UNIVERSITÀ DEGLI STUDI DI NAPOLI
"PARTHENOPE"
ISTITUTO DI STUDI ECONOMICI

THE ESTIMATION OF THE NAIRU AND THE EFFECT OF PERMANENT SECTORAL EMPLOYMENT REALLOCATION. THE ITALIAN EVIDENCE

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Vincenzo Di Maro

Abstract

This paper analyses the NAIRU making use of a cointegrated VAR and of Italian labour market data. We show that a cointegrated VAR represents a *statistically adequate* (using Aris Spanos's terminology) approach to the estimation of the NAIRU. This is an effective way to overcome several problems affecting standard structural approach. In particular the paper investigates whether permanent employment shift across Italian industrial sector [measured by the Neumann and Topel (1991)'s employment-based dispersion index] has an effect on unemployment. Among the findings of this paper we emphasize that sectoral employment shifts affect unemployment both in the long- and in the short-term. Moreover, we find that the effect of sectoral reallocation on the unemployment rate occurs only after some delay.

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1. *Introduction*

This paper concerns the estimation of the NAIRU (*non accelerating inflation rate of unemployment)* models.The purpose of this paper is twofold:

1) To analyse whether permanent employment shift across Italian industrial sectors has an effect on unemployment. Most of the unemployment change could be explained by sectoral, rather than aggregate shocks with remarkable policy implications. When the sectoral employment shift is high there are some sectors in boom and others facing a slump. Demand shifts can cause at least temporary increases in unemployment if workers who lose their jobs in contracting sectors take time to search or retool for new jobs in sectors that are expanding, or if the labour market shows substantial rigidity [see, among others Lilien (1982) and Abraham and Katz (1986)]. These allocative disturbances can be an outcome of rigidity of the labour market, that is market imperfections, high adjustment costs and serious wage rigidity.

2) To understand if the sectoral employment shift affects the natural rate of unemployment (NRU). Since the concept of NAIRU refers to an equilibrium solution, it is therefore necessary to verify whether employment shift among sectors is temporary or on the contrary is lasting since in that case NRU will be affected by employment imbalances across sectors.

The Italian economy is a good example of the so-called Eurosclerosis, showing serious labour market rigidities and it is particularly suitable for this kind of analysis.

As it will be shown in this paper the sectoral shift plays a relevant part in the persistence of unemployment in the Italian Economy, even though the negative effect of mismatch on unemployment occurs only after some delay. The findings of this paper confirm partially the ones of Chiarini and Piselli (2000), one of the first papers focusing on this issue with Italian data.

The paper is divided up as follows. Section 2 reports briefly some familiar NAIRU models and discusses why these models provide unreliable estimations of the natural rate of unemployment. Section 3 deals with the employment dispersion index proposed by Neumann and Topel (1991) and used in Chiarini and Piselli (2000) in a multivariate framework. The sectoral shift variable is defined using employment data from the Italian industrial sector. Section 4 refers to the econometric patterns involved in the estimate and offers a description of the data making use of the quarterly Italian labour market data from 1984:q1 to 1997:q4. The section also discusses how the NAIRU should be considered within a cointegrated VAR. This section continues reporting the results of the cointegration analysis carried out according to Johansen's maximum likelihood procedure. In particular the cointegration space has been identified by testing a set of over-identifying restrictions, driven by the theory, on the unrestricted cointegrating vectors. Section 5 makes use of the information related to the cointegration space to specify a VECM (*vector error correction model*) including the differences in non-stationary variables. Section 6 reports the results of the impulse response analysis to make clear the difference between short and long term behaviour of the variables. Section 7 draws some concluding remarks.

2 *The NAIRU: the textbook approach and the identification problem*

There are two main approaches to estimation of the NAIRU: the structural and statistical approach. Structural methods involve the specification of relationships governing wage and price determination (or sometimes labour demand) usually based on some underlying theory of agents' behaviour. The NAIRU is then derived from these relationships by computing the dynamic steady-state level of unemployment which will necessarily be consistent with stable wage and price inflation. On the contrary, statistical or direct methods[1] postulate the existence of an equilibrium unemployment rate but make no attempt to specify the underlying behaviour of economic agents. Instead, the NAIRU is derived from its supposed effects based on observation of the relationship between unemployment and inflation. [Westaway (1997)].

This paper follows the structural approach to the estimation of the NAIRU and refers to the imperfectly competitive bargaining approach popularised by Layard,Nickell and Jackman (1991). Such a model is known as the "battle of mark-ups" model, and is referred to as the "textbook approach". Within a stylised version of the "battle of mark-ups" model unemployment must move to balance the demands of employers (the feasible real wage) and the demands of employees (the bargained real wage).

This paper takes a critical look at this structural "textbook approach". Despite the vast literature [see Gordon(1982,1997), Sawyer(1998) and the papers quoted therein] making use of the "battle of mark-ups" model to estimate the NAIRU, this approach has many aspects open to criticism. It is well known that this approach to the NAIRU makes wage equation unidentified [see, among others Manning(1993), Bean(1994) and Westaway(1997)].

Figure 1 shows the implications of a lack of identification for the wage setting schedule. The textbook approach only estimates differently sloped versions of the wage curve that are crossing the price curve at the same point. In empirical work identification is obtained by imposing arbitrary zero restrictions or including *ad-hoc* dynamics. This view of identification is strongly misleading, since the reduced form is not estimated explicitly. What seems really wrong in these models is the fact that the reduced form estimated is invalid even if, as is commonly found, it seems to be very reasonable[2]. Rather than from a statistical reduced form, these conventional models calculate the NAIRU from estimated structural forms. The empirical evidence based on these solutions are seriously misspecified and, therefore, the policy implications erroneous. Following some seminal papers of Spanos (1988, 1989a, 1989b), we argue that this kind of approach tends to impose theory-oriented restrictions with little (if any) attention to statistical properties of the model[3].

[1] See Elmeskov(1993) and OECD(1994) for an example of this type of approach.

[2] Since the signs seem to be correct and the coefficients appear to be significant using the standard asymptotic standard errors.

[3] See also Chiarini and Piselli (2001).

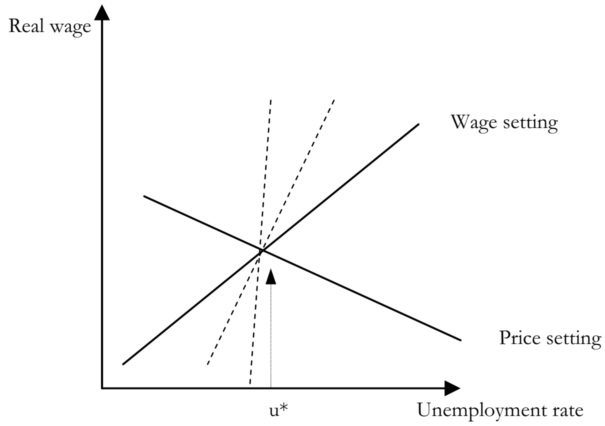

FIGURE 1: THE IMPLICATIONS OF LACK OF IDENTIFICATION FOR THE WAGE SETTING SCHEDULE

This paper aims to show how developments of cointegration within VAR methodology can represent an effective way to overcome problems of the textbook approach. Section 4, below, will focus on this subject.

3 *The employment dispersion measure*

Figure 2 shows the evolution of employment shares in Italy for 10 two-digit industries according to the ISTAT classification (see Table1) over the period 1982:1- 1999:4.

