Unemployment, Wage Pressure and Sectoral Shifts: Permanent and Temporary Consequences of Intersectoral Shocks

Bruno Chiarini, *Istituto Universitario Navale Di Napoli*

Paolo Piselli, *Banca D’Italia*


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*Address correspondence to Bruno Chiarini, Istituto Di Studi Economici, Facoltà Di Economia, Istituto Universitario Navale Di Napoli, Via Medina 40, 80133 Napoli, Italy. E-mail address: chiarini@nava1.uninav.it*

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1. INTRODUCTION

The purpose of this paper is to analyze the effect of permanent sectoral shifts on unemployment and the wage formation process. It is well known that the unemployment rate can increase independently of aggregate conditions if labour adjusts slowly to shifts of employment demand between sectors of the economy (Lilien 1982).

Whether sectoral, rather than aggregate, shocks are the key factor responsible for fluctuations in the unemployment rate is crucial. According to the sectoral shifts hypothesis, fluctuations in demand across sectors account for a substantial fraction of the variation in unemployment. 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 sector that are expanding, or if the labour market shows substantial rigidity.

The unemployment rate, therefore, is no longer an accurate measure of labour market activity. However, in the long run, the effect of sectoral reallocations should not affect the wage inflation process. If this hypothesis is correct, one feature of particular interest in estimates of conventional Phillips curve-type models is that the effect of excluding an employment based dispersion measure to filter out the effects of sectoral shifts entails unstable models.

A further interesting question is whether the employment imbalances across sectors are temporary or whether they persist over time and, therefore, whether they affect the natural rate of unemployment.

The main purpose of this study is to examine the effect of allocative disturbances using Italian labour market data and a VECM (vector error correction model). This exercise is useful because much of the empirical research has focused on United States and United Kingdom data, whereas the Italian economy has received little (if any) attention. This is quite strange, since Italy is seen as a good example of what is known as Eurosclerosis. Lilien argued that an important part of cyclical unemployment comes from permanent shifts in the composition of labour demand. Our findings emphasize that an important source of the persistence of unemployment comes from sectoral shifts. Thus, unemployment persistence may rise not (or not only) because of lower total demand for labour but rather, because of labour market imperfections, costs of adjustment and wage rigidity, workers do not move between sectors.
The paper is organized as follows. Section 2 reports on Neumann and Topel’s (1991) measure of sectoral shifts based upon dispersion in employment growth across industrial sectors. Sections 3 and 4 report empirical results for Italy using quarterly data (consumption real wage, labour productivity, price wedge, worked hours, unemployment rate and a sectoral shift variable) for the period 1975:1–1993:3. Using non-stationary variables and the maximum likelihood procedure proposed by Johansen, we first estimate a cointegration space, and subsequently, a VECM. The impulse response analysis and a geometrical representation of the cointegration space may give further support to the reason why it might be more useful to think in terms of ranges for equilibrium unemployment (see Cross & Darby, 1996). Section 5 uses the methodological and empirical work for drawing some policy implications. Section 6 concludes.

2. A SECTORIAL SHIFT MEASURE

Fig. 1 plots the evolution of employment shares in Italy for 10 two-digit industries (eight manufacturing industries plus energy products and construction). Reallocation is clearly evident by the distinct trends in the shares. Fig. 1 also illustrates how the composition of employment in the industry sector has been altered in the period examined. The changes have been particularly intense for fabricated metal products and construction, but the distinct evolutions in most industries’ shares (basic metal industries, chemical products, food, beverages and tobacco, energy products) indicate that reallocation has changed the composition of employment over the last two decades. A further characteristic is the variability of the industries’ relative shares through time. To this end, it is interesting to note that much of the variations in employment shares reported in Fig. 1 do not occur in and around recessions.

1See Juselius (1993) for an excellent discussion of the economic motivation underlying some recent developments in the macroeconometric analysis of time series.

