with Sp(β)=2, both relying on grafical evidence about recursive eigenvalues, as suggested in Hendry and Doornick (1993), and on theoretical arguments, given that we started the analysis having in mind a simple model based on a Phillips- type wage equation and a price mark-up equation[15].

[15] The number of the cointegrating vectors in the system is, roughly speaking, correlated to the amount of stability in the economic system they represent and, from the point of view of economic theory, it is preferable to have as many cointegrating vectors as possible, i.e. setting as many constraint as possible to the steady state. For example, when the long run matrix is full rank there exist as many long run relations

Having identified the basis of the cointegration space is only the starting point for the analysis. In order to make economic sense of the estimated cointegrating vectors we need to adress the issue of identification. This is possibly the most delicate issue in cointegration analysis because the order of testing is not clear.

As stressed by Johansen and Juselius (1992b) "Given an identified long run structure, identification of the short run structure can often be achieved by means of identifying restrictions on the weight parameters, or alternatively given identifying restrictions on the weight coefficients α, identification of the long run structure can be achieve. This suggests that the identification of the long run structure is closely related to the identification of the short run structure and viceversa through the weight coefficients α". One possible approach is to analyse the decomposition of the reduced rank matrix ($\Pi=\alpha\beta'$) starting from the primary problem of identification of the structural hypothesis lying in the cointegrating space. Though this approach is more appealing, it may incurr in problems about the estimation of the α-matrix, as we will turn out in our case (next section). When this is the case it is possible to attempt to achieve identification by jointly testing restrictions on α and β coefficients (for an application of this view see, for example, Hendry 1995, p.599).

as variables and there exist a unique steady state in the system, that is a unique point in the space spanned by the variables included in the system towards which the system itself converges in the long run. When, as in our case, there exist two cointegrating vectors and four variables, the steady state is represented by a plane rather than a point. In this case, the equilibrium relations constrain, in the long run, the four variable to lie in a plane in the R^4 space spanned by them, this plane is defined by the intersection of the two cointegrating relationships. The variance of the system is finite about the plane and infinite within the plane. This means that there exist two common trends, that is, two directions in which the variance of the system is unbounded. For an interpretation of the cointegrating vectors along the lines presented above see: Dickey, Jansen and Thornton (1994, p. 22-23).

