Robert Schuman Centre

European Unemployment:
Macroeconomic Aspects

Wage Setting, Wage Curve
and Phillips Curve:
The Italian Evidence

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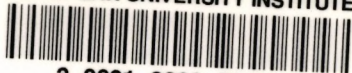and
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EUROPEAN UNIVERSITY INSTITUTE, FLORENCE

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European Unemployment: Macroeconomic Aspects

Wage Setting, Wage Curve and Phillips Curve: The Italian Evidence

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Abstract

The purpose of this paper is to investigate some issues of wage setting in order to assess if nominal inertia and wage flexibility characterize the Italian supply side, using multivariate cointegration models. Our estimates indicate that an explicit distinction between stationary and non-stationary variables and a joint analysis of long-run and short-run structure is crucial for achieving clearer results. To this end, we use quarterly time series data for manufacturing industry 1976:1-1993:4. Interesting results have been found concerning the empirical evidence of a long-run wage curve and the existence of a Phillips curve, through adopting alternative order reduction of the I(2) wage and price variables. Moreover some insights on regional (North-South) unemployment effects are pointed out and discussed.

JEL: J31

Introduction[1]

The purpose of this paper is to investigate some issues of wage setting in order to assess if nominal inertia and wage flexibility characterize the Italian supply side, using multivariate cointegration models. The literature on wage curve or wage setting in Italy is somewhat limited[1]. Some information comes from multi-country analyses (Bean et al. 1986; Grubb 1986; Layard at al 1991), other from macro-econometric models of the Italian economy (e.g. Modigliani et al. 1986; Baici 1992; Chiarini 1993; Terlizzese 1993; Tivegna 1993), but the existing empirical evidence on wage responsiveness to unemployment, wedge and labour productivity is rather unsatisfactory.

Several studies show significant effects of unemployment, others find no unemployment effect at all. Moreover, results on productivity and wedge effects are not homogeneous. All of this has obvious implications for macro-policy. For instance, in fighting against inflation", different policy mixes between pegged exchange rate, fiscal austerity and monetary restriction have been endorsed as alternative strategies to eliminate inflation differential, but how long it will take for the economy to shrink the initial differential and how much it will cost in terms of unemployment and output depends very much on nominal rigidities (see, Blanchard and Muet 1993 and Chiarini 1996 a,b, among others).

Our estimates indicate that an explicit distinction between stationary and non-stationary variables and a joint analysis of long-run and short-run structure is crucial for achieving better results. To this end, we use quarterly time series data for manufacturing industry 1976:1-1993:4.

The paper is organized as follows. In Section 1 the economic background is reported briefly. In Section 2 an unrestricted VAR model is specified and the analysis of cointegration is discussed. This is done by first estimating the cointegration space (Johansen 1988; Johansen and Juselius 1990) and then testing more specific hypotheses of economic interest within this space (Johansen and Juselius 1992; 1994). As a crucial issue, we replace the I(2) variables, wages and prices, with the I(1) alternative real wage. In Section 3 a "structural model" for the determination of the Italian wage is estimated using the derived short-run VAR in error correction form (VECM) with the cointegration relationships explicitly included. Section 4 goes back to the VAR in the I(1) space, adopting a

different solution for I(2) variables, wages and prices, focusing on the outcome of relaxing the short-run homogeneity assumption. In Section 5 an impulse-response analysis of the systems is illustrated. Results and implications are drawn in the economic identification paragraph at the end of the latter two sections, while some concluding comments can be found in Section 6.

Model diagnostic tests and cointegration analysis results are gathered in the Appendix.

1. Economic Background

The wage is determined through collective bargaining. Unions and firms (employers' federation) bargain over wage (W), and possibly employment (N), and firms retain the right to manage. Conventional bargaining models can be utilized, such as the following asymmetric Nash bargaining solution: max $[U (W,N)- U°]^{\alpha} [\Omega (W,N)-\Omega°]^{1-\alpha}$, with $0\leq\alpha<1$ and where U° and $\Omega°$ stand for the status quo outcomes, and U is the utility for unions and W° is the profit for firms. This *mechanism* of determining wages provides "the most general form of wage equation" (e.g. Nickell 1988; Layard and Nickell 1985). We may assume that real profits are a function of (Y/N) and (W/Q), Q is the price of output: $\Omega* = \Omega$ (W/Q,Y/N)K, whereas the union utility function has the following form, $U = NV + (N^s -N)V^*$, where V and V* are the union members' utilities for, respectively, the employed and the unemployed. These utilities are functions of the consumption wage and elements of the wedge between consumption and product wage. V* should also contain a set of variables that may influence the workers' well-being. The result of the Nash-bargaining solution is a general wage function of the form:

$$W/P = W [Y/N, Z, U, P*E/P]$$

where Z is a set of wage pressure elements and P*E/P is the level of competitiveness.

Thus the bargain outcome is real wage which, in turn, depends on productivity and on a set of indicators of the outside opportunities for workers. Similar solutions may be provided by different models (e.g. efficiency-wage; monopoly union model) and may be related to Sargan (1964) and Kuh (1967) type equations and the standard Phillips curve (e.g. Nickell 1988, Manning 1993, among others).

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2. A VAR Model for Wage-Setting

Data are quarterly, seasonally adjusted over the period 1976:1-1993:3. The system is in five stochastic variables $x' = [(w-p), \pi, (p-q), u, h,]$ with the intercept restricted to lie in the cointegration space. We use w_t s the wage variable, earnings per employee in the Italian manufacturing industry.[2] The price variable p_t is the consumption prices index, while q_t is the output price (value added deflator), u_t stands for the aggregate unemployment rate and π_t for the average labour productivity in the manufacturing industry.

The vector x' includes the (per capita) hours worked in the "large firms" h_t (firms with more than 500 employees) as a proxy for hours worked in the manufacturing sector. All the variables are in logs (the log of the unemployment rate represents the convex relationship between labour market pressure and wage growth), and non-stationary time series. However, for the variables w_t and p_t, conventional unit root tests and graphs of their first differences do not suffice to rule out that they are I(2) whereas the real wage $(w-p)_t$ appears I(1) (see Table A.1 in the Appendix).

The VAR is augmented by two shift dummy variables: Dummy1(=1979:2-1980:4=1) and Dummy2 (=1981:4=1;1982:1=1), which have been set as unrestricted. The first dummy accounts for the short run effects concerning the second oil shock, whereas the second one takes into account two anomalous data relative to the working hours variable.

The model is conditioned on S_t and S_{t-1}, where S measures the hours lost for labour conflicts. This variable (I(0)) is presumed to be weakly exogenous and to have only short-run effects.

In order to test for cointegration, we conduct our analysis using a VAR with 2 lags on all stochastic variables. The appropriateness of the lag order was tested commencing with a system lag length of four (using likelihood ratio test adjusted for degree of freedom).
With k=2 lags, the VAR model can be represented in a vector error-correction (VECM) form:

$$\Delta x_t = \Gamma \Delta x_{t-1} + \Pi x + \Psi D_t + u_t \; ; \; u_t \sim N\,i\,i\,d\,(0,\Sigma) \qquad (1)$$

where $\Gamma_i = -(I-A_1)$, $\Pi = -(I-A_1-A_2)$ and A_i is an (nxn) matrix. The vector D_t is a set of conditioning variables: non-stochastic variables such as intervention or shift dummies and others that are weakly exogenous. The dimension of Π is (nxn) and

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if its rank is equal to r<n, a representation of Π such that $\Pi = \alpha\beta'$ exists, where α and β are both (nxr) matrices.

This model is a cointegrating transformation of a VAR model (see Hendry 1995; Johansen 1995, among others). It can be regarded as a multivariate generalization of a model in differences with an ECM and if r<n, the matrix $\beta' x_{t-2}$, embedded in (1), represents up to (n-1) error correction mechanisms in the multivariate model, which ensure that the x_t converge to their long-run steady state solutions.

The estimation is carried out over the period 1976:1-1993:3 using recursive least square (RLS). Diagnostic tests (reported in Table A.2 in the Appendix) show a good descriptive power of the system.[3] None of the test is significant at the 1% level and only an autocorrelation statistic for h_t matters at 5%.

Cointegration Analysis

Testing for cointegration in the five-equation system provides the result reported in Table A.3, where μ stands for eigenvalue, Max and Tr are, respectively, the maximum eigenvalue and the trace statistics. As in the standard notation, the β matrix represents the estimated cointegrating vectors and the α matrix contains the relative weights (i.e. the adjustment coefficients).

Both the test statistics, testing the null of no cointegration, suggest that there is evidence of two cointegrating vectors (and three common trends) in the data set, given by the first columns of the β matrix. The hypothesis is supported by the test statistics adjusted for degrees of freedom (see Osterwald-Lenum 1992 for the asymptotic critical values) and by the analysis of the *companion matrix* of the VAR model (the matrix A of equation (1)). Following the suggestion of Juselius (1995), where the model is conditioned on dummies and weakly exogenous variables, we should look for additional information that can support the choice of stationary relations. In fact, variables included in the conditioning set D_t affect the asymptotic distributions and, therefore, the critical values provided by the Johansen procedure (and others) can be considered only as indicative.

All the nxk = 10 roots of matrix A are inside the unit circle and the three largest roots have as modulus, 0.9845, 0.9151 and 0.9151. (the fourth largest root is 0.576). This supports the choice of (n-r)=3 common stochastic trends and two stationary relations.

It is well known that any linear combination of the stationary vectors, found in the unrestricted cointegration space, is also a stationary vector. This implies that a

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direct interpretation of β_i may not necessarily be interesting from the economic point of view. However, the analysis of the individual β_i shows that the first unrestricted vector looks like a long-run wage function with theoretically expected signs. The second unresticted vector may be interpreted as a price equation. From the α matrix it seems that β_i should not appear in the equations for Δu_t and Δp_t, the variables being weakly exogenous for the long run relations.

Chart 1 (in the Appendix) plots the error corrections for the six components of the β matrix, showing that the first two (cointegrating vectors) look fairly stationary, and the disequilibrium is not large.