2Energy products, basic metal industries, non-metallic mineral products, chemical products, fabricated metal products, transportation equipment, food, beverages and tobacco, textiles, apparel, leather and footwear, wood products and furniture, paper, rubber and others and construction.
In order to measure sectoral shifts in labour demand, we use Neumann-Topel’s employment-based dispersion index. A dispersion index based on employment was used by Lilien (1982). He provided empirical evidence from the post-war United States of a positive correlation between aggregate unemployment rate and a sectoral shift proxy measured by weighted standard deviation of employment across sectors. Although Lilien’s sectoral shift hypothesis was substantially accepted in the 1980s, Abraham and Katz (1986) stressed the low power of this proxy in capturing sectoral shifts, as the business cycle is not neutral across different markets: a shift in employment in some sectors might be generated by a

Lilien suggests the standard error of interindustry employment growth rates as a possible measure. Ideally, a measure of dispersion in labour market conditions should be related to the economic cost associated with interindustry employment changes. Of course, there are other dimensions to structural changes in labour demand (educational, age, skill and geographical distributions). These changes affect labour demand within sectors, influencing the extent of structural adjustment.
cyclical effect instead of representing the consequence of a permanent relative change in labour demand. Thus, sectoral shifts and aggregate demand explanations may be observationally equivalent.⁴

Neumann and Topel consider Abraham and Katz’s critique and define a new proxy, decomposing the variability of employment shares in two components, with the aim of isolating the “permanent” changes in labour demand among sectors from changes due to local cycles or other unpredictable events (transitory effects). A “permanent” change in the distribution across sectors, producing mismatch, increases the unemployment rate in the short run. We use a similar index, based on employment distribution in 10 industrial sectors, according to Istat classification (eight manufacturing sectors plus the energy and construction sector).

To construct a measure of sectoral shifts we follow three steps: (a) define a direction of a permanent change; (b) compute the actual difference between current and past employment distribution; (c) determine the least squares projection of (b) onto (a).

Employment shares are defined for each sector $S_i$ as the ratio of sectoral employment to the total industry employment $l_i$, $i = 1, 2, \ldots, 10$. Considering $\lambda_i = (\lambda_{i1}, \lambda_{i2}, \ldots, \lambda_{i10})$ a vector of employment shares, Neumann and Topel define the direction of a “permanent” change in industry composition shift as follows:

$$\Delta \hat{\lambda} = \sum_{j=1}^{J} \sigma_{ij} \lambda_{i+j} - \sum_{j=1}^{J} \sigma_{ij} \lambda_{i-j}.$$  

Eq. (1) represents the difference between moving averages of future and past vectors of employment shares. The parameters $\sigma_i$ stand for smoothly declining weights while $J$ is the time horizon. In our case, we set $\sigma_i = 0.9^i$, with $J = 8$ quarters. Given Eq. (1), the difference between the current and the past employment distribution is defined as (Eq. (2)):

$$\Delta \lambda_i = \lambda_i - \sum_{j=1}^{J} \sigma_{ij} \lambda_{i-j}.$$  

Considering the difference as a linear combination of a permanent effect and a transitory component we can write that

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⁴See also Blanchard and Diamond (1989), Brainard and Cutler (1993), Chareyre and Kaufmann (1987), David (1987a), Lilien and Hall (1986), Padoa Schioppa (1991), Davis and Haltiwanger (1992) and Pissarides (1990) among others. Loungani, Rush, and Tave (1990), attempt to sort out this empirical equivalence, constructing a stock-price index to measure sectoral shift in labour demand. Causality is clear in the stock-index case: unemployment depends on up to three lags on the stock prices dispersion index. On the contrary, the correlation between unemployment and Lilien’s proxy is contemporaneous, enforcing the aggregate demand hypothesis versus sectoral shifts.
\[ \Delta \lambda_i = \beta \lambda_i + \Delta \lambda^p; \quad \beta = (\Delta \lambda' \Delta \lambda / \Delta \lambda' \Delta \lambda)^{1/2}. \]

\( \beta \Delta \lambda \) is the least squares projection of the current changes onto the vector measuring the direction of “permanent” changes, and \( \Delta \lambda^T \) is the transitory (orthogonal to \( \Delta \lambda \) ) component. In the Neu-
mann-Topel terminology, the vector measuring permanent shifts is given by the following expression, \( \Delta \lambda^p = \beta \Delta \lambda \).

When “permanent” shifts are zero, over time each sector keeps its share of employment and each component of \( \Delta \lambda^p \) is zero. If employment shifts, shares will change in the future compared to the past and, therefore, some components in \( \Delta \lambda^p \) will not be zero.