TABLE 1: 10 TWO-DIGIT INDUSTRIES OF ITALIAN INDUSTRIAL SECTOR ACCORDING TO THE ISTAT CLASSIFICATION

Tag is the same reported in figure 2	
ENERGIA:	Energy products
ALIM:	Food, beverages and tobacco
TESSILE:	Textiles, apparel, leather and footwear
LEGNO:	Wood products and furniture, paper, rubber and others
CHIMICA:	Chemical products
NON MET:	Non-metallic mineral products
METALLI:	Basic metal industries and fabricated metal industries
MECC:	Mechanical industry
TRASP:	Transportation equipment
COSTR:	Construction

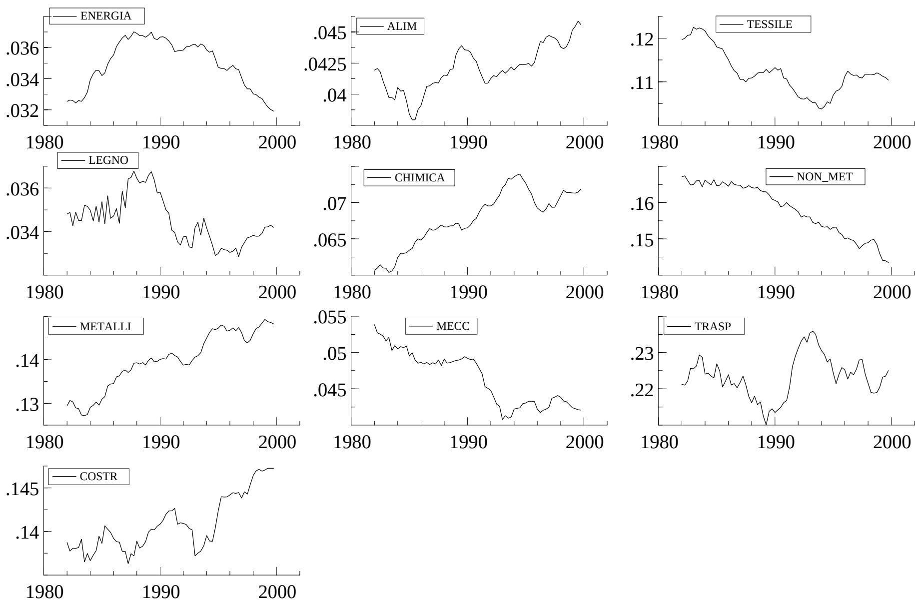

FIGURE 2: EMPLOYMENT SHARES IN 10 INDUSTRIES OF ITALIAN INDUSTRIAL SECTOR ACCORDING TO THE ISTAT CLASSIFICATION (1982:1 - 1999:4)

The figure shows distinct trends in the shares and makes clear the reallocation process undergone by the employment composition over the sample period. Sectors that suffered most of the reallocation are energy products, non-metallic mineral products, manufactured products, chemical products, construction and food, beverages and tobacco. As previously stated unemployment changes may be affected by a sectoral shock. Moreover permanent sectoral imbalances could affect the natural rate of unemployment.

An empirical testing of former statements requires a variable measuring employment dispersion. Since Lilien (1982) many authors have proposed a labour demand shift measure. Lilien's weighted standard deviation of employment across sectors was the sectoral shift measure most widely used in the '80s. However Abraham and Katz (1986) have shown that disentangling temporary and permanent effects of employment reallocation is important. They argued that the business cycle is not neutral across different markets and in some sectors the employment shift, instead of representing the consequence of a permanent change in labour demand, might be generated by cyclical effects.

Abraham and Katz's criticism has inspired the work of Neumann and Topel (1991) that show an effective way to disentangle variability of employment shares in two components: one accounting for lasting changes in labour demand across sectors (permanent effects) and another aiming to catch the changes due to local cycles or other unpredictable events (transitory effects).

In the following empirical analysis a variable measuring permanent employment shift will be used. The variable has been worked out according to Neumann and Topel's suggestions and to Chiarini e Piselli(2000) who used a similar variable relating to the period

1975:1-1993:3. Appendix A1 gives a detailed explanation of how the dispersion index was constructed.

The following Figure 3 shows the employment dispersion variable, *disp*, and the unemployment rate during the period 1984:1-1997:4.

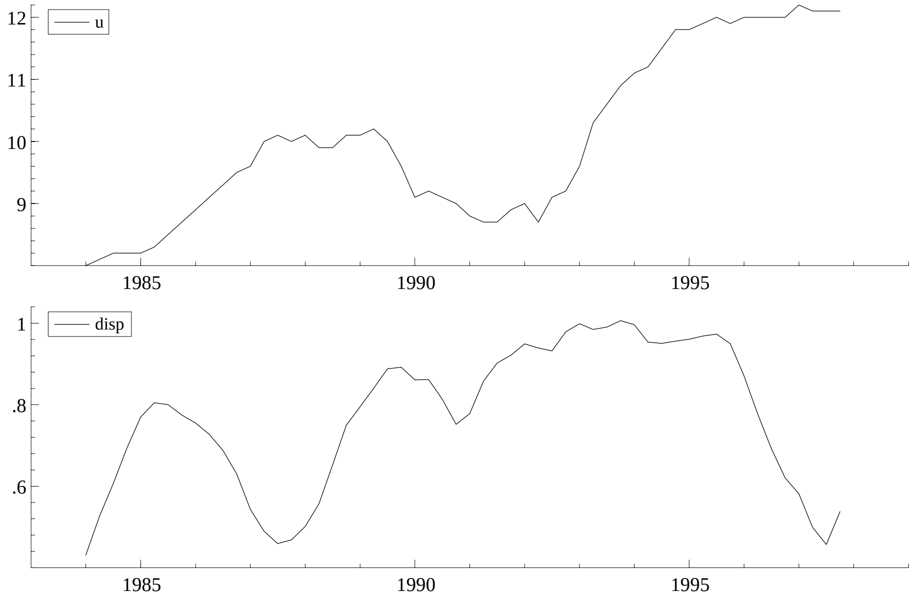

FIGURE 3: THE SECTORAL EMPLOYMENT SHIFT VARIABLE AND THE UNEMPLOYMENT RATE (1984:1-1997:4)

By the simple comparison between the graphs of the two series an interesting feature seems to transpire: when employment dispersion increases, unemployment rate tends to rise faster after some delay. This trend became clearly visible during 1985 when the employment dispersion index reached a peak and the unemployment rate began increasing markedly. Later on, during the whole period 1990-1995 the shift index was high and at the same time the unemployment rate also showed a huge rise, even if it only started to increase after 1992. Lastly, it's worth noticing that when the employment shift measure is low or, even decreasing, the unemployment rate is reasonably steady as is shown in Figure 3 focusing on the period 1987-1989 or after 1995.

4 *Cointegration analysis*

4.1 *The Data Set*

Data are quarterly seasonally adjusted over the period 1984:1:1997:4. The system includes 5 stochastic variables [wp, u, pq, pr, disp], respectively, consumption real wage, OECD standardised unemployment rate, price wedge (*consumption price minus product price*), labour productivity (*value added/labour units*) and the previously defined sectoral shift variable[4].

All the variables are non-stationary I(1) time series[5] in log form, except for unemployment rate, *u* and dispersion index, *disp*, and refer to Italian industrial sector[6].

4.2 *Criteria and diagnostic tests: lag length and descriptive power of the model*

Lag order of the VAR was chosen by joint analysis of criteria (see Table 2) and of diagnostic tests (see Table 3).

Akaike information criterion (AIC) picks out a lag order *k=3* (even though the value for *k=2* is very close to the *k=3* one), whereas the *Schwarz Bayesian criterion* (SBC) suggests a lag order of *k=1*. The adjusted LR test reject *k=6* against both *k=1* and *k=3*. These findings have to be completed by the study of diagnostic tests. In Table 5 tests relevant to a VAR(3) seem to state *white noise* residuals, except for some problems with the autocorrelation of the dispersion index equation residuals. Moreover, the diagnostic tests for VAR(3) are better than those for VAR(2) (not shown). Recursive tests (*1-step residuals +/- 2*SE* and *Break-point Chow-test*), shown in Figure 4, give evidence of a good descriptive power and a sufficient stability of the model.

TABLE 2: SELECTION CRITERIA FOR ORDER OF THE VAR

Order	LL	AIC	SBC	LR test	Adjusted LR test
6	768.9045	613.9045	465.7227	------	------
5	730.9960	600.9960	476.7145	CHSQ(25)= 75.8170[.000]	28.8105[.272]
4	683.9179	578.9179	478.5367	CHSQ(50)= 169.9733[.000]	64.5899[.080]
3	661.8414	581.8414	505.3605	CHSQ(75)= 214.1262[.000]	81.3680[.288]
2	636.6915	581.6915	529.1109	CHSQ(100)= 264.4261[.000]	100.4819[.468]
1	592.2659	562.2659	533.5855	CHSQ(125)= 353.2773[.000]	134.2454[.270]
0	313.7126	308.7126	303.9325	CHSQ(150)= 910.3839[.000]	345.9459[.000]

[4] Appendix A2 reports a more detailed account on variables.

[5] Appendix A3 reports results of unit root tests.

[6] In particular, the variables refer to the aggregate called, according to the ISTAT classification, "industry in strict sense", which excludes energy products and construction.