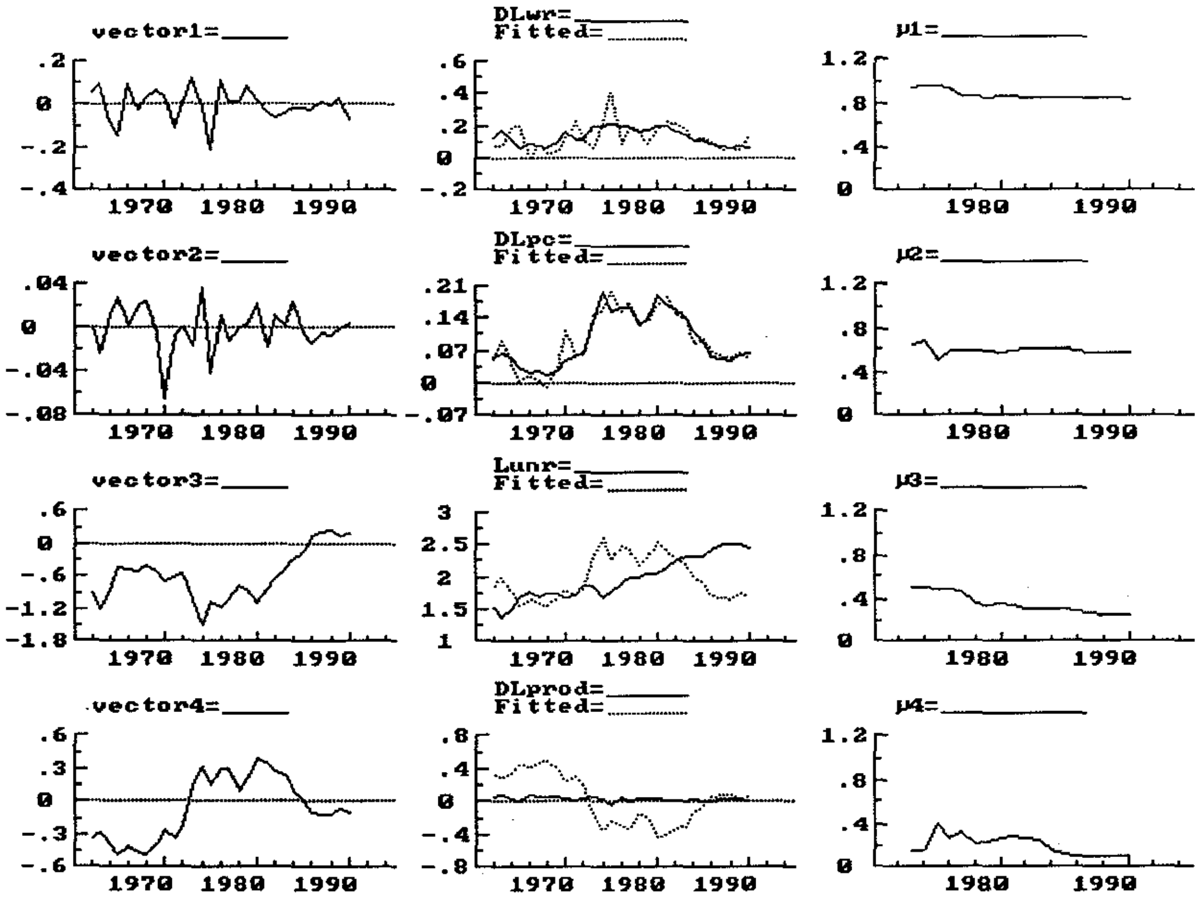

Figure 4 Cointegrating vectors and recursive eigenvalues

Before addressing the issue of structural hypotheses let us make confident about the well behavedness of the likelihood surface for the relevant estimated parameters. Figure 5 shows the 1-dimensional of the full sample likelihood surface for each of the 18 parameters of the system (Table 7), under the null of cointegrating rank r=2, 4×2 α −matrix, plus 2×4 β −matrix, plus 2 restricted constant parameters defined the number of parameters to be estimated. These plots suggest that the 1-dimensional likelihood surface is uni-modal in the neighbourhood of the optimum which is an index of well behaved cointegration estimates.