A measure of the importance of the shifts among sectors is provided by the Euclidean length of \( \Delta \lambda^p \) (3):

\[ \left| \Delta \lambda^p_i \right| = \left( \sum_{j=1}^{J} (\Delta \lambda^p_j)^2 \right)^{1/2} = \delta_i \]  \hfill (3)

where \( \lambda_i \) is a component of \( \Delta \lambda^p \) corresponding to the sector \( i \). The scalar \( \delta_i \) is our sectoral shift variable. Its length varies from 0 to \( \sqrt{2} \).

The choice of \( \sigma \) in the smoothing procedure described above seems to be important in describing the behaviour of \( \delta_i \). Although changes in the time horizon and in the weights may provide a certain effect on the sectoral shifts measure, we find \( \delta_i \) appropriate to take into account some peculiarities of the sample period.\(^5\)

**2.1. The Dispersion Measure \( \delta_i \)**

Fig. 2 presents the dispersion measure \( \delta_i \), constructed from Neum-
mann and Topel’s measure of sectoral shifts. Since data are quar-
terly, over the period 1973:1–1995:3, the measure is constructed for the period 1975:1–1993:3. The higher its value, the greater is the permanent shift in composition. The importance of sectoral shifts seems to vary considerably over time. \( \delta_i \) shows four major changes in the distribution of employment. A striking and pro-

\(^5\)We have considered a further employment-based dispersion index, based on employ-
ment distribution in 10 industrial sectors (as for \( \delta \)) and the service sector. However, the new index did not have a strong impact on our results.
of employment growth rates were growing simultaneously with a decline of economic activity (industrial production and GDP) in the first, second and fourth period, whereas the third period (1986–1987) was characterized by marked economic vitality.

The allocative effects due to raw material and oil price shocks can explain the large shifts increases between 1973 and 1977 (evidence of the allocative effects of oil price shocks is given in Hamilton, 1983; Loungani, 1986; Keane & Prasad, 1996, among others). Moreover, firms were proceeding with a capital-intensive adjustment process related to the significant wage increases that occurred in the late 1960s.

The 1980s opened with a recession. A new oil and raw material price increase shocked the economy, increasing production costs and starting off an inflationary process. These factors led the industrial structure to begin a renewal equipment process, lowering employment. The third peak is characterized by inflation deceleration and a revaluation of the exchange rate of the lira. The effects of joining the exchange rate mechanism of the EMS induced economic restructuring, achieved through massive investments in advanced technology equipment and labour offs. The
exchange rate revaluation effect provides compelling support for the sectoral shifts view also in the last peak, which gives evidence of the rationalization of the productive process, with a reduction of investments and employment.

Some authors (e.g., Davis, 1987b; Hamilton, 1983) stress that reallocation in response to a sectoral disturbance will occur when the economy is in a low state and the opportunity cost is smallest. Rissman (1993), according to this “reallocation timing argument”, emphasizes that it would make unemployment Granger-causing dispersion to the extent that the unemployment rate also reflects cyclical disturbances. Our dispersion measure $\delta$, does not seem to give full support to this hypothesis, relating allocative disturbances with an expansionary phase. Below, we use a VAR model to show that an innovation in unemployment variable has no effect on the other variables. That is, unemployment does not Granger-cause the structural shift measure.

3. COINTEGRATION ANALYSIS

3.1. The I(1) Model for Cointegration

We consider a general vector autoregressive model for a $n$-dimensional vector process $x_t$. When the data are $I(1)$ a useful reformulation of the model is to error correction form:

$$
\Delta x_t = \Gamma_1 \Delta x_{t-1} + \Gamma_2 \Delta x_{t-2} + \Pi x_{t-1} + \phi D_t + \varepsilon_t. \quad (4)
$$

The errors are assumed to be independent Gaussian variables with mean zero and variance $\Sigma$. $D_t$ is a vector of deterministic terms (constants, linear terms, intervention dummies). The lag length in Eq. (4) corresponds to the actual lag length of the VAR models discussed below, and has been determined by some information criteria. Starting from a VAR with four lags on all stochastic variables, simplification tests suggested that three lags sufficed (see Lütkepohl, 1991, for a survey of the criteria).