TABLE 3: DIAGNOSTIC TEST FOR VAR(3) (1984Q1-1997Q4)

lwp	:	AR 1- 4 F(4, 30)	=	1.7954 [0.1558]
u	:	AR 1- 4 F(4, 30)	=	1.0897 [0.3794]
lpq	:	AR 1- 4 F(4, 30)	=	1.4927 [0.2293]
lpr	:	AR 1- 4 F(4, 30)	=	0.20391 [0.9342]
disp	:	AR 1- 4 F(4, 30)	=	6.5874 [0.0006] **
lwp	:	Normality Chi^2(2)	=	1.8965 [0.3874]
u	:	Normality Chi^2(2)	=	0.12023 [0.9417]
lpq	:	Normality Chi^2(2)	=	0.80306 [0.6693]
lpr	:	Normality Chi^2(2)	=	2.3027 [0.3162]
disp	:	Normality Chi^2(2)	=	0.71756 [0.6985]
lwp	:	ARCH 4 F(4, 26)	=	0.056634 [0.9937]
u	:	ARCH 4 F(4, 26)	=	0.045778 [0.9958]
lpq	:	ARCH 4 F(4, 26)	=	0.43503 [0.7821]
lpr	:	ARCH 4 F(4, 26)	=	0.46969 [0.7574]
disp	:	ARCH 4 F(4, 26)	=	0.28937 [0.8822]
lwp	:	HET F(30, 3)	=	0.10418 [0.9999]
u	:	HET F(30, 3)	=	0.22266 [0.9898]
lpq	:	HET F(30, 3)	=	0.075625 [1.0000]
lpr	:	HET F(30, 3)	=	0.072627 [1.0000]
disp	:	HET F(30, 3)	=	0.13719 [0.9992]
Vector AR 1-4 F(100, 53)			=	1.3596 [0.1098]
Vector normality Chi^2(10)			=	5.2515 [0.8738]
Vector HET Chi^2(450)			=	459.02 [0.3742]

Testing for vector error autocorrelation from lags 4 to 4
Chi^2(25) = 51.185 [0.0015] ** F-form(25,94) = 1.5668 [0.0637]
lwp for Error Autocorrelation from lags 4 to 4
Chi^2(1) = 3.151 [0.0759] F-form(1,33) = 2.0859 [0.1581]
u for Error Autocorrelation from lags 4 to 4
Chi^2(1) = 4.5806 [0.0323] * F-form(1,33) = 3.1219 [0.0865]
lpq for Error Autocorrelation from lags 4 to 4
Chi^2(1) = 0.065803 [0.7975] F-form(1,33) = 0.041023 [0.8407]
lpr for Error Autocorrelation from lags 4 to 4
Chi^2(1) = 0.64187 [0.4230] F-form(1,33) = 0.40455 [0.5291]
disp for Error Autocorrelation from lags 4 to 4
Chi^2(1) = 0.26398 [0.6074] F-form(1,33) = 0.16519 [0.6870]

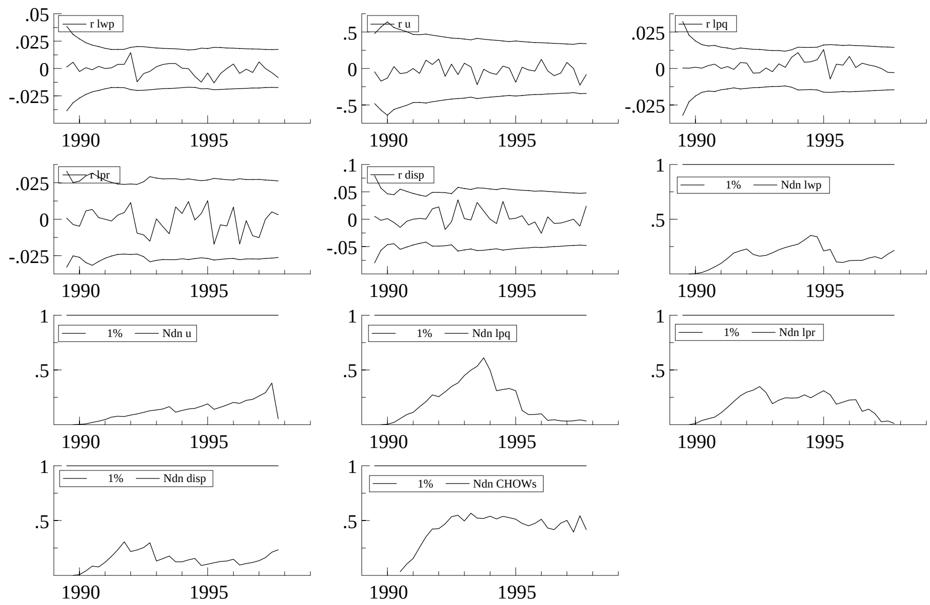

FIGURE 4: 1-STEP RESIDUAL +/- 2*SE AND BREAK-POINT CHOW TEST OF VAR(3)

In order to achieve *white noise* residuals it has been necessary to augment the VAR with three shift dummy variables: *d92q12* (1992:1-1992:2=1), *d91q1* (1991:1=1) and *d98q3* (1998:3=1). The last two variables (*d91q1* and *d98q3*) account for an outlier, respectively in the price wedge's and in the dispersion index's residuals whereas *d92q12* dummy variable is addressed to catch the remarkable shift in unemployment rate during 1992. It's worth noticing that data on 1992 show some problems both from a statistical and an economic point of view. Eurostat changed its employment survey methodology exactly in 1992 and, besides, it was one of the most troubled periods of the Italian economy in the last decade[7].

4.3 The statistical model

Criteria and diagnostic tests of the previous sub-section let one conclude that a VAR(3) is a good approximation of the data. Dealing with I(1) variables it is very useful to reformulate the model by applying the so-called "cointegrating transformation" in order to have a model in the error correction form:

(1) $$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \Gamma_2 \Delta x_{t-2} + \Pi x_{t-3} + \Psi D_t + \mu_t$$

[7] In the summer of 1992 the Bank of Italy nearly used all its foreign currency reserves to protect Italian currency from a speculative attack. In September of 1992 the Bank of Italy gave up protecting the currency and Italy shifted to a floating exchange rate system [Signorini and Visco (1997)].

where $\mathbf{\Gamma_1 = -(I - A_1)}$, $\mathbf{\Gamma_2 = -(I - A_1 - A_2)}$, $\mathbf{A}_i$ is a *(nxn)* matrix and the vector $\boldsymbol{D_t}$ is a set of conditioning non-stochastic variables such as intervention or shift dummies and others that are weakly exogenous.

It follows from the *Granger representation theorem*[8] that if the rank of matrix $\mathbf{\Pi}$, whose dimension is *(nxn)*, is equal to $r<n$, there exists a representation of $\mathbf{\Pi}$ such that $\mathbf{\Pi} = \boldsymbol{\alpha\beta}$, where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are both *nxr* matrices. [Johansen(95)]

Coefficients of the matrix $\boldsymbol{\alpha}$ represent the speed of adjustment to disequilibrium, while $\boldsymbol{\beta}$, also called the "cointegrating matrix", is the matrix of long-run coefficients such that the term $\boldsymbol{\beta X_{t-3}}$ embedded in (1) stands for up to *(n-1)* cointegration relationships. [see, among others Harris (1995), Hendry (1995), Favero (2001), Doornik and Hendry (1997) and Hamilton (1994)]

The problem of detecting the number and the shape of cointegrating vectors can be solved by applying Johansen's *reduced rank regression* [Johansen(95)], a widely used technique for selecting rank of $\mathbf{\Pi}$ (and so number of cointegrating vectors) and providing ML estimates of $\boldsymbol{\alpha}$ e $\boldsymbol{\beta}$ matrices. This last technique will be applied in sub-section 4.5 where cointegration analysis will be carried out.

4.4 The NAIRU within a cointegrated VAR

In section 2 we reviewed some of the problems affecting the standard structural approach to estimation of the NAIRU. In particular the textbook approach does not give the right attention to the problem of lack of identification of the wage equation and to statistical properties of the model [see Westaway (1997), Manning (1993), Bean (1994) and Chiarini and Piselli (2001)].

Following Spanos (1988, 1989a, 1989b) a "battle of mark-ups" model cannot be considered statistically adequate. In fact, the textbook approach considers the probabilistic structure of the data only as an afterthought and specifies the statistical model simply "*...attaching an error term to theoretical relationships"*. Spanos argued that the structure of data should be considered at the statistical model specification stage and, therefore, within the textbook approach there is no room for implementing *statistical adequacy*. Chiarini and Piselli (2001), following Spanos, have shown that these conventional models define a deducted form, rather than a statistical reduced form, calculating the NAIRU from estimated structural forms. Even if estimates of structural equations seem to be reasonable in terms of goodness of fit, t-statistic of regressors, correlations of residuals and expected signs of coefficients, their implicit reduced form is often not valid as it does not show "*white noise*" properties[9], and it's well known that properties of the reduced form are crucial for consistency of the structural form.