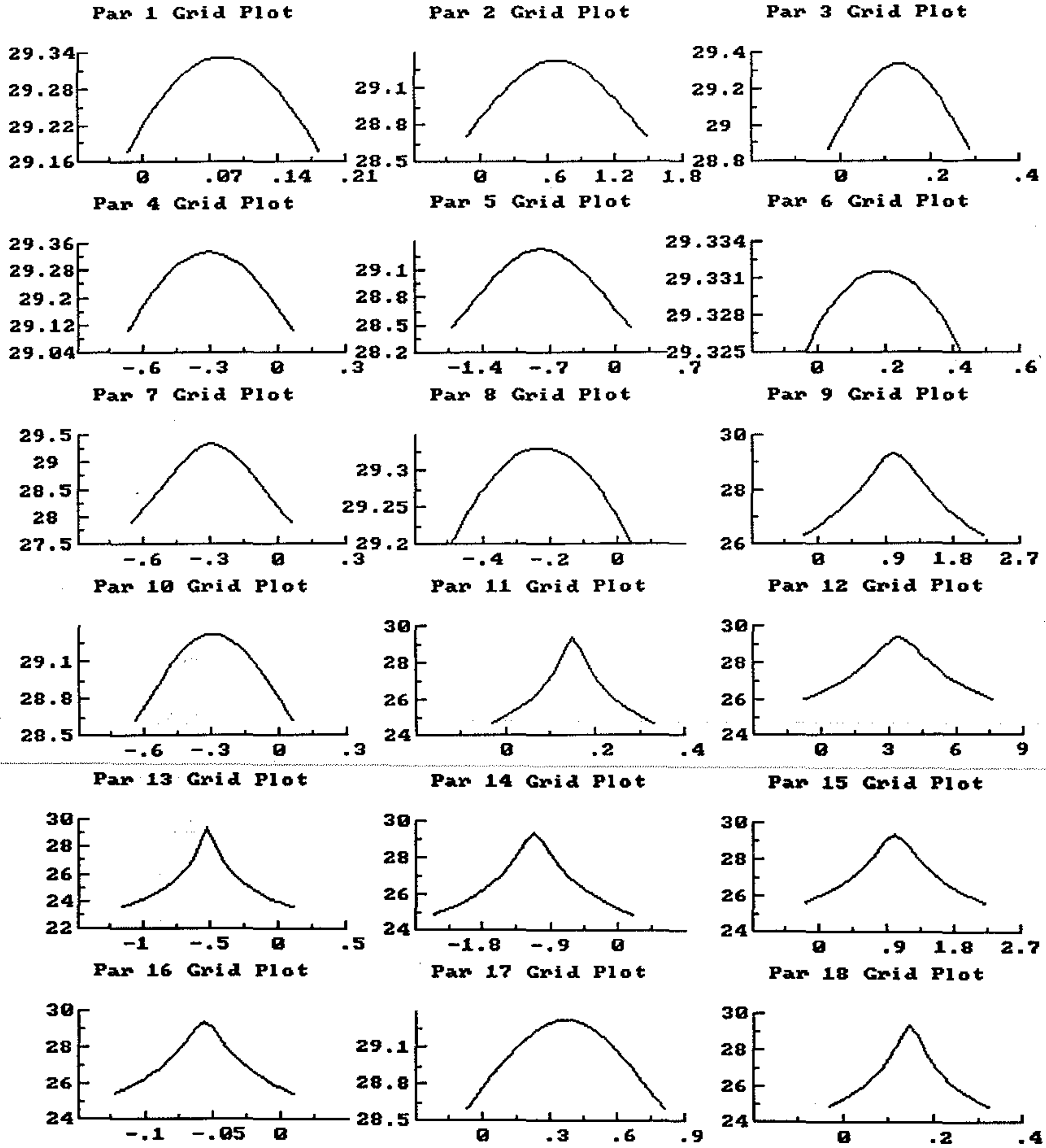

Figure 5 1-dimensional projections of the likelihood surfaces

3.4 Testing linear restriction on the cointegrating vectors and on the adjustment coefficients

The strategy to deal with this problem is the following: we start from the more general rank equal 2 model and we try to identify cointegrating vector as suggested by the theory, that is a wage equation and a price mark up equation. As we reported in the previous section, the test of structural hypothesis is a crucial part of the analysis, in particular when there is more than one cointegrating vector. The structural hypothesis are formulated as tests about the cointegration space asking whether some hypothetical relations can be assumed to lie in the stationary part of the whole space spanned by the variables in the system.

We impose linear restrictions on the estimated coefficients of the cointegrating vectors presented in Table 7. In order to perform identification of the basic relationships between the variables included in the system we rely on a very simple model of wage and price dynamics based on a mark-up rule for price inflation and a wage equation modelled as an "augmented" Phillips Curve, under the usual hypothesis that expected price inflation equals actual price inflation as for example in Mehra (1994). 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. Notice that all of the tests reported below are performed in Sp(β)=2, that is in the cointegrating space as a whole. In particular, the second cointegrating vector appears to be a candidate for a price mark-up relation of the sort:

$$\dot{p} = \dot{w} - \dot{q} + \mu, \qquad [11]$$

holding under the hypothesis of a long run constant mark-up[16]. By testing this restriction on the data we cannot reject the null:

[16]For a similar hypothesis see Mehra (1994) and Juselius (1991).

$$H_0 1 \in 2Sp(\beta): \begin{cases} \beta_2^{\ *} = (-1,1,0,1,\mu_2) \\ \beta_1^{\ *} = (a,b,c,0,\mu_1) \end{cases}, \text{LR(2)=5.25 (pv=0.0722)}$$

with the estimated coefficients equal to:

$$\hat{\mu}_2 = -0.0025, \hat{\mu}_1 = -0.18, \hat{a} = 1.512, \hat{b} = -1.27, \hat{c} = 0.054.$$