$P = \alpha \beta'$ has reduced rank $r$ and $\alpha$ and $\beta$ are $(n \times r)$ matrices. Johansen shows that the reduced rank of $\Pi$ implies that the stochastic part of $\beta'x$ is stationary and, therefore, with this model we can consider the economic relations $\beta x = \xi$ and assume that the agents react to the disequilibrium error $\beta x_{t-1} - \xi$ through the coefficient $\alpha$. 
The representation theorem of Granger (a way of finding the MA representation from the AR representation, and vice versa, see Johansen, 1995) shows that the solution of Eq. (4) is given by:

\[ x_t = x_0 + C(1) \left( \sum_{i=1}^{\gamma} \varepsilon_i + \sum_{i=1}^{\rho} \phi D_i \right) + C^*(L)(\varepsilon_t + \phi D_t), \]  

(5)

where \( C^*(L) = [C(L) - C(1)](1 - L)^{-1} \) and \( C(1) = \beta_{\perp} (\alpha_{\perp}', \Gamma \beta_{\perp})^{-1} \).

\( \alpha_{\perp} \) and \( \beta_{\perp} \) are orthogonal complements of \( \alpha \) and \( \beta \), respectively. The orthogonal complement of \( \beta \), \( \beta_{\perp} \), is defined as the \((n \times (n - r))\) matrix such that \( \beta' \beta_{\perp} = 0 \) with \( r(\beta) = m - n \) and \( r(\beta_{\perp}, \beta) = n \). Thus, \( \beta_{\perp} \) spans the null space of \( \beta \). Notice that the process \( x_t \) has a deterministic trend of the form \( C(1)(\varepsilon_t + f D_t) + C^*(L)\varepsilon_t \).

It is seen from Eq. (5) that the cumulative shocks \( \alpha_{\perp} \sum_{i=1}^{\gamma} \varepsilon_i \) drive the economic variables to lie in the space spanned by \( \beta_{\perp} \) (the attractor set). The matrix \( C(1) \) describes how the stochastic and deterministic trends determine the long-run components of \( x_t \). This matrix is of reduced rank and we will see below that, in our case, it implies that there exist \( n - r = 4 \) common trends (a geometrical representation of a simplified version of the space spanned by \( \beta_{\perp} \) is reported in Appendix A).

A necessary condition for the I(1) model (5) is that \( \alpha_{\perp}' \Gamma \beta_{\perp} \) has full rank (otherwise the process \( x_t \) is integrated of second order). In both models (4) and (5), restrictions are required to identify, respectively, uniquely relations and trends (i.e., if \( \beta' \) is a cointegration matrix, so is \( \beta'F \) with any non-singular \( r \times r \) \( F \)).

To identify uniquely the individual long-run relations, it is usual formulate linear restriction \( R_i \) as full rank matrices \((n \times r_i)\), \( R_i \beta = 0, i = 1, \ldots, r \). Johansen and Juselius (1994) define a matrix \((n \times (n - r))\), \( H_i = R_{i1} \), and conveniently write \( \beta = H_i \phi_i \) for some \( n \times r_i \) dimensional vector \( \phi_i \). The matrix \( H_i \) reflects linear (economic) hypotheses to be tested against the data. This allows to write the rank-condition for identifying the equations. In the cointegration analysis below, we find \( r = 2 \) stationary relations, which requires the following condition \( r_{ij} = r(R_iH_j) \neq 1, i \neq j \).

The cointegration analysis illustrated below, shows that we restrict the space \( H \) further (imposing and testing over-identifying restrictions). From Eqs. (9) and (10), it may easily be verified that this is done without decreasing the rank-condition.
3.2. The Data Set

Data are quarterly, seasonally adjusted over the period 1975:1–1993:3. The system is in six stochastic variables $x' = [(w - p), \pi, (p - q), h, U, \delta]$. The variables are, respectively, consumption real wage, labour productivity (value added/labour units), price wedge (consumption price minus product price), worked hours, unemployment rate and the sectoral shift variable defined above. The variables refer to the industry sector while the worked hours variable is a proxy, being per capita hours worked in the “large firms” (firms with more than 500 employees). All the variables are in log form, except for the sectoral shift variable $\delta$. Finally, all the variables are non-stationary ($I(1)$) time series (the univariate evidence is provided in the JPM website version).