[8] See Engle and Granger (1987) and Johansen (1995).

[9] Chiarini and Piselli (2001) give an example of that analysing the empirical behaviour of wages and price determination in Italy over the period 1975:1-1993:4.

A reduced form multivariate framework, such as a cointegrated VAR model, can be an effective way to specify a model with the *statistical adequacy* property. The probabilistic structure of data becomes a key point within a cointegrated VAR and a statistical model will be specified only if probability assumptions are not rejected by data [see Hendry (1995) and Favero (2001)]. Data assume a role that it is not secondary compared to that of the theory. The statistical model, within this approach, is a convenient summary of the sample information that, at the same time, considers the theoretical model [Spanos (1989a)].

The NAIRU models, such as the "battle of mark-ups, should be integrated into a framework for modelling multivariate relationships. In this context, the empirical model needs two distinct characteristics: it will need to capture the equilibrium relationships among the variables and it needs to capture the adjustment to equilibrium following a shock. Thus, we should examine conditions for stability and stationarity of the model and show how the presence of non-stationary variables is potentially useful in providing information about the long-run relationships.

Using a cointegrated VAR for estimating the cointegration space, that is the space which stationary relations span[10], a sensible set of theory-grounded restrictions should be able to detect a firm and an union real wage equation. In the "battle of mark-ups" model unemployment moves to balance the bargained (union) and feasible (firm) real wage.

The relevant question, at this stage, is the following: after its movements to equilibrate firm and union real wage will the unemployment rate come back every time to the same level, that is to the natural level? or, in other words, is the NAIRU a single fixed point in the cointegration space?

The answer to the latter question is not straightforward. However, relying on the concept of the stationary (cointegrated) space it is possible to use a helpful geometrical explanation. For instance, Chiarini and Piselli (2001) argue that the NAIRU estimates using conventional reduced form models (or the Phillips curve), are misleading since the natural rate of unemployment cannot be considered as a fixed point, as in the traditional NAIRU models, calling for the *ceteris paribus* assumption for all the data involved. In a simplified model of only 3 variables and 2 cointegrating vectors, the NAIRU has to be considered as a straight line, defined by the orthogonality of the two stationary relationships.

From a stricter geometrical point of view the NAIRU is defined as the attractor set, that is the orthogonal complement of β, written as $\beta_\perp$ (see Figure 5). β is the cointegrating matrix, whose r columns are the r cointegrating vectors and the orthogonal complement of β is defined as the [$nx(n-r)$] matrix $\beta_\perp$ such that $\beta'\beta_\perp = 0$. In actual models, with n variables and r cointegrating vectors, the NAIRU has to be considered as a hyperplane with dimension n-r. Within a cointegrated VAR, a shock to one variable implies a shock to all the variables in the long run and so the *ceteris paribus* assumption is not allowed. The only way of analysing the economic relationships in this context remains the impulse response analysis, that will be carried out in section 6. If the NAIRU is analysed in a multivariate context, eventually, specifying a fixed value for it makes no sense.

[10] Cointegration space is the set of all possible linear combinations among unrestricted cointegrating vectors. The only way to identify unique relationships is to impose some restrictions on unrestricted vectors.

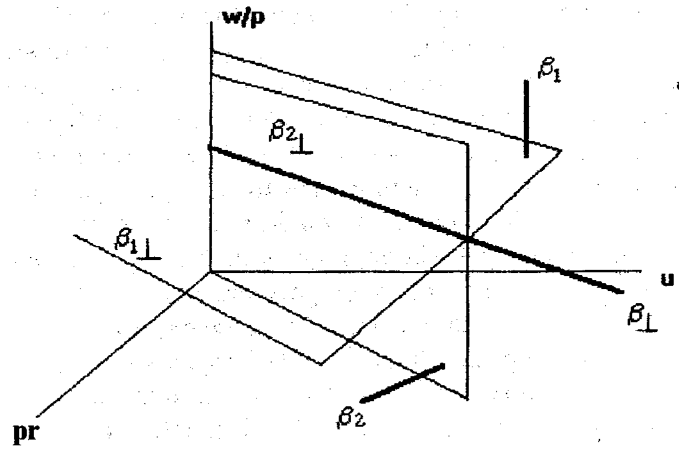

FIGURE 5: THE NAIRU IN THE COINTEGRATION SPACE

4.5 Johansen's reduced rank regression

Applying Johansen's *reduced rank regression* and testing for cointegration provide the result reported in Table 4 where Max and Trace are, respectively, *Max eigenvalue* and *Trace statistics* in the standard version and adjusted for degrees of freedom (Max^ and Trace^).

TABLE 4: COINTEGRATION ANALYSIS (1984Q1-1997Q4)

eigenvalues

0.510605	0.397133	0.270864	0.223992	0.165959

Ho:rank=p	Max	using T-nm	95%	Trace	using T-nm	95%
p == 0	37.87**	27.15	34.4	104.5**	74.92	76.1
p <= 1	26.82	19.23	28.1	66.62**	47.77	53.1
p <= 2	16.74	12	22.0	39.8**	28.54	34.9
p <= 3	13.44	9.637	15.7	23.06**	16.53	20.0
p <= 4	9.618	6.896	9.2	9.618**	6.896	9.2

standardized \beta' eigenvectors

lwp	u	lpq	lpr	disp	Constant	Trend
1.0000	0.021531	0.48839	-1.0251	-0.0069453	0.60454	0.0017509
121.87	1.0000	128.55	-9.2104	-10.910	597.36	-0,67599
0.69914	0.015608	1.0000	-0.39300	0.069711	1.8739	-0.0038503
1.1047	-0.026567	-0.58767	1.0000	-0.054941	9.9006	-0.0036223
12.020	0.62023	-18.534	-1.6242	1.0000	43.305	0.035317

standardized \alpha coefficients

lwp	0.12411	-0.0014037	-0.27240	-0.028016	-0.0039689
u	2.8183	-0.0067621	4.6708	0.32737	-0.15964
lpq	0.29730	-0.0018186	0.10495	0.075423	0.0027409
lpr	0.73246	-0.00092115	0.073468	-0.11650	0.0014080
disp	0.41849	0.0058612	-0.55182	0.33119	-0.0052317

Max eigenvalue suggests the presence of one cointegrating vector while *Trace statistics* selects five. One should not be surprised by this diverging result since the critical values of tests, due to conditioning set $\boldsymbol{D_t}$ affecting the asymptotic distribution of these ones, are considered to be indicative.

For this reason it is necessary, as Juselius (1995) pointed out, to gather further information analysing the magnitude of eigenvalues and roots of the companion matrix[11] of VAR model as well as graphs of the eigenvectors. To begin with, the first eigenvalue in Table 4 is high enough (0,5106) while second one is already too low (0,3971).

The graphs of the Figure 6 depict the first two eigenvectors and show that the second vector cannot be considered stationary. Notice also, the wide range in which vector floats (+5, -2,5).

All the roots of the companion matrix (they are n*k=15) are inside the unit cirde and the four largest ones have as modulus a value very close to 1 (0,9868 0,9868 0,9097 0,9097) while the fifth (0,8073) seems far enough from 1. As a final test it's worth analysing results of the ADF (*augmented Dickey-Fuller*) [Dickey and Fuller (1979)] test carried out on the first two eigenvectors (see Table 5). ADF states on one hand stationarity of the first vector but on the other it provides little information about second one.

All previous information supports the choice of one cointegratingvector only and *n-r=4* common stochastic trends.