Identification restrictions were attempted on the unrestricted cointegrating vectors. In particular we test a number of structures, theoretically motivated, on

$$((w-p), \pi, u, (p-q), h, \text{const}) \in Sp(\beta)$$

In particular, the hypotheses tested were expressed as hypotheses about a restricted vector in the space spanned by the first two cointegration vectors, and they may classified in groups: in the first group, the stationarity of the real wage and each of the variables is tested (for instance, is real consumption wage around the productivity trend stationary? and so on). In the second group, hypotheses regarding whether an estimated combination of unemployment and productivity and a combination of unemployment and price wedge, cointegrate. A further group involves the working hours variable and the real wage. The likelihood ratio tests, asymptotically distributed as χ^2 with the appropriate degrees of freedom, indicated that stationarity had to be rejected for all the cases.

These results imply that combinations of at least three variables are needed to find cointegration. Therefore, we imposed r-1 exactly identifying restrictions and a normalizing coefficient on each cointegration vector, raising a number of testable questions (for instance, whether real consumer wage around the productivity trend cointegrates with the unemployment rate). This process ends up with the following description of the cointegrating space:

β_1 $(w-p) = \pi - 0.114u - 1.9(p-q) - 0.30h - 5.41\gamma$.
β_2 $(w-p) = 0.483\pi + 0.277h - 6.07\gamma$.

where γ represent a constant term.

Two restrictions were imposed on the seond vector and one on the first. The LR test for these hypotheses, distributed as a χ^2 (1) under the null gives a value of 0.2266 [0.634].

The restrictions on identifying the two separate cointegrating vectors have been tested jointly with testing for u and π being long-run weakly exogenous for both the stationary relations. The LR test for these joint hypotheses is χ^2 (4) = 7.242 [0.203], indicating that we cannot reject these reductions at a 5% level.

An Economic Interpretation of Cointegration Vectors

Dickey, Jansen and Thornton (1991) emphasize that, in general, cointegration vectors cannot be interpreted as representing structural equations because they are obtained from the reduced form of a system where all of the variables are jointly endogenous.

Structural results are difficult to extract from the reduced form equations. However, they could be thought of as representing constraints that an economic structure imposes on the movement of the variables in the system in the long-run. Thus, we have two linear combinations for which the variance is bounded. To this regard, Johansen (1995c) points out that the foremost problem that one should consider is to build a statistical model that describes the data well, and express the economic theory in terms of the parameters of the statistical model.

If a structural relation is modelled by stationary relations (cointegrating relations) and if the property of stationarity and non-stationarity can be deducted from the reduced form, then the cointegrating relation captures the economic notion of a long-run economic relation. In fact, the multivariate model (1) is given in ECM form, but it is essentially a statistical description of the data. When the data are I(1), the process $\{x_t\}$ contains information about the equilibrium forces and the adjustment towards long-run stedy states and the driving forces that push the economy out of equilibrium (see also Juselius 1993).

The two cointegration relations seem support the existence of a long-run real wage relation. According to Johansen and Juselius (1994), it seems plausible to assume that the long-run structure may cointain evidence about two behavioral relations (in this case, unions versus firms). Therefore our stationary relations may be interpretable, respectively, as a "supply" (firm) real wage relation and a "demand" (union) real wage relation. This interpretation seems to be supported by the fact that the first relation includes variables correlated with the "supply" (the wedge between consumer and producer prices and the unemployment rate) and

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uncorrelated with the "demand". These hypotheses have been tested on the coefficients as over identifying restrictions.

Furthermore, the different sign of the coefficient found for the worked hours variable in the two relations, seems consistent with this interpretation. If hours are increased (and employment reduces for a while), wages pressure will reduce and the amount of work available will have to be increased. Employment will revert to its former level. However, a rigid labour market may obstruct these movements.

The last question is whether the estimated coefficients are economically meaningful. Cointegrated systems must be interpreted cautiously. The coefficients of the cointegrating relation cannot be interpreted as elasticities. This is true even if the variables are in logs, because all the other dynamic relations between the variables which are specified in a VAR model are ignored. Lütkepohl (1991; 1994) shows that the *ceteris paribus* assumption may not have a meaning. The analysis requires that short-term dynamics and intertemporal adjustment processes generated by equilibrium errors are taken into account. Inpulse response analysis, taking into account the full system, may provide a more reliable conclusion. This means that the coefficients provided by the cointegration analysis are only indicative.

To summarize, the analysis of the long-run structure of the data in terms of stationary cointegration relations provides the following results:

(i) the hypothesis that real consumer wages cointegrate with labour productivity is rejected as is the stationarity of real wages and the unemployment rate, whereas the hypothesis that these two cointegrate with price wedge and working hours is accepted. This relation cannot interpreted individually. A further equation describes the equilibrium relations between the five variables.

(ii) The possibility that the "wage gap" (defined as $(w-p)-\lambda\pi$) cointegrates with the working hours variable is not rejected by the data.

(iii) The low degree of wage flexibility implies weak equilibrating mechanisms.

(iv) Finally, it is worth stressing that restricting the unemployment effect to zero in both cointegrating vectors is rejected by the data: $\chi^2 (2)=13.972[0.0009]$.

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3. The Short-Run Structure

Now we map the data to $I(0)$ series using first differences of the variables, reformulating the model with the error-correction terms explicitly included, ECM_1 and ECM_2. The system determines four variables: $\Delta(w-p)_t$, $\Delta(p-q)_t$, Δh_t and Δu_t. Table A.4 (in the Appendix) reports some model statistics of the parsimonious VAR model (PVAR), obtained dropping the variables with F-tests for the nullity on the retained regressors and adding two identities for the cointegrating vectors. The PVAR satisfies the properties of parameter costancy (see Chart 2: 1-step residuals with $\pm 2\sigma$ and individual equation break-point Chow test), normality and homoskedasticity. The multivariate test for autocorrelation is significant at the 5% level; mostly due to the proxy variable h_t.

Imposing specific restrictions on each equation (dropping insignificant regressors), we procede to estimate a simultaneous (structural) model using the FIML procedure (Table 1). The LR test of over-identifying restrictions indicates that the structural model encompasses parsimoniously the conditional PVAR: $\chi^2 (26)=31.23[0.220]$. The multivariate tests suggest that the model has approximately white noise, normally distributed errors.

It should be noted that the short-run behaviour of the weakly exogenous variable p_t is not modelled, being zero the coefficients of Δy_{t-1} [$y'= (w-p), (p-q), h, u$] and therefore, strongly exogenous: y_t is not the Granger cause of π_t. On the contrary, Table 1 shows that although the change in unemployment rate does not react to disequilibrium errors, it still reacts to lagged changes of real wage.

TABLE 1

Real wage model. Structural Model (FIML)

Equation 1 for $\Delta(p-q)$

Variable	Coefficient	t-vaue
Δp_1	-0.13378	-2.429
ECM1_1	-0.15382	-4.959
ECM2_1	0.12068	3.025
S_1	-0.0088173	-3.518
Dummy1	-0.011042	-3.614
Δu_1	0.039776	3.211

Equation 2 for $\Delta(w-p)$

Variable	Coefficient	t-vaue
$\Delta(p-q)\_1$	-0.43981	-1.881
$\Delta(w-p)\_1$	-0.31435	-3.103
S_1	0.012877	1.921
Dummy1	0.025202	3.118
ECM2_1	-0.14124	-3.214

Equation 3 for Δu

Variable	Coefficient	t-vaue
$\Delta(w-p)\_1$	0.48222	1.828
Δu_1	-0.26155	-2.286
$\Delta\pi\_1$	-0.75636	-1.761
S_1	0.042684	2.194

Equation 4 for Δh

Variable	Coefficient	t-vaue
Δu_1	0.11372	3.204
$\Delta\pi$	0.61845	4.532
$\Delta\pi\_1$	-0.20539	-1.321
ECM1_1	-0.10975	-1.228
ECM2_1	0.24935	2.173
Dummy2	0.046912	5.691

Vec.AR1-5 $F(80,160)=1.4394$ [0.0267]*
Vec.normality $\chi^2(8) =9.6676$ [0.2891]
Vec.HET $F(210,313)=0.92526$ [0.7275]

Economic Identification

The dominant features of the estimated model are:

-**i**) lagged unemployment and labour productivity do not enter significantly in the wage equation, while the long run equation shows a long-run productivity coefficient equal to unity.
The lack of evidence of short-run effects of productivity changes on real wages implies that a fall in productivity growth determines a wage gap (as in the 1980s).

-**ii**) The cointegration vector β_1 shows that unemployment carries a long-run coefficient of 0.114; moreover, it seems to be weakly exogenous for the parameters of both the stationary relations.

Hence, our empirical analysis implies considerable less long-run wage responsiveness to unemployment as compared to the findings in many multicountry studies.

Changes in the unemployment rate do not affect wage growth, whereas changes in the latter increase the unemployment rate. This result will be confirmed by the inpulse-response analysis (carried out below, see Charts 5-5.3) and can be found only in a multivariate context, often neglected by the macroeconomic literature.

Following Nymoen (1992), one may state that if real wage does not drift too far from productivity over time, that is if the wage gap w-p-$\lambda\pi$ is stationary, then there is wage flexibility. Our results confirm serious rigidity, showing that wage gap cointegrates with unemployment, worked hours and relative prices.

-**iii**) We find, however, support for the wage curve (as defined by Blanchflower and Oswald 1994a,b): a relationship between the level of (real) pay and the unemployment (and local unemployment) rate. Although here the wage curve is a long-run equilibrium relationship, the estimated unemployment coefficient (-0.11) appears to be approximately the same as that found in many countries through the microeconometric analysis of Blanchflower and Oswald.[4] However, as discussed in Section 2, this is only indicative, the coefficients do not represent elasticities. Considering the short-term dynamics of the full system we reach a quite different result (see Section 5).