3.3. The Cointegration Space and the Natural Rate of Unemployment

Using a VAR(3) model and Johansen’s tests and ML estimation method (Johansen, 1988; Johansen & Juselius, 1990), two cointegration relations were found. The variables characterize most empirical works based on the bargaining framework. In searching for a restricted cointegrating space, we set down a stylized version of the battle of the mark-ups approach to the natural rate of unemployment.

It is well known that this model makes the wage equation unidentified (Bean, 1994; Manning, 1993). Suppose that wage-setting and pricing equation are estimated in the following form:

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1. There is a vast literature on seasonal adjustment and the impact of SA on cointegrated time series (e.g., Franses, 1996). In particular, the application of linear moving average filters may yield non-invertible MA processes for the SA time series. Similar results hold for Census X-11 corrected time series (the seasonal adjustment method often applied in practice). See Maravall (1995) for a discussion of the non-invertibility of SA time series.

2. The degree of integratedness of the unemployment rate and the sectoral shift variable may be considered a bit puzzling: in fact, both the variables are bounded and seem to be $I(1)$. However, the unemployment rate rises significantly during the last 20 years, changing its average level (see Fig. 2 for $\delta$). Moreover, the order of integration is a statistical property of the series within the considered sample and not necessarily a property of the generating process.

3. The following multivariate tests show the congruency of the initial general system: vectAR(1,5): $F(180,108) = 1.246[0.106]$; vectNorm: $\chi^2(12) = 8.657[0.732]$; vectHet: $F(756,22) = 0.05[1.00]$. 

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where the vector $X_1$ contains demand side variables and $X_2$ is a vector of variables like union power and other variables that determine wage pressure. As stressed by Manning (see also Westaway, 1996), even if the cross-equation restrictions $\mu_2 X_1 = -\phi_1 X_1$ are imposed, so that the invariance of the natural rate of unemployment to demand factors holds, Eq. (6) remains unidentified. This is easily seen from the fact that adding a multiple of the price-setting equation to the wage equation alters only the coefficients of the wage-setting schedule (Eq. (6)):

$$(w - p) = \sigma_0 + \sigma_1 U + \sigma_2 X_1 + \sigma_3 X_2$$

where $\sigma_0 = \Phi \mu_0 - (1 - \Phi) \phi_0$; $\sigma_1 = \Phi \mu_1$; $\sigma_2 = \Phi \mu_2 - (1 - \Phi) \phi_1$; $\sigma_3 = \Phi \mu_3$; $0 < \Phi < 1$ (for a critical analysis, see Chiarini & Piselli, 2000).

In this context, the multivariate cointegration approach may play a new role in ensuring that the wage-price structure defined does not suffer from the potential lack of identification of wage-setting relations.

The statistical model was checked to be “identifying” (see Johansen & Juselius, 1994). Subsequently, further restrictions on the parameters space were imposed and tested, obtaining the following LR test, $\chi^2(3) = 2.763[0.4296]$. The restrictions placed on the two vectors are as follows: three restrictions on $\beta_1$, comprising two zero constraint ($w - q = 0$ and $U = 0$) and a homogeneity constraint ($w - p = \pi$). There are also two restrictions placed on $\beta_2$, ($w - p = \pi$ and $\delta = 0$). These restricted cointegration vectors were tested for lying in the cointegration space, jointly with testing for labour productivity and unemployment being weakly exogenous. We found the productivity variable exogenous to the system ($\chi^2(6) = 5.635[0.4653]$).

$$(w - p) - \pi + 1.38 h + 0.0156 \delta + 0.77$$

$t$-statistic in parentheses). A simplified geometrical representation of this cointegration space is provided in Appendix A.

It is important to emphasize that $U_i$ is weakly exogenous. This means that the process determining $U_i$ ($\Delta U_i$), which is pushed
along the attractor set by the common trends, does not react to the disequilibrium error \( \beta' x \), via the adjustment coefficients. Of course, this does not mean that \( U_t \) does not move together the others variables (see Johansen, 1992).

In the cointegration space defined by the two stationary relations, unemployment allows a long-run equilibrium between the bargained real wage and the demand of employers (the feasible real wage). It is worth stressing that setting the unemployment restrictions equal to zero in both the cointegrating vectors, provides non-stationary relationships among the variables.