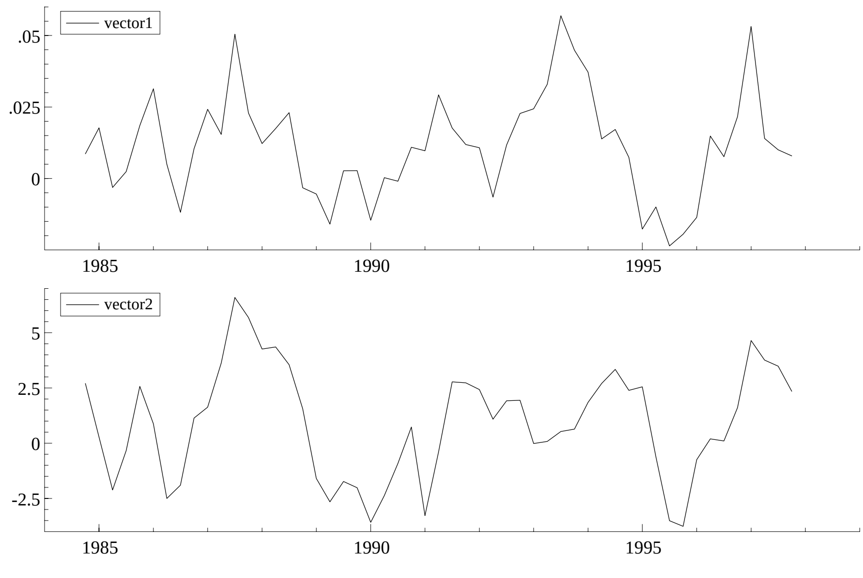

FIGURE 6: FIRST TWO EIGENVECTORS

[11] The companion matrix of the VAR model refers to the matrix A of equation (1). It is interesting to analyse the roots of the companion matrix as these provide useful information about the stability of the system. If all roots have modulus inside the unit circle then all the variables are I(1). When any of the modulus is outside the unit circle, some of the variables could be I(2).

TABLE 5: ADF TEST ON THE FIRST 2 EIGENVECTORS (1984Q1-1997Q4)

```
Critical values: 5%=-2.921 1%=-3.568; Constant included

                t-adf
vector1 ADF(3) -3.3472*
vector1 ADF(2) -2.9053
vector1 ADF(1) -3.1532*
vector1  DF    -3.4576*
vector2 ADF(3) -3.2089*
vector2 ADF(2) -2.5627
vector2 ADF(1) -3.5327*
vector2  DF    -2.6347
```

4.6 Identification of cointegrating vector

Dealing with only one cointegrating vector it suffices to normalise for one of the variables to achieve identification. If the vector is normalised for the real wage (*wp=1*), the just-identified vector is represented by the equation (2), where t-statistics are in brackets.

(2) $$wp = -0,02\underset{(4,87)}{u} - 0,48\underset{(2,31)}{pq} + 1,025\underset{(5,39)}{pr} + 0,0069\underset{(0,33)}{disp} - 0,60\underset{(0,72)}{INT} - 0,00179\underset{(1,38)}{trend}$$

However, besides this just-identified relationship, we have attempted to test for cointegration several theoretically-motivated structures, such as for instance cointegration between unemployment and the dispersion index [Chi^2(3)=16,005 (0,0011)**] as well as cointegration between the real wage and the dispersion index [Chi^2(3)=14,086 (0,0028) **] which has been rejected by the appropriate LR test. Imposing the over-identifying restrictions (*wp=pr*) it was identified the following restricted cointegrating vector that spans the cointegrating space as LR test confirms [Chi^2(1)=0,0057668 (0,9395)].

(3) $$wp = pr - 0,0211196\underset{(6,59)}{u} - 0,49360\underset{(2,43)}{pq} + 0,00777\underset{(0,42)}{disp} - 0,71541\underset{(12,05)}{INT} - 0,00159\underset{(1,79)}{trend}$$

As a consequence in the long run unemployment cointegrates with real wage, productivity, price wedge and dispersion index. Employment dispersion index, *disp*, enters in cointegrating vector with a low coefficient (0,00777), that is also not significant (t-ratio 0,42), even though the sign is correct[12]. Actually it does not make much sense interpreting coefficients of cointegrating vectors as structural parameters[13] and as regards this subject

[12] The positive sign of dispersion index is consistent with Chiarini and Piselli (2000)'s findings. They found, however, 2 cointegrating vectors for the period 1975:1-1993:1 making use of a similar data set.

[13] If variables are in log form one could consider coefficients of vectors as elasticity. Lutkepohl(95) has shown that this is correct only if one supposes that when a variables moves, the others hold constant. The

Lutkepohl (1993, 1995) showed why it is wrong to consider coefficients as elasticity. In any case, a low coefficient of dispersion index could also mean that mismatch, brought about by sectoral reallocation, does not affect real wage in the long run.

It is now clear that to understand the links among variables you have to consider not only the long run properties of the system but also the short run structure of the model. This last subject will be fully covered in the next section.

5 Modelling the short run

5.1 From long-run findings to short-run structure

The stationary relationship discussed in the previous section represents an error correction term. It is now possible to reformulate the VAR model using first differences of variables and including explicitly the error correction term, *ECM*. Hence the VECM including long and short run information becomes:

(4) $$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \Gamma_2 \Delta x_{t-2} + \alpha ECM_{t-3} + \Psi D_t + \mu_t$$

where $\Delta x_t = [\Delta wp, \Delta u, \Delta pq, \Delta pr, \Delta disp]$, *ECM* represents error correction mechanism and $\boldsymbol{D}_t$ includes an intercept and three dummy variables (*d92q12, d91q1, d97q2*).

From the OLS estimate (results are not shown but can be provided upon request) of model (4), it emerges that there is some correlation among residuals of different equations. In particular simultaneity emerges between the equations of productivity and real wage and also between the equations of productivity and price wedge. It is known that the OLS estimate is not able to deal with these kind of problems, then the FIML estimate, or *full information maximum likelihood*, considering the whole information from data (simultaneous links too), is more suitable in these cases. The FIML procedure is used to estimate a parsimonious version of the VECM that is achieved dropping insignificant repressors out.

The simultaneous structural model is presented in Table 6, where there are also multivariate diagnostic tests suggesting that the model has *white noise* and normally distributed errors, while univariate autocorrelation test for unemployment equation's residuals shows some problems [AR 1-4 F(4,32)=3,303 (0,0225)*].

Moreover LR test of over-identifying restrictions indicates that the structural model encompasses parsimoniously the VECM [Chi^2(35)=16,4379 (0,9968)]. Recursive tests, *1-step residuals* and the *Break-point Chow test* (not shown), eventually claim a good stability of the model.

latter is clearly an unreal assumption within a simultaneous system whereas, citing Lutkepohl's words, "*..economists would like to know what actually happens in the real world.*"

T-values in brackets

$$\Delta wp = \underset{(1,60)}{-0,0086458\Delta u_{t-2}} \underset{(2,21)}{-0,24901\Delta pq_{t-2}} \underset{(0,75)}{-0,059630\Delta pr_{t-1}} \underset{(4,27)}{-0,031332\Delta disp_{t-1}} +$$
$$+\underset{(2,02)}{0,0028189 INT} \underset{(1,01)}{-0,071483 ECM} + \underset{(2,70)}{0,021622 d91q1}$$

$$\Delta u = \underset{(1,05)}{3,7808\Delta wp_{t-2}} \underset{(3,39)}{-0,45376\Delta u_{t-1}} \underset{(1,16)}{-1,0556\Delta disp_{t-1}} + \underset{(1,41)}{1,2343\Delta disp_{t-2}} +$$
$$+\underset{(1,82)}{2,6540 ECM} \underset{(1,63)}{-0,22290 d92q12}$$

$$\Delta pq = \underset{(2,91)}{0,38914\Delta wp_{t-1}} \underset{(1,44)}{-0,19838\Delta wp_{t-2}} \underset{(1,49)}{-0,0083045\Delta u_{t-2}} \underset{(4,21)}{-0,42248\Delta pq_{t-2}} +$$
$$+\underset{(3,41)}{0,22955 ECM_{t-1}} + \underset{(3,30)}{0,0043650 INT} \underset{(3,48)}{-0,026606 d91q1} \underset{(2,20)}{-0,017353 d97q2}$$

$$\Delta pr = \underset{(1,22)}{0,24997\Delta wp_{t-1}} \underset{(0,81)}{-0,16390\Delta wp_{t-2}} \underset{(4,44)}{-0,57773\Delta pr_{t-1}} + \underset{(2,10)}{0,072085\Delta disp_{t-2}} +$$
$$+\underset{(4,55)}{0,51557 ECM_{t-1}} + \underset{(2,02)}{0,0041062 INT}$$

$$\Delta disp = \underset{(2,51)}{-0,046180\Delta u_{t-2}} \underset{(1,57)}{-0,51108\Delta pq_{t-1}} + \underset{(3,77)}{0,94075\Delta pr_{t-2}} + \underset{(8,48)}{1,0006\Delta disp_{t-1}} +$$
$$\underset{(3,33)}{-0,39071\Delta disp_{t-2}} + \underset{(3,21)}{0,71935 ECM} + \underset{(2,72)}{0,072805 d91q1} \underset{(2,37)}{-0,064186 d97q2}$$

Vector diagnostic vector test

Vector AR 1-4 F(100,102) =	1.0609 [0.3833]
Vector normality Chi^2(10)=	12.516 [0.2520]
Vector Xi^2 F(375,103) =	0.58946 [0.9998]

5.2 Main features of the estimated model

The structural model's findings offer various hints to be discussed. In particular it is worth discussing the following ones:

a) Lagged unemployment changes do enter significantly the unemployment equation. This seems to give evidence of the presence of unemployment persistence, and so of the *hysteresis* phenomenon. [see, among others Blanchard and Summers (1986), Lindbeck and Snower (1987) and Blanchard (1991)].

b) There is the possibility of a wage gap (*wp-pr*) in the short-run as productivity changes have no significant effects on real wage. Therefore if productivity growth falls, real wage is expected to remain unaffected, determining a wage gap, as Italy experimented[14] in the 80s.