It is useful to conclude the discussion on wage-unemployment relationship mentioning some limitations on analysis. When bargaining is centralized or co-ordinated, unemployment takes on a key role. We should expect, therefore, a powerful impact of unemployment on wages. However, in the short-run this

depends on the extent to which wages are indexed. Moreover, the Italian labour market presents a remarkable dualism, with the Northern labour market close to full-employment and the Southern regions characterized by an unemployment rate of around 20%.[5] This indicates that some "dispersion" variable should be included in the wage equation to capture a short-run effect for demand pressure (e.g. the seminal paper by Archibald 1969).[6]

The importance of the dualistic nature of the Italian labour market has been investigated introducing the unemployment rate of the northern regions of Italy into the model. The long-run equilibrium model is found to be not so different from the standard model discussed above. Three main results emerge from the FIML estimate. First, the long-run equilibrium unemployment coefficient is -0.103. Second, changes in the northern unemployment rate do lead to wage reactions in the short run. Third, the short-run unemployment rate is no longer determined by changes in real wage.[7]

On the contrary, the southern unemployment rate is insignificant both in the short- and long-run (see also Bodo and Sestito 1991). Obviously all of this has strong policy implications. For instance, if wage increases are determined in low-unemployment regions one may reduce unemployment in other regions without worsening wage inflation. This policy intervention, involved by a "kinked" Phillips curve, horizontal at the wage growth corresponding to the southern regions, is extremely difficult in Italy, where national and sectorial wage setting is not a function of regional unemployment and productivity.[8]

-iv) An implication of the above conclusions is that outsiders having no role in wage determination may generate high unemployment persistence (e.g. Blanchard and Summers 1986, Lindbeck and Snower 1987, Blanchard 1991).

-v) As expected, real wage adjustment to deviations from disequilibria is rather sluggish: the estimated speed of adjustment parameter is -0.14.

-vi) The short-run effect of the wedge (p-q) is rather strong; the estimated short-run dynamic impact is -0.34. A negative and permanent (long-run) price effect is also provided by the cointegration restrictions. Any expansionary macro-policy which affects the relative price between domestic output and consumption, reducing the real product wage, increases employment.

-vii) The hours lost for labour conflicts have a positive (although modest) short-run effect on real wage and unemployment, and a negative effect on the wedge growth.

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-viii) Both unemployment and productivity are significant in the relative price equation but the signs seems to suggest a counter-cyclical effect.

4. An Alternative Wage Model

In a previous section, we noted that the nominal variables w_t and p_t might be I(2). From the graphs of the data and unit roots tests it seems indeed that the first differences of these variables could be descibed by a I(1) process.

Instead of introducing the variable $(w-p)_t$ one can eliminate the I(2) component working with inflation rate Δp_t and wage growth Δw_t . It is plausible, from the economic point of view, to assume lack of homogeneity in the short-run, although this implies a short-run wage and price equation in second differences.[9]

The variables used in this model are the same as for the real wage model. However the model is characterized by a new conditioning set D_t , which includes two dummies: dummy3 (1979:4-1980:1=1) to account for the second oil shock and dummy4 (1981:1=1) to cope with an outlier. With the variables Δw_t , Δp_t , π_t, u_t , $(p-q)_t$, h_t we can repeat the I(1) analysis. The results are reported in Table A.5 in the Appendix.

The trace test and maximal eigenvalue test now suggest the existence of three cointegrating vectors in the data, whereas test statistics adjusted for degrees of freedom suggest two stationary relations. The additional information given by the 12 roots of the companion matrix seems to support three common stochastic trends (the first largest root, 0.987, is followed by a complex pair of roots with modulus 0.86) and, therefore, three cointegration relations (see Chart 3).

It is interesting to note that, unlike the unrestricted estimates β_i of the real wage model, now the unrestricted eigenvectors can be given a direct interpretation in terms of hypothetical relations. The first β vector primarily reflects a price equation. The second cointegration vector seems to suggest a stationary relation describing the wedge (between consumer and producer prices) long-run determinants, while the third vector suggests a stationary relation between wage and price growth, and unemployment and productivity.

By imposing some restrictions on each cointegration vector and on the associated a matrix we obtain the following long-run relations as economically acceptable long run price, wedge and wage equations:

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$$\Delta p = \Delta w + 0.479(p\text{-}q) - 0.09\pi$$
$$(p\text{-}q) = \Delta p - 0.105\Delta w + 0.236\pi - 0.105h$$
$$\Delta w = \Delta p - 3.92(p\text{-}q) - 0.17u + \pi - 0.82h$$

The restricted cointegration vectors, tested for lying in the cointegration space yield χ^2 (12) = 9.9131 [0.6236]. Three restrictions were imposed on the first vector and two on the second and third vector.

From a statistical point of view, testing for cointegration between variables (differences) that require to be differenced twice to induce stationarity and other I(1) variables (levels) is appropriate. From an economic point of view, determining a general formulation of overidentifying restrictions on the individual equations is more difficult.

The first cointegration vector describes price growth. Long-run movements in prices and wages are correlated. Moreover, we find that long-run increases in the ratio between producer and consumer prices in part will be counteracted by an increase in the real wage growth. There is a long-run tendency for the wage growth to follow labour productivity. The second cointegration vector suggests a positive association between the wedge between consumer and producer prices and labour productivity and a negative relation between the price wedge and the wage growth. This seems consistent with a price setting based on unit costs. These variables cointegrate with the worked hours variable. The third stationary relation is a plausible long-run wage equation. A Phillips curve augmented by productivity and prices. A rigid labour market may lead firms to use working hours which reduce wage increases. In the long-run, rises in the wedge are born entirely by labours (see Modigliani et al 1986; Padoa Schioppa 1986 for permanent tax effects for Italy. See also Layard et al. 1991). This result confirms the one found for the real wage model.

Notice that in this model the constant term does not appear in the cointegrating relations. This result may indicate that static equilibrium is a very restrictive long run solution for this model. According to this explanation, the constant term in the long run equation of the previous (real wage) model depends on the steady state growth rate of the exogenous variables. On the contrary, the present model does not include exogenous variables in the long-run, and rates of wage and price growth are explicitly included in the cointegration analysis.

Parsimony has been achieved by removing the insignificant regressors in the VAR model in I(0) space, and testing whether this reduction in the model is supported by an F-test. The diagnostic results from estimating the parsimonious version of the model are given in Table A.6 in the Appendix. Multivariate tests

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and the recursive properties of the model (Chart 4) are satisfactory. The model has constant coefficients and white noise, normally distributed errors.

The structural model, estimated in FIML, is reported in Table 2. The model encompasses parsimoniously the PVAR χ^2 (29)=19.34 [0.9120].

TABLE 2: Nominal wage model. Structural Model (FIML)

Equation 1 for Δ(p-q)

Variable	Coefficient	t-vaue
Δu_1	0.048909	3.520
ΔΔp_1	-0.18599	-2.220
Δπ_1	-0.15943	-2.847
ECM1_1	-0.0086276	-1.348
ECM2_1	-0.019320	-5.093
Constant	-0.16212	-4.273
S_1	-0.0043868	-1.621

Equation 2 for Δu

Variable	Coefficient	t-vaue
Δu_1	-0.28731	-2.784
Δπ_1	-0.62392	-1.446
ΔΔw_1	-0.90274	-3.123
ECM1_1	-0.25510	-3.838
dummy3	-0.073164	-1.986
Constant	0.51264	3.867
S_1	0.052859	2.323

Equation 3 for ΔΔp

Variable	Coefficient	t-vaue
Δu_1	0.031799	2.670
ΔΔp_1	-0.30397	-3.941
ΔΔw_1	-0.082128	-2.540
ECM1_1	-0.037223	-4.967
ECM2_1	0.014461	3.289
ECM3_1	-0.12472	-4.566
Δπ_1	-0.15660	-3.288
dummy3	0.021984	5.541
Constant	0.28971	5.407
dummy4	0.044121	1.892

Equation 4 for ΔΔw

Variable	Coefficient	t-vaue
Δ(p-q)_1	-0.54046	-2.049
ΔΔp_1	-0.62920	-3.065
ECM1_1	-0.16543	-10.662
ECM2_1	0.035973	10.879
ECM3_1	-0.29211	-6.264
dummy4	0.026180	3.867

Equation 5 for Δh

Variable	Coefficient	t-vaue
Δu_1	0.091435	2.106
ΔΔp_1	0.58008	2.022
Δπ	-0.67234	-3.708
ECM2_1	0.047000	2.314
ECM3_1	-0.30957	-3.011
Constant	0.42958	2.253

Vector AR 1-5 F(125.162)= 1.2749 [0.0731]
Vector normality $\chi^2(10)$= 12.47 [0.2548]
Vector HET F(345.339) = 0.93737 [0.7253]

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Economic Identification

There are a number of features worthy of comment:

i) as in the real wage model, it is clear that the phenomenon of the unemployment hysteresis does appear in the wage model.

ii) As in the real wage model, unemployment matters only in the long-run, and is inversely related to wage growth (with a coefficient of -0.17).
We cannot restrict the long-run Phillips curve to be vertical: setting unemployment to be zero throughout the cointegrating space yields χ^2 (5)=38.6396[0.000]. Thus, considering a vector time series which follows a VECM process, we cannot conclude that the Italian data are strongly supportive of the wage curve rather than the Phillips curve.

iii) Using regional (North-South) unemployment rates we may prove results analogous to the real wage model ones: the northern unemployment rate has a slightly stronger effect than the aggregate unemployment rate, whereas the restriction to zero of the coefficient for the southern unemployment rate is not rejected by the data (results are provided upon request).

iv) Unlike the real wage model, unemployment cannot be considered weakly exogenous for the equilibrium parameters, whereas labour productivity should be considered exogenous to the system.

v) The effect of "unexpected" changes from wage and price setting, are rather different in the respective equations (see also the impulse-response analysis reported below). An increase in wage growth yields a negative dynamic impact of 0.082 on the acceleration of prices, that relative to a change in inflation on the acceleration of wages is 0.629. The extent of nominal inertia in wage setting depends negatively on the extent of indexation (for most of the sample Italy is characterized by a high wage indexation). Inertia is also inversely related to the extent of synchronization and wage contract renewals in Italy are more or less synchronized in the manufacturing sector. With stickier nominal wages (a partial and backward-looking indexation mechanism limited the real wage rigidity in the very short-run), real wage decreases in the short run because of unexpected increases in prices. Short-run price homogeneity does not seem to be present.

vi) There is evidence (although the t-value is rather low) of a negative relationship between unemployment changes and favourable productivity shocks. Further notable features of the unemployment equation are the strong effects of wage acceleration, and a positive effect of the hours lost for labour conflict.