Normalising with respect to wage inflation the other cointegrating vector turns out to be:

$$\beta_1^* = (-1, 0.84, -0.036, 0, 0.12)$$

this may be interpreted as a wage equation in the spirit of an "augmented" Phillips Curve:

$$\Delta w = 0.84\Delta p - 0.036u + 0.12 \qquad [12]$$

To further characterise the just identified long run relationships we implement now some other LR tests representing the following hypothesis: the coefficient of unemployment rate in zero in the wage equation under different hypothesis concerning the other coefficients of the $\beta_1^{\ *}$ cointegrating vector ($H_0 2$, $H_0 3$), the trade-off between real wage and unemployment rate ($H_0 4$), $\hat{\mu}_2 = -0.0025$ in the mark up equation is restricted to zero ($H_0 5$), the multivariate integration analysis for u and Δq (respectively $H_0 6, H_0 7$). The hypothesis $H_0 2$, $H_0 3$, $H_0 4$ are largely rejected, while $H_0 5$ is accepted leaving [12] unchanged and the mark up equation [11] slightly modified in a long run relation without constant term. Hence, focusing on $H_0 5$, data seem to support the basic representation of comovements in wage and price inflation in terms of a simple Phillips curve - price mark-up system. We will reconsider this interpretation in a moment, by testing if α-coefficients are compatible with the identification achieved by imposing linear restrictions on β-coefficients. However, independently from the selected identification restrictions, the hypotheses in Table 8 suggest that a trade-off between nominal variables and unemployment rate is data consistent. Altogether, the empirical analysis gives some evidence against the assumption of a vertical long-run Phillips curve in terms of nominal wage inflation as for example in Juselius (1991) on Danish data.

As the whole analysis here crucially depends on the degree of integration of the variables included in the system, in particular of the unemployment rate and productivity, before moving to testing α-coefficients we check their degree of integration in the multivariate framework, in order to support the previous (non clear) interpretation of the univariate analysis outcomes. We reject the hypotheses of I(0) process for unemployment rate and growth of labour productivity as reported in $H_0 6, H_0 7$ in Table 8.

Table 8 Restrictions on the cointegrating vectors

$H_0 2 \in 2Sp(\beta): \begin{cases} \beta_2^* = (-1,1,0,1,\mu_2) \\ \beta_1^* = (*,*,0,0,\mu_1) \end{cases}$ LR(3)=22.58 (p-value=0.00),

$H_0 3 \in 2Sp(\beta): \begin{cases} \beta_2^* = (-1,1,0,-1,\mu_2) \\ \beta_1^* = (-1,*,0,*,\mu_1) \end{cases}$ LR(3)=22.58 (p-value=0.00),

$H_0 4 \in 2Sp(\beta): \begin{cases} \beta_2^* = (-1,1,0,1,\mu_2) \\ \beta_1^* = (-1,1,*,0,\mu_1) \end{cases}$ LR(3)=22.91 (p-value=0.00),

$H_0 5 \in 2Sp(\beta): \begin{cases} \beta_2^* = (-1,1,0,1,0) \\ \beta_1^* = (a,b,c,0,\mu_1) \end{cases}$ LR(3)=5.46 (p-value=0.14), with $\hat{a} = -1, \hat{b} = 0.84,$ $\hat{c} = -0.036, \hat{\mu}_1 = 0.12$

$H_0 6 \in 2Sp(\beta)$: $\beta^* = (0,0,1,0,\mu)$ LR(3)=12.6 (p-value=0.0018),

$H_0 7 \in 2Sp(\beta)$: $\beta^* = (0,0,0,1,\mu)$ LR(3)=19.53(p-value=0.0001)

3.5 Identification through joint linear restrictions on the adjustment coefficients and long run coefficients

In this section we further inspect the identifying restrictions we found in the previous section by testing weak exogeneity. In particular we test if the identified cointegration vector enters the relevant ECM representation and regard this as an indication of robustness of the identification achieved by testing the cointegrating coefficients.

Table 9 summarises some of the tests we performed on the α-coefficients. We can not reject $H_0 8$, this means that, given the identification of the two cointegrating vectors in $H_0 5$ in Table 8, the ECM of wage does not adjust to the wage inflation disequilibrium and that the ECM of prices does not adjust to the price inflation disequilibrium, which is clearly inconsistent. Furthermore, we test $H_0 9$ which represents the hypothesis fully consistent with the previous identification of the cointegrating vectors, LR tests lead to their rejection.