[14] In the 80s the "*scala mobile*", a system of wage indexation, was revised. Income policy, which was able to revise wage indexation, started in 1983 with a general agreement that brought about a 15% decrease of average rate of wage indexation. Then, in first semester of 1984 government started linking wages' rise to estimated inflation target. Later wage indexation mechanism was radically revised and average indexation rate fell to about 50%. [Signorini and Visco (1997)].

c) Lagged changes of the employment dispersion index enter the unemployment equation. In particular *t-1* term has a negative sign while

t-2 has a positive sign. In other words when changes in the employment dispersion of the previous period tends to rise, the actual change of unemployment should fall. This result is quite peculiar and a possible explanation will be provided below, referring to impulse response analysis.

d) Lagged unemployment change enters the real wage equation with the expected negative sign but the coefficient is very low (-0,0086) as well as it not showing a satisfactory t-value (1,60). This result does not seem very surprising considering that the Italian labour market is highly unionised, with a centralized bargaining process where the "outsiders" have no role to play.

e) Real wage cointegrates in the long-run with unemployment, price wedge and employment dispersion measure as well as productivity. Moreover, it was tested if the wage gap (*wp-pr*) exists in the cointegration space, that is if it is stationary, with the data rejecting this hypothesis [Chi^2(3) = 16.603 [0.0009] **]. These two outcomes indicate a high degree of real wage rigidity that is confirmed by the short-run structure: since the cointegration relation, *ECM*, enters the wage equation with a coefficient (-0,071) that is non-significant (*t-ratio* 1,01).

f) Price wedge has a negative effect on the real wage (the coefficient is -0,24). Hence any expansionary macro-policy that makes price wedge rise reduces the real wage and, therefore, the unemployment.

Before discussing the impulse response analysis is, however, necessary to discuss some features analysis of this paper does not show.

The analysis is carried out on the aggregate Italian labour market neglecting some important features of this market. For instance, the Italian labour market presents a remarkable dualism with the Northern labour market close to full employment whereas the unemployment rate in the southern regions is about 25%[15]. In fact in many studies considering a dualistic labour market [see Bodo and Sestito (1991) and Chiarini and Piselli (1997)], the southern unemployment rate is found not significant both in the short and long-run with strong policy implications.

Many other meaningful variables, besides, should have been included in the model. Actually there is always some limit on the number of variables which can be included in the VAR model as well as on the maximum number of lags[16] and, therefore, it may happen that some variables have to be excluded prior to modelling[17]. For instance it would have been

[15] In 1965 unemployment in the Southern Italy was 6% while in the Northern and Central Italy was 5,6%. In 1996 unemployment in the South was equal to 28,8%, about three times as large as unemployment in the rest of the country, that reached 9,5% [Brunello,Lupi and Ordine (2000)].

[16] For instance for a system of 6 variables with 5 lags in each equation, the total number of regressors is 30. This could make the entire modelling process impossible if we have only a small data sample.

[17] As Spanos(1989a) pointed out, choice of variables to be included in a VAR model is a way to insert economic theory in this type of models, also referred to as *atheoretical macroeconometrics*.

interesting to insert private services sector real wage[18] in the model in order to assess if the private services can play the role of a *leading sector*, compared to industrial sector.[19]

Moreover some variables accounting for the composition of union membership could have made limitations of analysis less evident. Over the last fifteen years a striking change in the structure of union membership has been observed, with a remarkable increase in the number of pensioners and a significant rise in membership in the southern regions. [see Santi (1988), CESOS (1995) and Merolla (1996)].

Chiarini (1999) pointed out that it is implausible to rule out union membership components, like active workers' and pensioners' membership, in empirical analysis. In fact his econometric analysis shows that pensioners' and active workers' membership tend to inversely affect wages and employment level.

Section 4 and 5 provided a good description of long- and short-run properties of the model. However in order to achieve a complete outline of relationships linking variables of the system it is necessary to examine results of the impulse response analysis. To this the paper turns in the next section.

6 *Impulse response analysis*

Impulse response analysis consists of simulating an impulse in one of the variables of the system and analysing response of the whole system. In particular this analysis amounts to a dynamic simulation, from an initial value of zero, where a shock[20] at $t=1$ in a variable is traced through [see Lutkephol (1991) and Lutkepohl and Reimers (1992)].

Examination of the graphs of responses (responses to a shock in Δwp, Δpq, Δpr and $\Delta disp$ are shown, respectively, in figures from 7 to 10) gives useful information on both the stability of the system and relations between variables. When an impulse in a differenced variable is quickly absorbed by other endogenous variables[21], one can conclude that there is a good stability in the system, one of the main properties the model should have for making results of simulations reliable. Moreover it is possible to know if impulses have a

[18] Actually it was tested a 6 variables VAR, augmented with private services real wage but it showed not completely satisfying statistical properties.

[19] From standardised labour units data (ISTAT's ULA) over the period 1982:1 -1999:4 one can see clearly the reallocation in employment shares. Private services' share on total employment has risen from less than 25% in 1982 to about 33% in 1999 whereas industrial's share has fallen from more than 28% to about 22%.

[20] The one-time impulse is of size one-standard deviation. Another possibility could have been to impose a shock of unit size.

[21] In the sense that all the endogenous variables return to their previous value if no further shocks hit the system.

permanent (long-run) effect on levels of the variables by examining accumulated responses[22].

Figures 7-10 show that endogenous differenced variables when shocked tend to return to their steady-state value after only few periods, indicating that the system is quite stable.

A real wage innovation (see Figure 7) has the effect to make the unemployment rate grow for about 5 periods. Afterwards it converges to zero. The innovation seems to have a permanent effect on unemployment that settles at a higher equilibrium value. This outcome is completely consistent with Layard, Nickell and Jackman (1991)'s unemployment theory[23]. Real wage impulse seems to have no permanent effect on labour productivity. This still further confirms the presence of a wage gap (*wp-pr*).

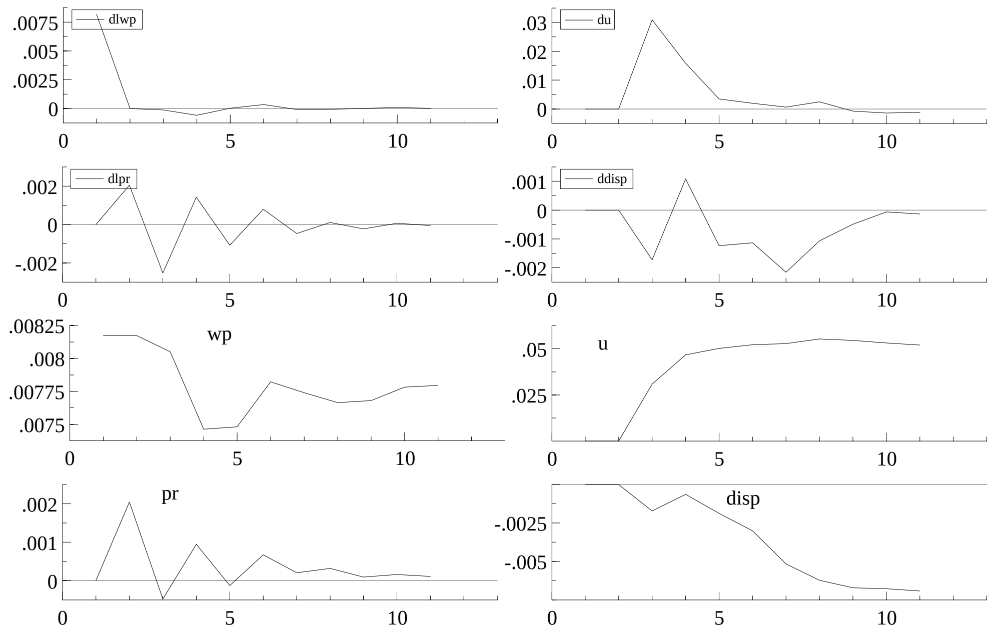

FIGURE 7: RESPONSE TO AN IMPULSE IN ΔWP

[22] A shock imposed in period *t=1* brings about a change in other variables for a number of steps (in this analysis steps refer to quarters). Each step can be a negative or a positive change. When changes are absorbed regarding variables in differences (that return to the initial value of zero), the new permanent level of the variables can be worked out by adding up changes of various steps, that is accumulating the responses.