5. Impulse Response Analysis

In charts 5-5.3 we show the responses of the variables to a one-time impulse in, respectively, $\Delta(w\text{-}p)_t$, $\Delta(p\text{-}q)_t$, Δu_t and Δh_t, of size one standard deviation. Therefore, the units at the vertical axes equal to the standard deviations of the residuals corresponding to the variables whose effects are considered. Impulse response analysis amount to dynamic simulation, from an initial value of zero, where a shock at t=1 in a variable is traced through (see. Lütkepohl 1991; Lütkepohl and Reimers 1992).[10]

For instance, in the first chart, a real wage innovation is seen to induce the unemployment rate to grow for one period and then it converges to zero. The impulse leads to a permanent (long-run) increase in the level of u_t. This is a consequence of the nonstationary features of the data. Unemployment seems to settle at a different equilibrium value. The real wage innovation has no significant effects on Δh_t and $\Delta(p\text{-}q)_t$.

In Chart 5.1 the effects of an impulse in $D(p\text{-}q)_t$ are displayed. This leads to a short-run decline in real wage and, therefore, the shock improves unemployment and worked hours.

The effect of the shock in any of the variables dies away quite rapidly in the sense that the differences of all the endogenous variables return to their previous values if no further shocks hit the system. This means that the system is stable. Notice, however, that the impulse in $\Delta(p\text{-}q)_t$ generates significant permanent effects on $(w\text{-}p)_t$ and u_t. These variables do not return to their previous equilibrium values. The dynamics of the ECM_i contains information about the equilibrating forces that make the process (the variables) adjust towards long-run steady states.

Chart 5.2 shows there is no significant effect on real wage due to an increase in Δu_t. These results indicate a low degree of wage flexibility. Therefore, unemployment does not Granger-cause the other variables of the system whereas the converse is true. Finally, the response of $\Delta(w\text{-}p)_t$ to an impulse in Δh_t is clearly insignificant.

Innovation responses of the nominal wage model are depicted in Charts 6-6.2. The reaction of the variables does not contrast sharply with the previous results, although there is an obvious difference in the dynamic reaction of the variables $\Delta^2 p_t$ and $\Delta^2 w_t$, and in the speed of adjustment of the ECM_i. An important issue is whether the extent of indexation attenuates nominal inertia in wage setting. Chart 6.2 shows that real wage growth is not negatively affected by unanticipated increases in inflation.

6. Conclusions

The paper investigates the degree of real wage flexibility and real wage resistance in the Italian economy. Since wages and prices are of order I(2), we use two wage models for solving the higher nonstationary order of these variables: we replace the I(2) variables with a I(1) alternative by reformulating the model to consider real wages and, subsequently, by using their first differences. Although the latter specification implies an unusual short-run wage equation in second differences, we show that the main characteristics of wage determination are the same but the models reveal additional information: imposing homogeneity in the short-run may be not appropriate even in a high indexed economy.

Interesting results have been found concerning the empirical evidence of a long-run wage curve and the existence of a Phillips curve. In a multivariate context the short-run Phillips curve disappears: changes in unemployment do not affect wages, whereas increasing wages rise unemployment. As the estimated models imply low (if any) responsiveness of wages to variations in unemployment, they generate unemployment persistence.

Cointegration analysis shows that one cannot reject a long-run trade-off between unemployment and wage growth, but a different result is reached if the full system is taken into account: there is no significant permanent (or transitory) effect on real wage (or wage growth) due to a transitory increase in unemployment.

The unemployment effect on wage essentially reflects the demand pressure in the northern regions whereas the remarkable unemployment in the southern regions is negligible.

Notes

1) See also Padoa Schioppa (1990), Lucifora and Sestito (1993), Baici e Dell'Aringa (1994); Checchi (1995). There is a growing interest on microdata analisys but a few works are available. See, for instance, Dell'Aringa and Lucifora (1990), Erikson and Ichino (1994) among others. They can provide interesting insights on the process of wage setting that cannot be gained from the aggregate data. For instance, the empirical work on wage drift in Italy conducted by Ordine (1996), using annual data (1983-1988) on the published accounts of 105 Italian three-digit industries, shows that insider factors (productivity, inventories, insider workers power) are extremely relevant in determining the wage drift while the rate of unemployment has a prominent role in determining minimum contractual wages.

2) Labour input is measured in the National Account data in standardized labour units. These measure the labour input in terms of full-time workers, taking into account the informal economy (irregular workers; second job; work done by illegal immigrants).

3) In order to keep the length of this paper within reasonable limits, the full output result is not reported. Further results may be provided by the authors upon request.

4) We did not find evidence in favour of strong convexities in the wage curve. See, for instance, K.Johansen (1995).

5) The difference between regional unemployment rates remained steady at around 6 percentage points during the first half of the 1980s. At the beginning of the 1990s the differential reached 14 percentage points.

6) A variable used by Archibald (1969) to measure the degree of the dispersion of the local unemployment rates around the over-all figure is $\sigma^2 = \sum_{i=1}^{n} \tau_{it} (u_{it} - u_t)^2$, where n = number of labour markets, τ_i = proportion of the total labour force in market i and u_i = unemployment rate in that market, but it might be useful to reconsider the past literature on wage inflation and the regional wage structure.

7) Wage and unemployment dyanamics is described by the following equations (standard errors in bracket):

$$\Delta(w-p)_t = -0.312\Delta(w-p)_{t-1} - 0.042\Delta u_{N,t-1} - 0.45\Delta(p-q)_{t-1} + 0.013S_{t-1}$$
$$(0.099) \qquad (0.021) \qquad\qquad (0.229) \qquad\quad (0.006)$$
$$- 0.027\text{Dummy}1 - 0.145\text{ECM}2_{t-1}$$
$$(0.007) \qquad\qquad (0.043)$$

$$\Delta u_{N,t} = -0.248\Delta u_{N,t-1} + 0.244\Delta(w-p)_{t-1} - 1.287\Delta p_{t-1} + 0.046S_{t-1}$$
$$(0.116) \qquad\qquad (0.453) \qquad\qquad (0.578) \qquad\quad (0.026)$$
$$- 0.032\text{Dummy}1$$
$$(0.037)$$

where u_n = unemployment rate (northern regions). The full system is provided upon request.

8) For the problems of the interdependence of wage bargains, the wage-leading regions and wage differentials and the role of actual and potential migration (mobility), see Bodo and Sestito (1991).

9) Software programs incorporating the I(2) features of the Johansen approach (see Johansen 1995(a,b) and Juselius 1995) are not available. Butter and Wijngaert (1992) do not adopt this specification and attempt to achieve a cointegration regression under the presumption that wages, the wage space (price inflations plus productivity growth) and its component parts are of order I(2).

10) This amounts to graphing powers of the companion matrix. See Hendry (1995), Lütkepohl (1991).

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APPENDIX
Table A.1: Unit root Tests

Series	ADF	Critical Values	Model °
p	-1,845	-3,476	C,T
Δp	-4,214**	-3,476	C,T
u	-1,659	-2,904	C
Δu	-3,263*	-2,904	C
(p-q)	-2,064	-3,476	C,T
Δ(p-q)	-5,315**	-3,476	C,T
Δw	-2,087	-2,904	C
ΔΔw	-8,137**	-2,904	C
(w-p)	-2,886	-3,476	C,T
Δ(w-p)	-4,796**	-3,476	C,T
S	-3,027*	-2,904	C
h	-1,712	-2,904	C
Δh	-3,691**	-2,904	C
UN	-1,694	-2,904	C
ΔUN	-2,92*	-2,904	C
Δp	-1,637	-2,904	C
ΔΔp	-4,121**	-2,904	C

° Similar test
C: ADF(5) with Constant;
C,T: ADF(5) with Constant and Trend;

Variables and sources

p	labour productivity	ISTAT
u	unemployment rate	ISTAT
p	consumer price index	ISTAT
q	value added deflator (manufacturing industry)	ISTAT
w	wage (manufacturing industry)	ISTAT
S	hours lost in labour conflicts	ISTAT
h	per capita worked hours**	ISTAT
UN	unemployment rate, North of Italy	ISTAT

* per standard labour unit
** firms with more than 500 employees

TABLE A.2
Real wage model.
Diagnostic Tests: VAR I(1)

---

p   :AR 1-5F(5,50)=1.9606 [0.1009]
(w-p) :AR 1-5F(5,50)=0.78514[0.5653]
(p-q) :AR 1-5F(5,50)=2.0402 [0.0889]
u   :AR 1-5F(5,50)=2.2162 [0.0671]
h   :AR 1-5F(5,50)=3.17   [0.0146]*
p   :Norm. χ^2(2)= 2.0223 [0.3638]
(w-p) :Norm. χ^2(2)= 2.8326 [0.2426]
(p-q) :Norm. χ^2(2)= 1.5567 [0.4592]
u   :Norm. χ^2(2)= 1.248 [0.5358]
h   :Norm. χ^2(2)= 3.9121 [0.1414]
p   :ARCh4 F(4,47)=1.67582 [0.6121]
(w-p) :ARCh4 F(4,47)=0.68242 [0.6076]
(p-q) :ARCh4 F(4,47)=1.2297 [0.3111]
u   :ARCh4 F(4,47)=0.13402 [0.9691]
h   :ARCh4 F(4,47)=0.76195 [0.5553]
p   :HET F(20,34)=0.75588 [0.7427]
(w-p) :HET F(20,34)=0.53151 [0.9311]
(p-q) :HET F(20,34)=1.0437 [0.4439]
u   :HET F(20,34)=0.74645 [0.7524]
h   :HET F(20,34)=0.78967 [0.7074]
Vec.AR1-5 F(125,132)=1.2761 [0.0838]
Vec.normality χ^2(10)=13.146 [0.2156]
Vec.HET F(300,284)=0.73354 [0.9959]

TABLE A.4
Real wage model.
Diagnostic Tests: VAR I(1)