The overall evidence suggests to reformulate the identification strategy, now we start from the restrictions on α-coefficients to acquire some information on the β-coefficients. The acceptance of $H_0 10$ supports evidence that the only variable that is not weakly exogenous with respect to the first cointegrating vector is wage inflation and that price inflation is weakly exogenous within the system. Furthermore, in the context of a VAR(1), weak exogeneity of prices implies that wages do not Granger cause prices[17]. The acceptance of $H_0 11$ show that the cointegrating vector that we identified

[17]We also checked the robustness of these results with respect to the addition of one more lag length in the initial VAR representation. In particular, considering the reparameterisation of a VAR(2) in ECM form (under the null of rank equal 2 of the long run matrix), H10 and H11 cannot be rejected confirming the result of weak exogeneity and the identification proposed in the text. Following Mehra (1994, pp.154-156) the analysis in the context of an error correction representation which requires testing zero coefficients on the lags and on the adjustment coefficients confirms that wage and price inflation are related with Granger causality running from prices to wages and not viceversa. Also following Mehra (1994 pp. 156-159) we replicated the analysis in the standard VAR representation using FIML estimation

as a wage equation only enters the ECM representation for the wage inflation and also that it is the only cointegrating relation to enter this ECM.

In particular, hypothesis $H_0 11$ implies that:

$\hat{\alpha}_1^* = (-0.40,0,0,0)$, $\hat{\alpha}_2^* = (0,0,-0.71,-0.44)$, $\hat{\beta}_1^*(-1,0.82,-0.059,0,0.17)$ and $\hat{\beta}_2^* = (-11.55,-1,-2.36,-64,44,8.66)$. $\hat{\beta}_2^*$ can be normalised with respect to unemployment or productivity, in these cases becomes: $\hat{\beta}_{2u}^* = (-4.89,-0.42,-1,-27.30,3.66)$ or $\hat{\beta}_{2q}^* = (-0.18,-0.015,-0.036,-1,0.13)$. The long run wage equation, under the accepted hypothesis on α-coefficients, and confirming constancy with respect to different hypotheses, it turns out to be nearly identical to the one presented in the previous section. The estimated long-run wage equation is the following:

$$\Delta w_t = 0.82\Delta p_t - 0.059u_t + 0.17.$$

Hence, by joint restrictions on the components of the reduced rank long run matrix ($\Pi=\alpha\beta'$), we are able to identify a long run wage equation. The second cointegrating vector can either be interpreted as a productivity equation or as an unemployment equation[18]. However, $H_0 11$ enables us to specify an error correction representation for the short run dynamics wage inflation, which will be modelled in the next section.

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procedure. F tests of zero restrictions on lagged wage inflation and lagged price inflation respectively in price and wage equations of the system can not be rejected, supporting the existence of no-Granger causality from wages to prices. All the results are available on request.

[18]In order to get information about the second cointegrating vector we estimated recursively the error correction representation for $\Delta\Delta q$ and Δu. The former exhibits a structural break detected by recursive Chow tests. The latter does not exhibits any structural break. This give indications that the second cointegrating vector $\hat{\beta}_2^*$ may represent an unemployment equation.

Table 9 Testing weak exogeneity

$$H_0 8 \begin{cases} \alpha_1^*(0,*,*,*), & \beta_1^*(-1,*,*,0,*) \\ \alpha_2^*(*,0,*,*), & \beta_2^*(1,-1,0,-1,*) \end{cases} \quad \text{LR(4)= 8.05 (p-value=0.089)}$$

$$H_0 9 \begin{cases} \alpha_1^*(*,0,0,0), & \beta_1^*(-1,*,*,0,*) \\ \alpha_2^*(0,*,0,0), & \beta_2^*(1,-1,0,*,*) \end{cases} \quad \text{LR(4)=45.72 (p-value=0)}$$

$$H_0 10 \begin{cases} \alpha_1^*(*,0,0,0), & \beta_1^*(*,*,*,*,*) \\ \alpha_2^*(0,0,*,*), & \beta_2^*(*,*,*,*,*) \end{cases} \quad \text{LR(3)=5.44(p-value=0.14)}$$

$$H_0 11 \begin{cases} \alpha_1^*(*,0,0,0), & \beta_1^*(*,*,0,*,*) \\ \alpha_2^*(0,0,*,*), & \beta_2^*(*,*,*,*,*) \end{cases} \quad \text{LR(3)=6.58 (p-value=0.09)}$$

3.6 Short run dynamics

Results on weak exogeneity and the acceptance of the hypothesis $H_0 11$ enable us to estimate a single ECM representation for the wage equation conditioning upon price inflation, unemployment and for which labour productivity is excluded from the cointegrating vector. Hence, we concentrate on the first cointegrating vector as the unique error correction mechanism to which wage dynamics adjust in the long run, $H_0 11$ allows to consider price inflation, unemployment rate and productivity growth as weakly exogenous with respect to this short run equation and hence an error correction model in the same spirit as in Phillips original work turns out to be the result of a progressive marginalisation strategy as, for example, in Hendry and Doornick (1993).