[23] In Layard, Nickell and Jackman's "battle of mark-ups" model the equilibrium solution for natural rate of unemployment gives: $u^* = \frac{\beta_0 + \gamma_0}{\beta_1 + \gamma_1}$ where β_0 *and* γ_0 represents respectively wage push and price push, and β_1 *and* γ_1 are wage and price flexibility. A real wage impulse can be thought of as an exogenous increase of numerator in previous expression that clearly makes unemployment rise.

The reaction of the dispersion measure is peculiar since the variable *disp* settles at a lower equilibrium value after the real wage shock. This result, together with the previous one (unemployment increases after real wage impulse), could be interpreted against the hypothesis of a positive relationship between unemployment rate and sectoral employment dispersion.

Actually when an exogenous shock brings about a rise in unemployment, it means that there is a lower demand pressure in the labour market and so reallocation across sectors is likely to be quite low.

Figure 8 displays the effects of an impulse in the dispersion measure variable. As expected the unemployment level is permanently higher after the shock. This confirms that when permanent[24] reallocation across sectors is high, there is a mismatch effect having a negative outcome on employment. Actually interpretation of the unemployment response to an employment dispersion's shock seems to be much more intricate than it appears : unemployment tends to decrease just after the impulse and only after some delay (about 6 periods) it settles at a permanently higher level. Real wage and productivity's response is consistent with the temporary reduction in unemployment as real wage goes down and productivity increases after the shock. This peculiar behaviour of unemployment following the innovation in the employment dispersion across sectors, suggests the following interpretation. When employment dispersion is high, reallocation across industrial sectors is in progress, or in other words this means that there are growing sectors where labour demand is high and other sectors from which employees are being expelled. At the outset of the reallocation, the firms in dynamic sectors tend to employ more workers in order to allow their expansion and, therefore, they bargain over lower real wage. Some features of the Italian labour market could explain how firms manage to lower wages during sectoral reallocation: first of all firms can employ a huge number of low-specialised unemployed people, especially coming from southern regions of Italy. Moreover in the Italian industrial sector there is only low demand for hi-tech employment[25]. Hence, firms in growing sectors do not find it difficult to bargain lower wages since they need mostly non-specialised workers. Employment growth in expanding sectors is greater than the fall in contracting sectors' employment, and so, at the beginning of the sectoral reallocation, aggregate unemployment decreases. After some time (impulse response analysis indicates 4 periods, or one year) unemployment begins to rise and, finally, settles at a higher permanent level than the initial value. This happens when the mismatch effect becomes so great that it is able to offset the employment gain of the outset of the reallocation.

[24] The employment dispersion index used in the estimate is built as to catch permanent part of employment reallocation across sectors.

[25] There are several indicators that give evidence of a poor public and private demand in Italian hi-tech industries and in R&D sector. A comprehensive survey on the subject is the European Commission's "Key Figures 2001" on research policies [European Commission(2001)], that shows several indicators worthy of comment: patents per million population shows that Italy applied for only 62 patents in 1999 with an EU's average of 125. Total R&D expenditure in relation to GDP, then, is lower than in the EU as a whole (Italy:1,04 EU:1,92 in 1999) as is R&D expenditure financed by industry in relation to industrial output (Italy: 0,60 EU:1,42 in 1999).

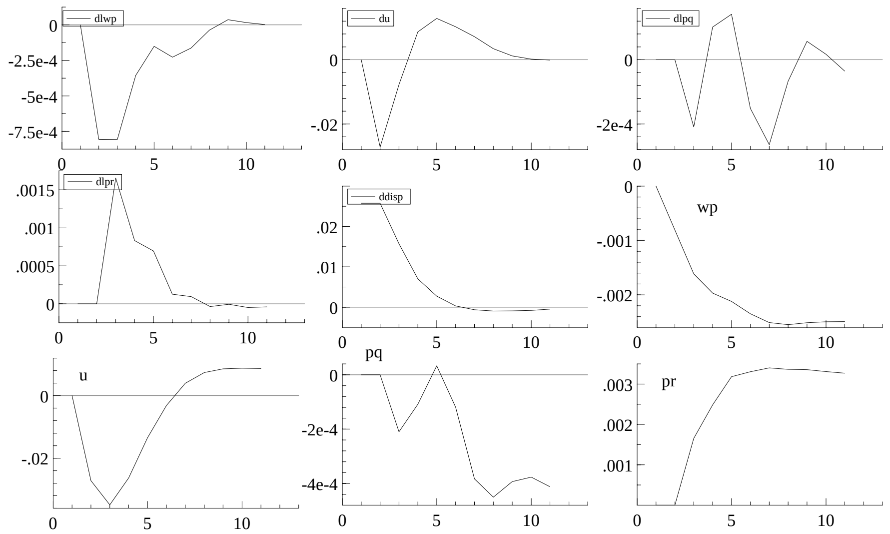

FIGURE 8: RESPONSE TO AN IMPULSE IN ΔDISP

Previous interpretation of unemployment reaction to a shock in employment dispersion is clearly limited and this should be seen only as a suggestion for further analysis on unemployment-sectoral reallocation links.

As far as response to a shock in labour productivity is concerned, one can see from Figure 9 that unemployment tends to decrease just after the shock is imposed. Then unemployment starts growing only when employment dispersion begins to rise. It is again confirmed that employment reallocation has an effect on unemployment only after some delay. Something peculiar is that the real wage settles at a lower level after the productivity shock, even though change is very low (-0,01).

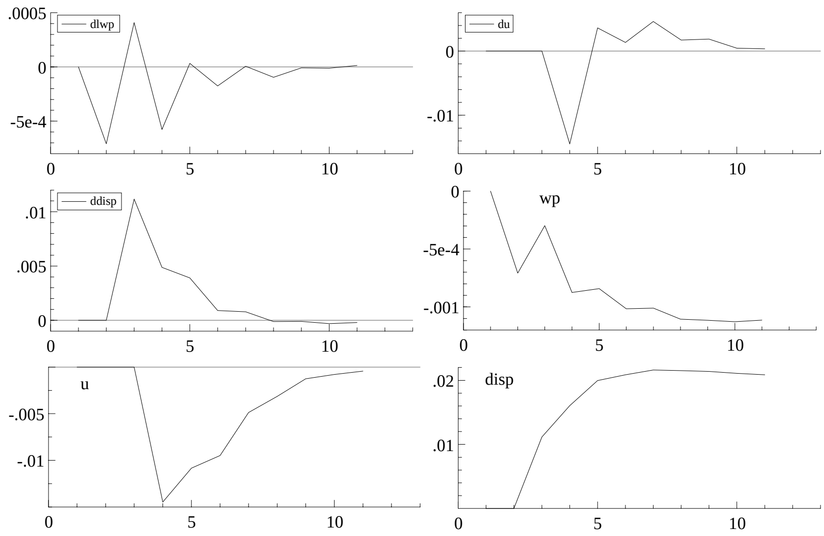

FIGURE 9: RESPONSE TO AN IMPULSE IN ΔPR

Lastly Figure 10 reports the system's reaction to a price wedge shock. The impulse has a striking effect on dispersion measure and, therefore, on unemployment. In fact both variables take a lower equilibrium value. The response of the real wage and the productivity is consistent with previous movements as they both settle at a lower value. Hence any expansionary macroeconomic policy that brings about a price rise should be able to reduce unemployment.