Δ(p-q) :AR 1-5F(5,51)=1.0527 [0.3974]
Δ(w-p):AR 1-5F(5,51)=1.2299 [0.3088]
Δu   :AR 1-5F(5,51)=1.4718 [0.2154]
Δh   :AR 1-5F(5,51)=5.6594 [0.0003]**
Δ(p-q) :Norm. χ^2(2)=2.9758 [0.2259]
Δ(w-p):Norm. χ^2(2)=0.54405 [0.7618]
Δu   :Norm. χ^2(2)=0.26012 [0.8780]
Δh   :Norm. χ^2(2)=0.79904 [0.6706]
Δ(p-q) :ARCh4 F(4,48)=0.10423 [0.9805]
Δ(w-p):ARCh4 F(4,48)=1.2758 [0.2926]
Δu   :ARCh4 F(4,48)=0.20105 [0.9366]
Δh   :ARCh4 F(4,48)=0.77152 [0.5491]
Δ(p-q) :HET F(21,34)=0.51916 [0.9418]
Δ(w-p):HET F(21,34)=0.74014 [0.7637]
Δu   :HET F(21,34)= 0.4298 [0.9776]
Δh   :HET F(21,34)=0.55999 [0.9179]

Vec.AR 1-5 F(80,132)=1.3911 [0.0466]*
Vec.normality χ^2(8)=4.7566 [0.7832]
Vec.HET F(210,249)=0.62076 [0.9998]

TABLE A.3
Real wage model. Cointegration analysis

μ: 0.462179 0.383195 0.207429 0.0882889 0.0675254 4.42744e-015

Ho:rank=p	Max	Max°	95%	Tr	Tr°	95%
p == 0	42.18**	35.97*	34.4	101.9**	86.9**	76.1
p <= 1	32.86*	28.03	28.1	59.71*	50.93	53.1
p <= 2	15.81	13.48	22.0	26.85	22.9	34.9
p <= 3	6.285	5.361	15.7	11.04	9.416	20.0
p <= 4	4.754	4.055	9.2	4.754	4.055	9.2

Max=maximum eigenvalue statistic; Tr=trace statistic;
° =small sample correction

standardized b eigenvectors

p	(w-p)	(p-q)	u	h	Constant
1.000	-0.3237	-3.596	-0.1761	-0.9024	-0.6462
1.777	1.000	-8.360	-0.5438	-2.776	8.977
1.237	-2.476	1.000	-0.04138	0.2554	-13.23
-2.062	4.319	-3.373	1.000	6.232	-10.02
-0.6892	2.078	1.317	-0.8203	1.000	6.088

standardized a coefficients

p	0.03041	-0.03787	-0.08620	-0.002652	-0.01601
(w-p)	0.001281	-0.04810	0.09459	-0.002580	-0.01544
(p-q)	0.1662	0.0008411	-0.003952	0.003121	0.004066
u	0.3289	-0.02597	-0.0007584	-0.05530	0.02876
h	0.03906	0.02812	-0.05071	-0.004680	-0.04582

TABLE A.5
Nominal wage model. Cointegration analysis

μ: 0.653355 0.490702 0.408387 0.0945457 0.0749002 0.0154549

Ho:rank=p	Max	Max°	95%	Tr	Tr°	95%
p == 0	72.04**	59.33**	39.4	166.7**	137.3**	94.2
p <= 1	45.88**	37.78*	33.5	94.68**	77.97**	68.5
p <= 2	35.69**	29.39*	27.1	48.8*	40.19	47.2
p <= 3	6.754	5.562	21.0	13.11	10.79	29.7
p <= 4	5.294	4.36	14.1	6.353	5.232	15.4
p <= 5	1.059	0.8722	3.8	1.059	0.8722	3.8

Max=maximum eigenvalue statistic; Tr=trace statistic;
° =small sample correction

standardized b eigenvectors

p	Δu	Δp	Δw	(p-q)	h
1.000	-0.3562	-6.456	6.412	-2.899	-1.275
-5.376	1.000	-16.61	2.538	17.68	2.640
-0.1520	0.05624	1.000	0.2577	0.8791	0.2935
-1.385	1.820	0.07771	1.000	-6.073	0.4405
-0.05403	-0.1391	-1.603	-0.3532	1.000	1.036
0.8284	0.1551	2.652	1.354	1.409	1.000

standardized a coefficients

	p	Δu	Δp	Δw	(p-q)	h
p	0.002350	0.02890	-0.1580	0.006345	-0.03785	-0.01182
Δu	0.2661	0.006795	-0.6898	-0.07530	-0.03530	-0.02027
Δp	0.02972	0.004064	-0.2625	0.002807	0.01831	-0.0003639
Δw	-0.1848	0.0008371	-0.5120	-0.01251	0.01684	-0.001989
(p-q)	0.01503	-0.02618	-0.1230	0.005652	0.006131	-0.002495
h	-0.01536	0.02449	-0.7022	0.0004720	-0.1252	0.005345

TABLE A.6
Nominal wage model. Diagnostic Tests: VAR I(0)