In Table 10, we report the results about the short run dynamics of the acceleration of the wage inflation, this turns out to depend on the acceleration in the price inflation and on the disequilibrium represented by a simple long run Phillips curve as specified above. The size of the estimated coefficients imply large and immediate responses of the acceleration of the wage inflation to contemporaneous acceleration of price inflation and to disequilibrium via the error correction term. As the error correction term comes from the estimation of model [3], the intercept in the short run specification turns out to be not significantly different to zero. All the short run parameters have the expected sign; the adjustment coefficient to the long run disequilibrium and the coefficient of the acceleration of price inflation are significantly different to zero. A Wald test for the zero restrictions on the coefficients of unemployment changes and productivity growth changes suggests that these variables are not significant to explain the acceleration of wage inflation.

Concerning the statistical attributes of the ECM representation of wages, the various diagnostic checks, also reported in Table 10, indicate the absence of misspecification of any sort: there is no evidence of serial correlation, heteroscedasticity and non normality of the OLS residuals and the Reset test for functional form also confirms the good approximation of the model to the data generating process.

Continuing in a progressive research strategy, we inspect the constancy of the ECM by recursive analysis. Plots from one to five in Figure 6 record the recursive estimates of the parameters of the model together with a confidence region based upon plus-or-minus twice its estimated standard error at each sample size. As the sample size increases, the estimated coefficients changes relative to its estimated uncertainty, with the final estimate lying inside the initial confidence interval. Such evidence indicates reasonably constancy of the estimated parameters. Figure 6 also reports the one-step residuals and the corresponding calculated equation standard errors and the 1-step and N decreasing Chow tests. These tests also indicate a good measure of constancy for the ECM.

Table 10 Short run dynamics

$$\Delta\Delta w_t = 0.002 + 0.90\Delta\Delta p_t - 0.072\Delta u_t + 0.009\Delta\Delta q_t - 0.98ecm_{t-1}$$
$$(0.65) \quad (4.81) \quad (-1.31) \quad (0.068) \quad (-4.45)$$

$R^2 = 0.67$, DW=2.05,

$F_{WT}(2,24)$=0.216 (p-value=0.80), $F_{WT}(3,24)$=0.616 (p-value=0.61)

$F_{aut}(1,23)$=2.13 (p-value=0.15), $F_{het}(1,27)$=0.16 (p-value=0.68),

$\chi^2_{nor}(2)$=4.19 (p-value=0.12), $F_{reset}(1,23)$=0.23 (p-value=0.63)

Note: $F_{WT}(2,24)$ is the Wald test for linear restrictions: zero on constant and $\Delta\Delta q_t$, while $F_{WT}(3,24)$ is the Wald test for zero on constant, $\Delta\Delta q_t$ and Δu_t. F_{aut} is the Godfrey's test for autocorrelation, F_{het} is the White test for the presence of heteroscedasticity, χ^2_{nor} is the Jarque-Bera test for normality, F_{reset} is the Reset test for functional form. t-statistics of the coefficients are in brackets (.).