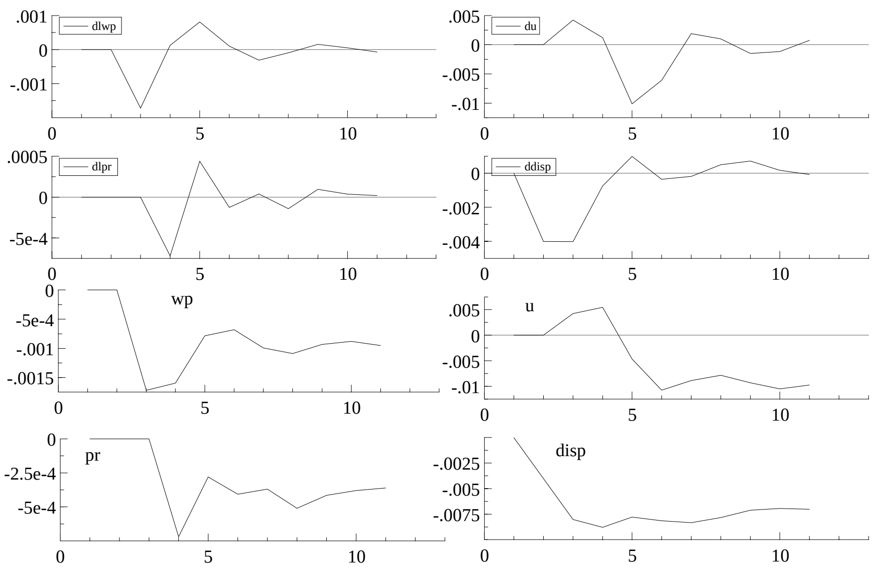

FIGURE 10: A RESPONSE TO AN IMPULSE IN ΔPQ

7 *Concluding remarks*

In this paper we investigate the estimation procedure of the NAIRU models. In particular the paper deals with the short- and long-run effect of the employment sectoral reallocation on the unemployment rate in the Italian economy.

A cointegrated VAR was used in order to carry out the estimates. It is well known that this kind of methodology is an effective way to overcome several problems affecting the standard structural approach to the estimation of the NAIRU. Following Spanos' suggestions, a model that has the property of *statistical adequacy* is specified in this paper.

Empirical evidence provides several interesting results. Cointegration analysis shows that real wage cointegrates with unemployment, productivity, price wedge and the employment dispersion variable in the long run. Then stationarity of the wage-gap (*wp-pr*) is rejected by the appropriate LR test. This seems to give evidence of a serious rigidity of real wage.

A further result is interesting: a rise in the permanent shift of the employment composition brings about a permanent increase in the unemployment. This means that permanent employment reallocation is a significant determinant of unemployment.

It is also worthy of comment that the effect of permanent shift in employment composition on unemployment occurs only after some delay. In particular impulse response analysis shows that at the outset of the employment reallocation, the unemployment tends to decrease. After some time (impulse response analysis indicates 4 quarters) the mismatch effect, brought about by the reallocation, becomes so large that unemployment begins to rise and settles at a higher permanent value.

Finally, within a multivariate cointegrated space, the NAIRU has to be thought of as a hyperplane with dimension *n-r* (*n*=number of endogenous variables and *r*=number of cointegrating vectors), rather than as a single point. In this context it is possible to investigate the NAIRU only taking into account the full system, that is short-term structure and impulse response analysis.

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APPENDIX A1 - THE EMPLOYMENT DISPERSION MEASURE

The employment dispersion variable used in the empirical part of this work has been derived on the basis of Neumann-Topel's employment-based dispersion index . In particular it was used an index based on employment distribution in 10 industrial sectors according to the ISTAT classification, as that used in Chiarini and Piselli (2000).

Construction of the sectoral shift measure undergoes 3 steps:

a) *define a direction of a permanent change*.

Employment shares, λ_i, are defined as the ratio of sectoral employment to the total industry employment, or $\lambda_i = \frac{n_i}{N},\ i = 1, 2, \ldots, 10$, where n_i is the employment in i-th sector and N is the total industry employment. Considering $\lambda = (\lambda_1, \lambda_2, \ldots, \lambda_{10})$ a vector of employment shares, Neumann and Topel define the direction of a "permanent" change in industry composition shift as follows:

(A1) $$\Delta\hat{\lambda} = \sum_{j=1}^{J} \sigma_i \lambda_{t+j} - \sum_{j=1}^{J} \sigma_i \lambda_{t-j}$$

Equation (A1) represents the difference between moving averages of future and past vectors of employment shares. The parameters σ_i stand for smoothly declining weights while J is the time horizon. In this paper σ_i =0,9^J with J=8 quarters[26].

b) compute the actual difference between current and past employment distribution.

Given the equation (A1) the difference between the current and the past employment distribution is defined as:

(A2) $$\Delta\lambda_t = \lambda_t - \sum_{j=1}^{J} \sigma_i \lambda_{t-j}$$

c) determine the least squares projection of (b) onto (a).

Considering the difference (A2) as a linear combination of a permanent effect and transitory component it is possible to write:

[26] The choice of σ in the smoothing procedure seems to be important in describing the bahaviour of employment dispersion index. σ should be chosen so as to take into account pecularities of the sample period.

(A3) $$\Delta\lambda_t = \beta\hat{\lambda}_t + \Delta\lambda_t^T$$

where $\beta = \dfrac{\Delta\hat{\lambda}'\Delta\lambda}{\Delta\hat{\lambda}'\Delta\hat{\lambda}}$, $\beta\Delta\hat{\lambda}$ is the least square projection of the current changes onto the vector measuring the direction of "permanent" changes, and $\Delta\lambda_t^T$ is the transitory (orthogonal to $\Delta\hat{\lambda}$) component. In the Neumann and Topel's terminology, the vector measuring permanent shifts is given by the following expression.

(A4) $$\Delta\lambda_t^P = \beta\Delta\hat{\lambda}$$

When "permanent" shifts are zero, over time each sector keeps its share of employment and each component of $\Delta\lambda_t^P$ is zero. If employment shifts, shares will change in the future compared to the past and, therefore, some components will not be zero. A measure of the importance of the shifts among sectors is provided by the euclidean lenght of $\Delta\lambda_t^P$:

(A5) $$\left|\Delta\lambda_t^P\right| = \left[\sum_{j=1}^{J}\left(\Delta\lambda_i^P\right)\right]^{\frac{1}{2}} = \delta_t$$

The scalar δ_t is the sectoral shift variables used in the estimate.

APPENDIX A2 - VARIABLES AND SOURCES

In the following table an outline of time series used in this paper is displayed. In the first part of table there are time series (upper-case) from which variables used in the estimates (lower-case) have been derived.

Sample period : 1984:1 -- 1997:4

Data frequency : quarterly

Sources : all variable refers to ISTAT's "Quarterly National Accounts" except for unemployment rate, u, that comes from ISTAT's "Quarterly Labour Force Statistics", but it is in standardised version provided by OECD.

Name	***Description***	
RET	Total gross compensation	
ULAT	Total standardised labour units	
CFIK	Private final consumption expenditure, constant prices	
CFIC	Private final consumption expenditure, current prices	
VAPBC	Base prices added value, current prices	
VAPBK	Base prices added value, costant prices	
STUR	Standardised unemployment rate (OECD)	
		Algebra
w	Gross compensation per employee	RET/ULAT
p	Private final consumption deflator	CFIC/CFIK
wp	Real wage	w/p
q	Base prices added value deflator	VAPC/VAPBK
u	Standardised unemployment rate(OECD)	STUR
pq	Price wedge (or relative prices)	p-q
pr	Labour productivity	VAPBK/ULAT
disp	Sectoral employment shift measure	see Appendix A1

APPENDIX A3 - UNIT ROOT TEST

A3.1 Augmented Dickey-Fuller (1984:1 1997:4)

Critical values: 5%=-3.497 1%=-4.142; Constant and Trend included

	AR order	ADF-t
wp	1	-2,012
u	2	-1,9294
pq	3	-2,5933
pr	1	-2,8959
disp	2	-2,5866

Critical values: 5%=-2.918 1%=-3.56; Constant included

	AR order	ADF-t
Δwp	1	-5,0404**
Δu	1	-3,1207*
Δpq	1	-8,5347**
Δpr	1	-5,9907**
$\Delta disp$	1	-3,4324*

A3.2 Perron-modified(P-mod) [Perron (1989)] and Phillips-Perron (Ph-P test) [Phillips and Perron (1988)] test for variables Δu e $\Delta disp$

	AR order	ADF-t	P-mod	Ph-P test	Dummy
Δu	1	-3,1207*	-8,0429**	-3,2461*	Dadd92q1
$\Delta disp$	1	-3,4324*	-3,8721**	-3,2325*	Dadd91q1

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