$\Delta(p-q)$: AR 1- 5F(5, 50) =   0.88934[0.4954]
Δu : AR 1- 5F(5, 50) =   2.1162 [0.0788]
$\Delta\Delta p$: AR 1- 5F(5, 50) =   0.99175[0.4323]
$\Delta\Delta w$: AR 1- 5F(5, 50) =   2.2503 [0.0636]
Δh : AR 1- 5F(5, 50) =   3.3009 [0.0118] *
$\Delta(p-q)$: Normality $\chi^2(2)$=   0.23742[0.8881]
Δu : Normality $\chi^2(2)$=   0.3724 [0.8301]
$\Delta\Delta p$: Normality $\chi^2(2)$=   1.7153 [0.4242]
$\Delta\Delta w$: Normality $\chi^2(2)$=   2.6732 [0.2627]
Δh : Normality $\chi^2(2)$=   7.5719 [0.0227] *
$\Delta(p-q)$: ARCh 4 F(4, 47) =   0.04700[0.9957]
Δu : ARCh 4 F(4, 47) =   0.4073 [0.8024]
$\Delta\Delta p$: ARCh 4 F(4, 47) =   0.59343[0.6691]
$\Delta\Delta w$: ARCh 4 F(4, 47) =   0.49089[0.7424]
Δh : ARCh 4 F(4, 47) =   2.104 [0.0952]
$\Delta(p-q)$: HET   F(23, 31) =   1.679 [0.0888]
Δu : HET   F(23, 31) =   0.89305[0.6055]
$\Delta\Delta p$: HET   F(23, 31) =   1.1362 [0.3650]
$\Delta\Delta w$: HET   F(23, 31) =   1.2077 [0.3080]
Δh : HET   F(23, 31) =   0.44454 [0.9761]
Vector AR 1-5 F(125,132) =   1.1645 [0.1940]
Vector normality $\chi^2(10)$ =   13.286 [0.2081]
Vector HET   F(345.263) =   0.6968 [0.9992]
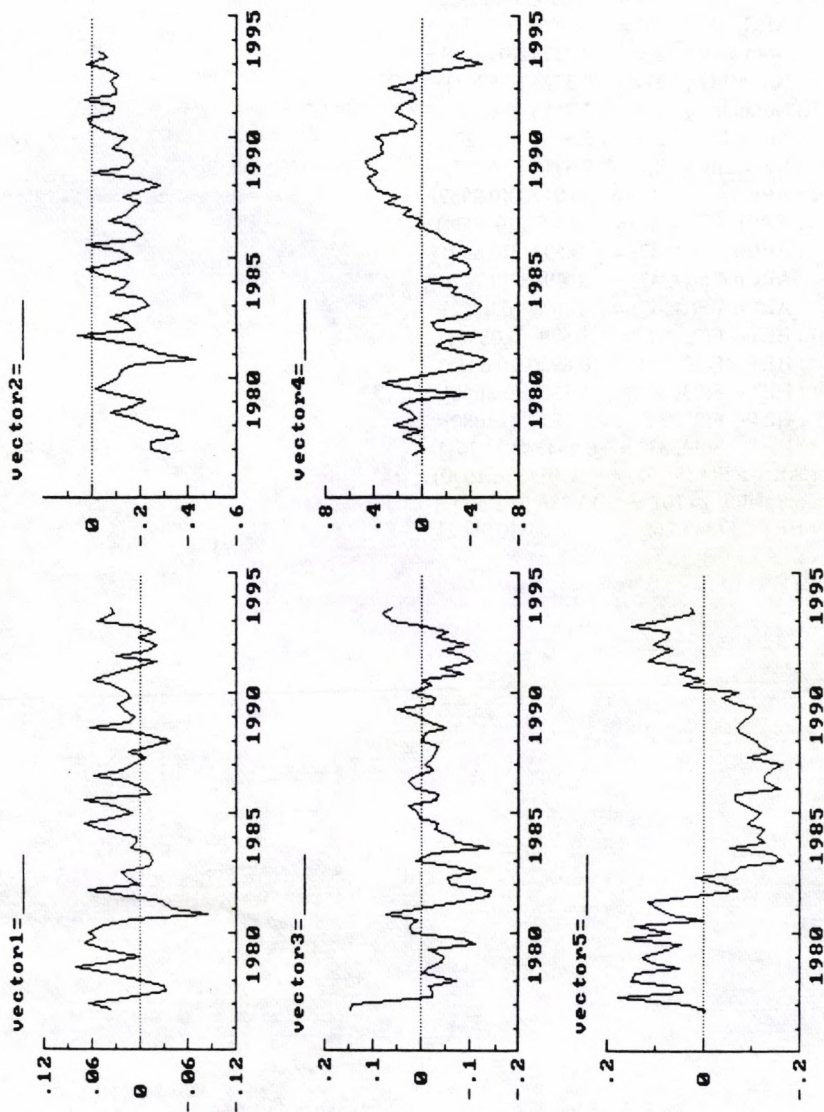
Chart 1. Real Wage Model. Cointegration Vectors
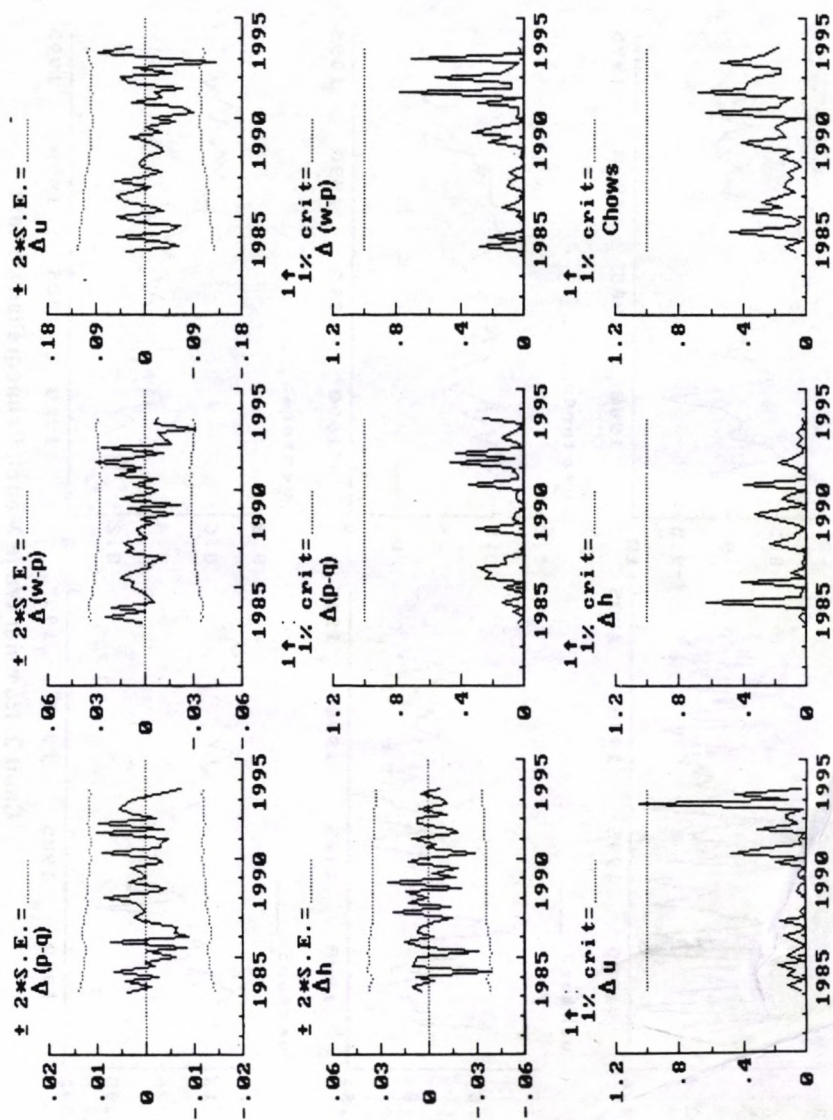
Chart 2. Real Wage Model. Recursive Tests.
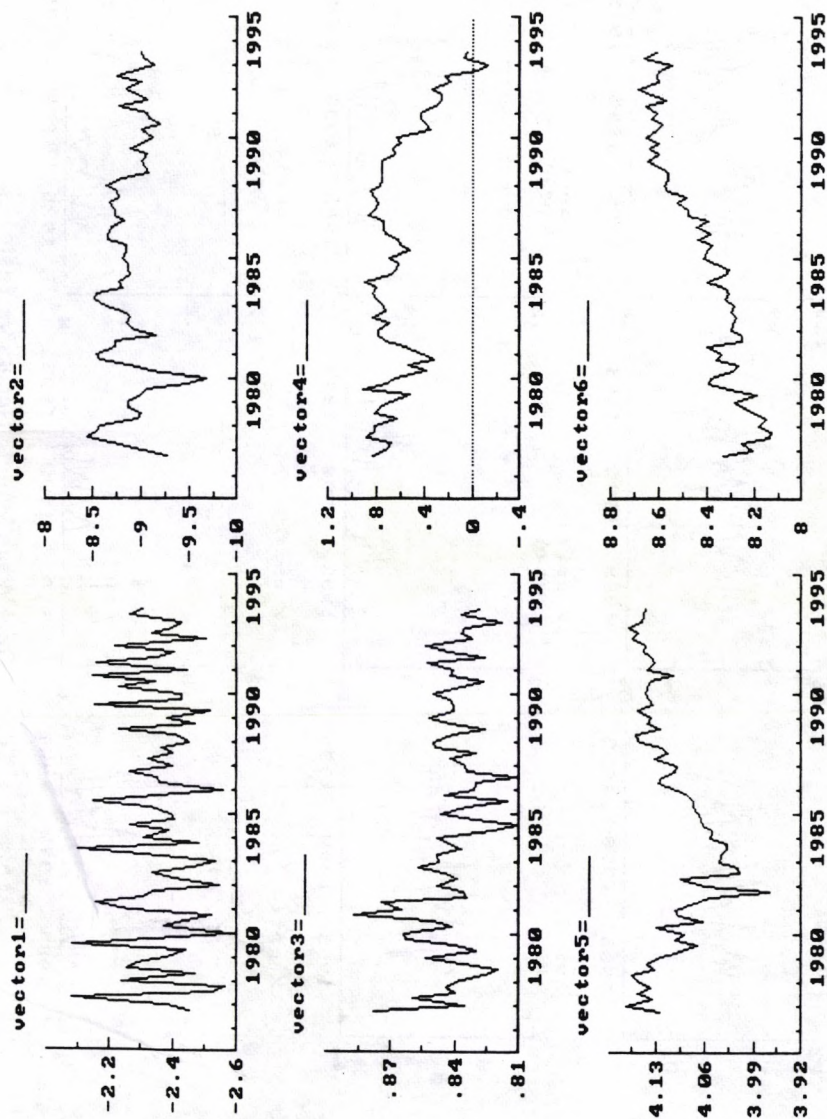
Chart 3. Nominal Wage Model. Cointegration Vectors
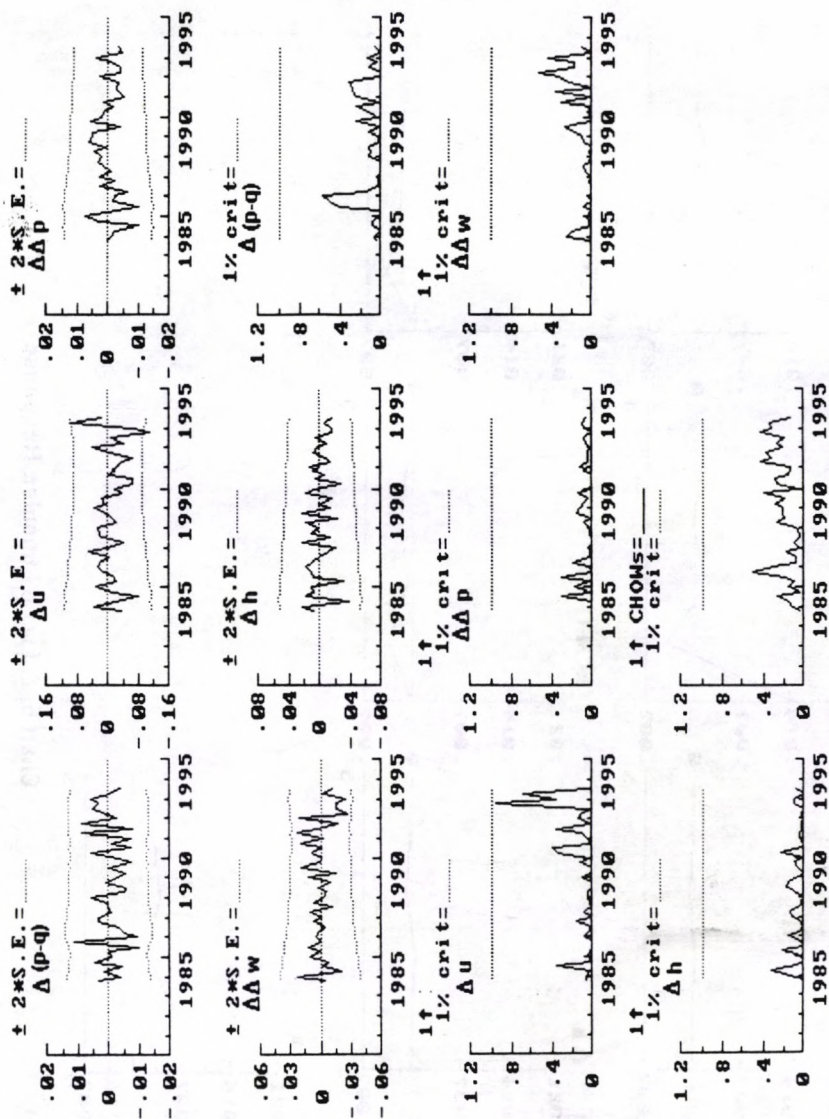
Chart 4. Nominal Wage Model. Recursive Tests
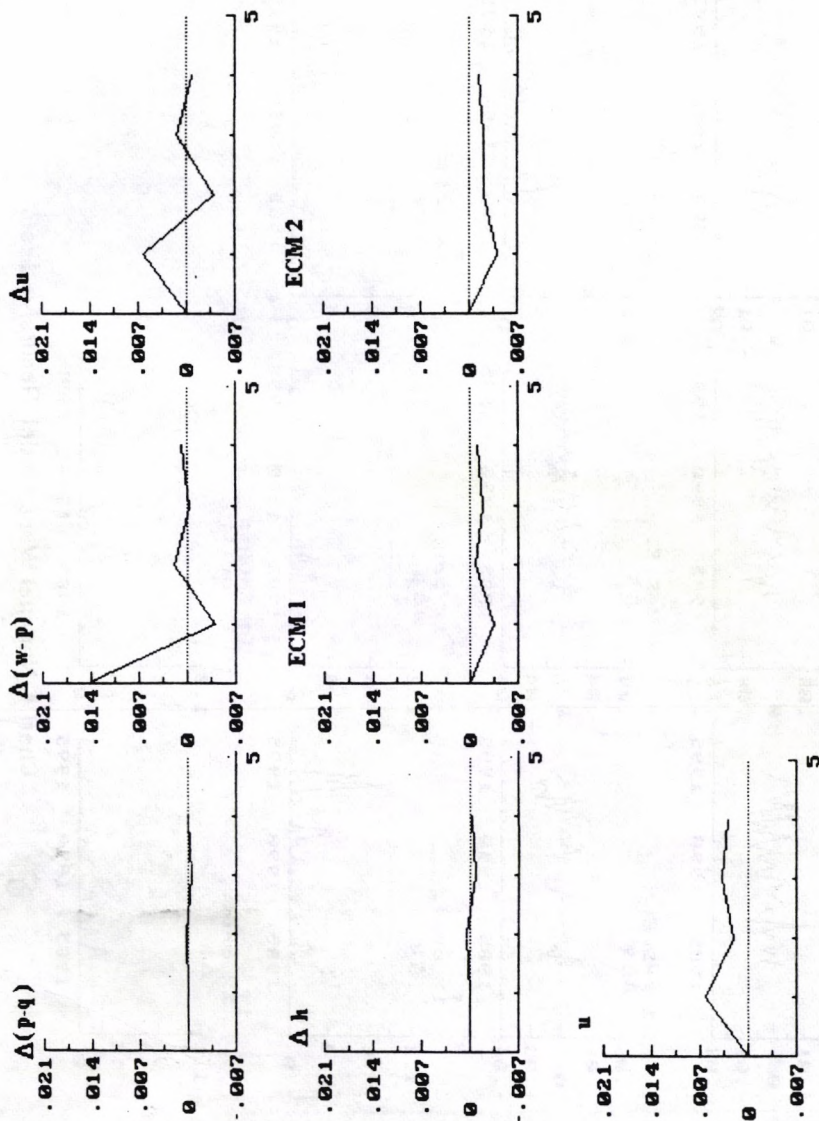
Chart 5. Δ (w-p) **Impulse Response.**
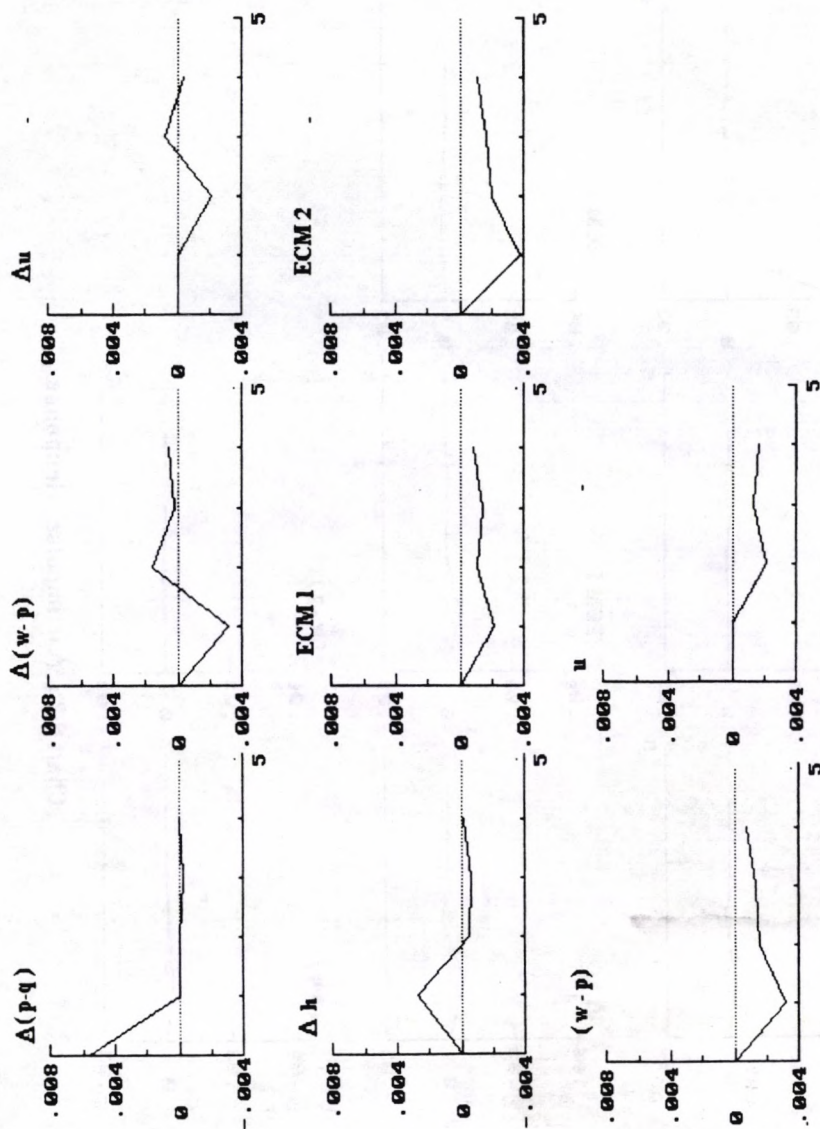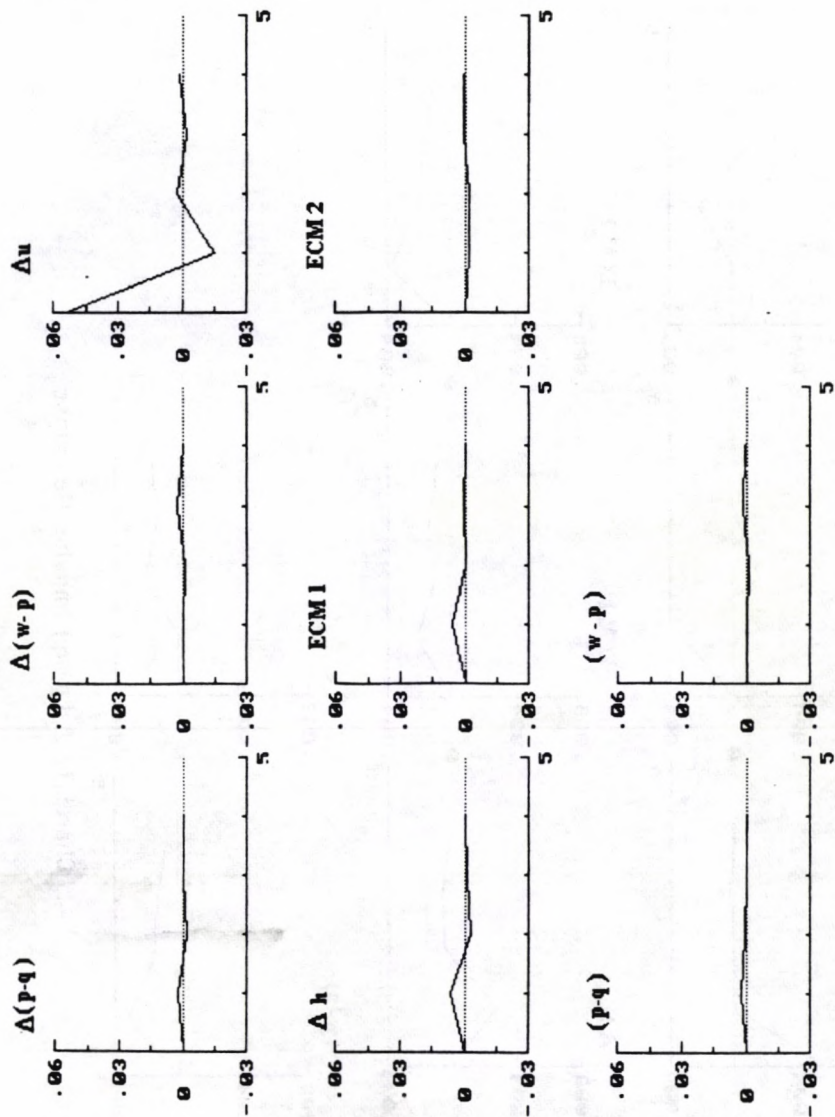
Chart 5.1. Δ (p - q) Impulse Response