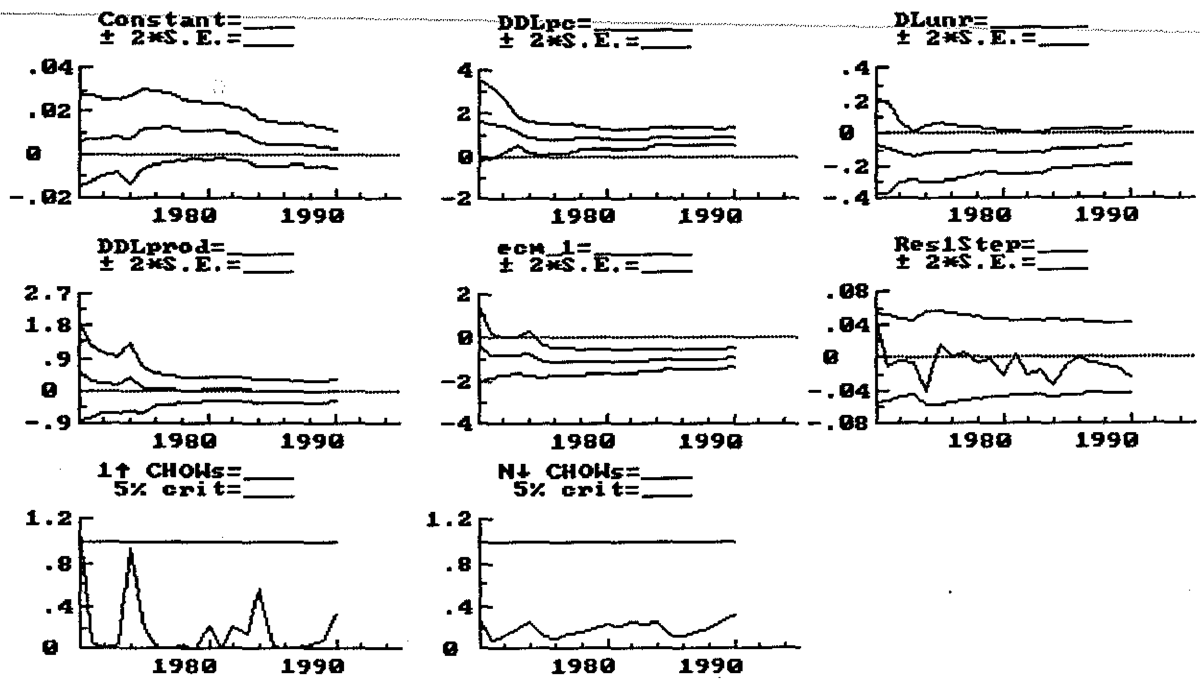

Figure 6 Analysis of the model's constancy using recursive estimates

4. Conclusions

This paper analyses dynamic relations between wage and price inflation, unemployment rate and labour productivity on Italian annual data (1960-1990) using the heading of cointegration and a battery of misspecification tests and tests for parameter constancy.

In order to model comovements in the series, we selected a satisfying (in terms of diagnostic tests and tests for constancy) and parsimonious (via the tests of model reduction) VAR specification. We applied Johansen procedure to test the presence of stochastic trends in these variables (multivariate integration analysis) and their long run comovements taking into account two cases of relation between the constant term and the reduced rank matrix (unrestricted and restricted intercept in the cointegrating space). In both cases we found, via the unadjusted tests of cointegration, two long run elationships between the variables.

Next, we performed tests for structural hypothesis lying in the estimated cointegrating space. These tests gave evidence for the existence of a significant trade-off between wage inflation and unemployment. In order to tackle with the problem of normalisation and identification of the two cointegrating vectors we also performed weak exogeneity tests on the adjustment coefficients to the (dis-)equilibria. From these tests it turns out that one of the two cointegrating vectors can be identified as a Phillips-type equation. As far as the second cointegrating vector, tests on the adjustment coefficients lead to the rejection of the interpretation of this vector as a price mark-up equation, reverting the results obtained by testing structural hypothesis on the long run coefficients. Furthermore, zero restrictions on the coefficients of adjustment lead to identify a short run representation for the wage equation.

The final ECM representation for the wage equation is reasonably stable and passes diagnostic tests showing that, in the short run, the acceleration of wage inflation largely depends on the contemporaneous inflation of prices and on the disequilibrium represented by a long run Phillips relation.

The overall conclusion from this work seems to be that a trade off between a nominal variable (wage inflation) and a real variable presumably related to the business cycle (unemployment) exists and is reasonably stable in the form of an ECM process. As far as the inflation process, the data set considered in this study only allows to take into account the inflationary effect of labour market disequilibria. The weak exogeneity in price inflation suggests that this variable never adjusts to the disequilibria in the long run relationships. Furthermore, there is evidence of no-Granger causality from wages to price inflation. These results do not support the simple price mark-up view of inflation process and suggest that a richer information set is required to model inflation in the Italian economy, to analyse the relevance of the monetary channel and of the external channel as determinants of price inflation (as, for example, Juselius 1991). However, the wage-type Phillips curve model performs rather well on the sample at hand and past inflation and past labour market disequilibrium, as tracked by the unemployment rate, can be used to predict future wage growth.

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