Chart 5.2. Δu Impulse Response
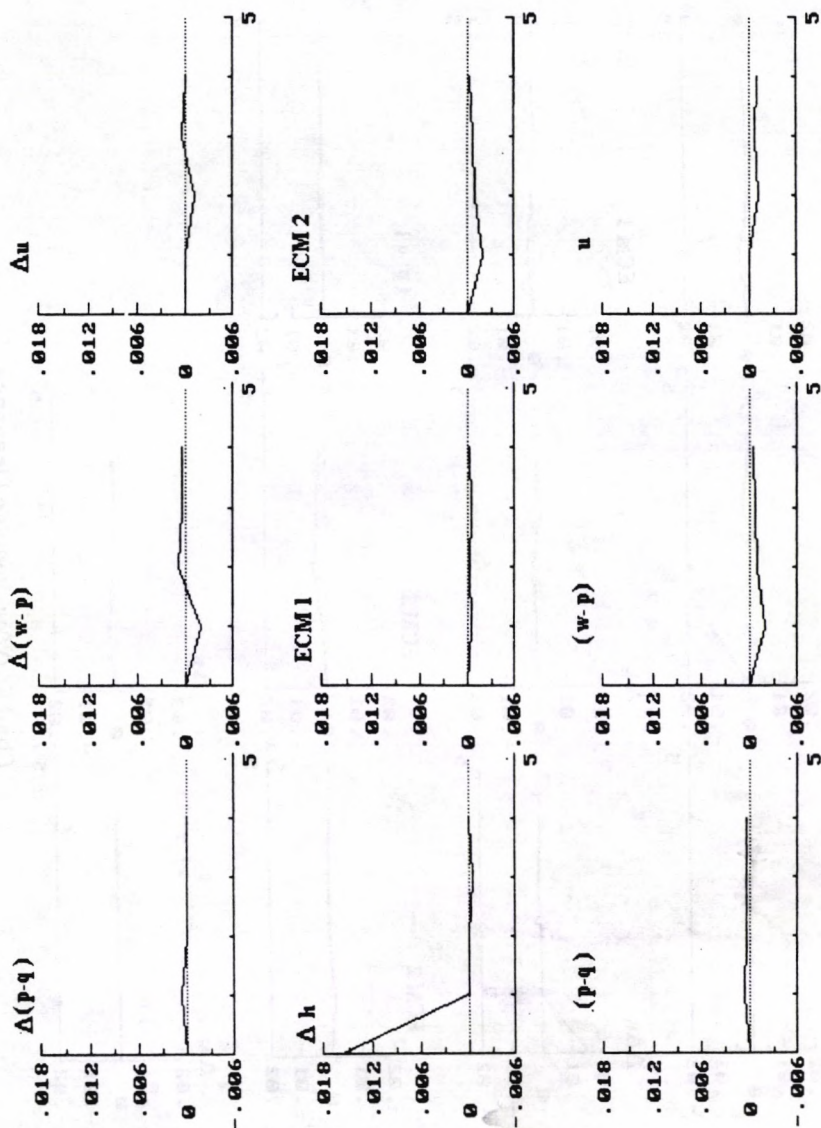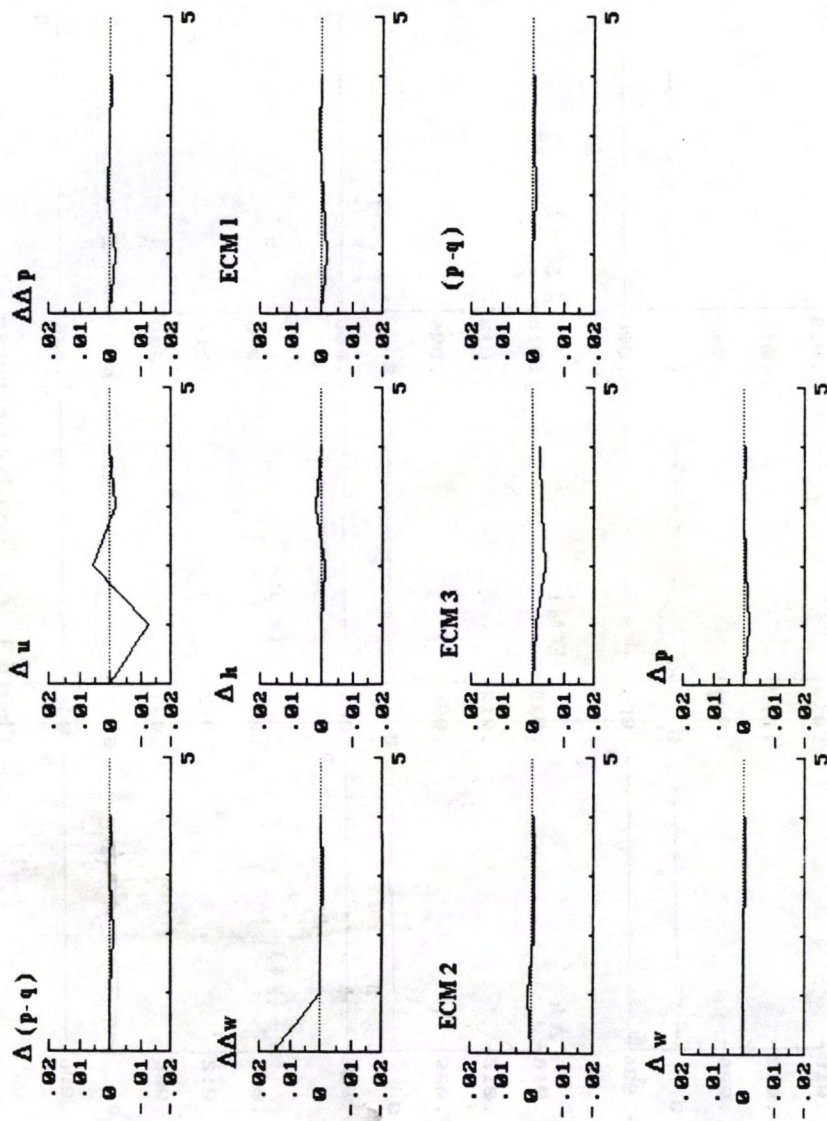
Chart 5.3. Δh Impulse Response

Chart 6: The ΔΔw Impulse Response
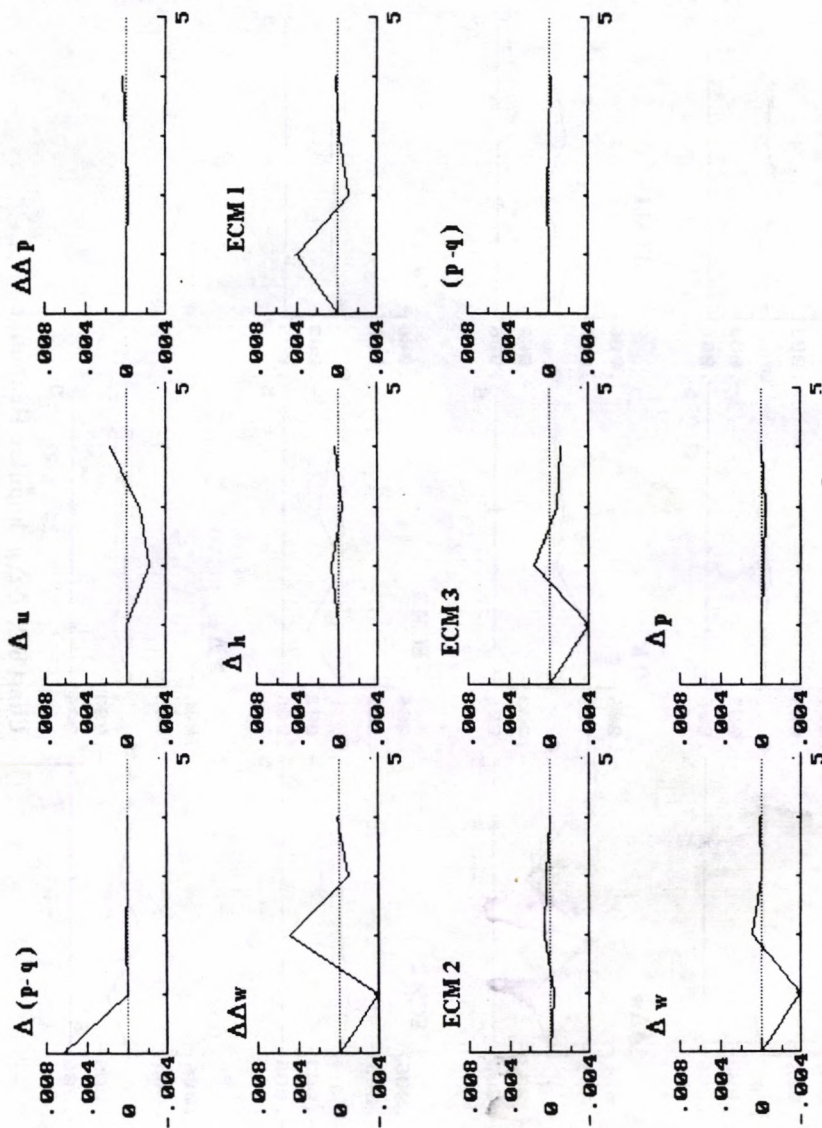
Chart 6.1. Δ[p - q] Impulse Response
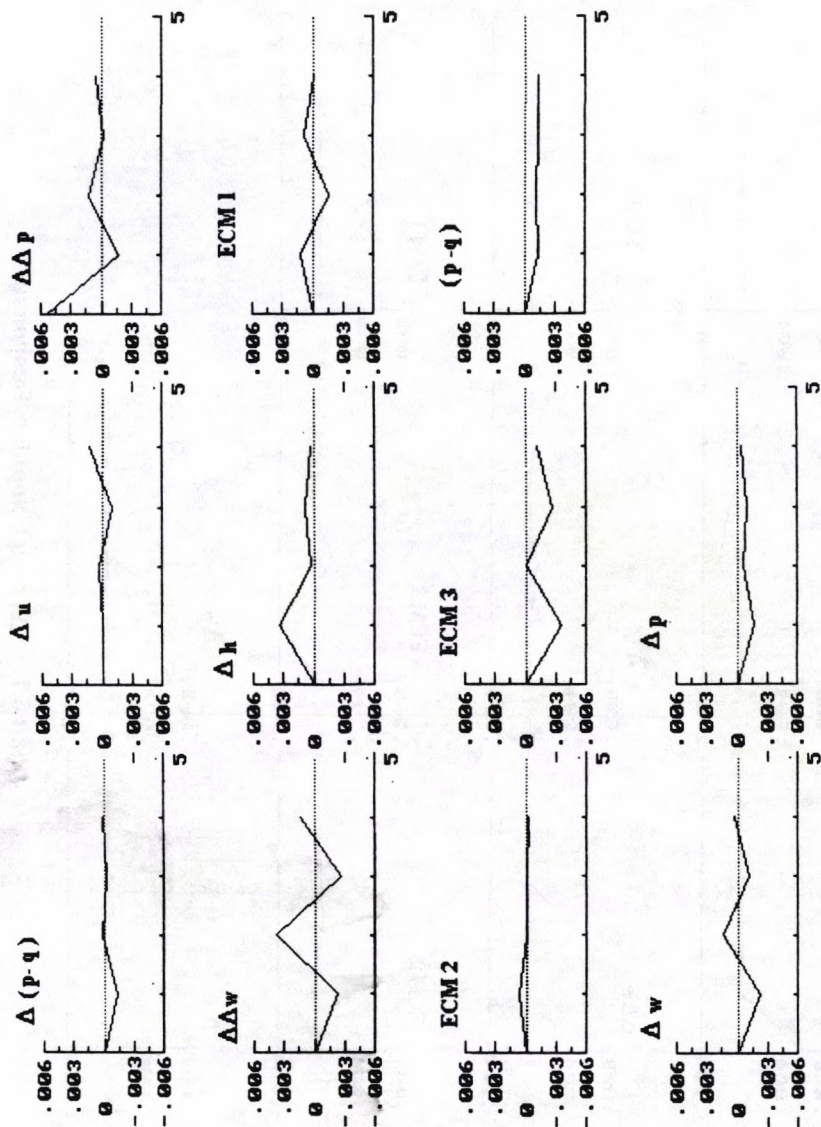
Chart 6.2. $\Delta\Delta p$'s Impulse Response

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