Restricting the demand factors in the price and wage equation to be the same with opposite sign does not provide an acceptable result in terms of LR test. However, our stationary relations identify wage and price-setting schedules without relying on arbitrary (or ad hoc) identifying exclusion restrictions (see Manning, 1993). The first cointegrating vector, Eq. (9), may be interpreted as a price equation. It will depend on unit labour cost, a measure of intrasectoral labour reallocation and a constant (that, for instance, can be related to the degree of product market competition).

Labour reallocation tends to determine a lower real wage that, for given productivity, price setters are willing to concede. The second stationary relation (Eq. (10)) may be interpreted as a structural equation for the mark-up of real wages (labour share) that depends negatively on the unemployment rate and prices wedge. In the context of Eqs. (6) and (7), we have:

\[
(w - p) = \sigma_0 - \sigma_\delta u + \sigma_\gamma X_1 + \sigma_\delta X_2 + \sigma_\delta
\]

where \( \sigma_\delta = (1 - \Phi)\phi_2 \). Clearly, Eq. (11) does not have the same form as the original wage-setting schedule (Eq. (6)). However, it is important to stress that the stationary relations (9) and (10) cannot be interpreted individually, ignoring the whole system. That is, the short-run dynamics and the adjustment processes (see Chiarini & Piselli, 2000; Lütkepohl, 1991). Below, we return to this problem using impulse responses analysis.

Notice that both the equations constitute testable questions as to whether the hypothetical relation can be assumed to lie in the stationary part of the space spanned by the non-stationary variables. The likelihood ratio tests indicate that stationarity has to be rejected in all the alternative cases.

### 3.4. An Economic Interpretation

The restricted cointegration space illustrates clearly that, in this framework, the natural rate of unemployment will depend on
This result seems to question the sectoral shifts hypothesis as a phenomena associated with temporary consequences of intersectoral shocks. In fact, it shows that the shifting composition effect of labour demand does not net out on average, influencing the stationary space (wage and unemployment) in the long run (although the service sector, in the period examined, absorbed much of industry’s loss, the reallocation shocks affected the unemployment rate). As we shall see below, taking the short-term dynamics into account and carrying out the impulse response analysis provides a similar conclusion. Thus, more than frictional (or transitional) unemployment related to individuals changing jobs across sectors, we obtain a persistent imbalance between supply of and demand for labour across sectors. Increased mismatch reflects the failure of relative wages to adjust or inadequate mobility.

This (long-run) mismatch phenomenon is perhaps more worthy of study than the conventional composition hypothesis with short-run turbulence, and it arises essentially from the supply side. The literature on Italian labour market mobility seems to provide support for these results: low workforce adaptability; low willingness to change employment status; high retraining costs; low geographical mobility; high employment protection regulations on hiring and firing; high inefficiency of public placement offices (see among others, Attanasio & Padoa Schioppa, 1991; De Luca & Bruni, 1993; Galli et al., 1996; OECD, 1995).

Moreover, wage differentials have not been big enough to match the cost of moving from one sector to another. In the period examined, wage structure within sectors was influenced by two factors. The first is the wage indexation mechanism (scala mobile). In 1975, an agreement between unions and entrepreneurs considerably increased the degree of coverage and the frequency of adjustments of the scala mobile, altering the structure of remuneration by reducing wage differentials and the importance of the other remuneration components. Although a series of modifications were introduced, this mechanism ceased to exist in the last years of the sample considered. The second element, affecting wage structure, was the egalitarianism promoted by the unions
during collective bargaining (see, for instance, Erikson & Ichino, 1994). At least until the mid-1980s, intersectoral (and regional) wage differentials narrowed, and wage dispersion in manufacturing picked up in the second half of the decade, but in the first years of the 1990s, its value did not reach that obtained at the beginning of the 1970s. Thus, although wage levelling has been opposed by employers by using wage drift, the sectoral shifts could not produce downward wage pressures in some sectors and, as a reaction, upward wage pressures in others.

4. IMPULSE RESPONSE FUNCTIONS

Having determined the cointegration space, we reformulate and estimate (by FIML) the model using a general to specific modelling approach and imposing specific restrictions on each equation (the model is reported in the JPM website version).

The impulse responses for the model are shown in Figs. 3–5. We find that fluctuations in the pace of labour reallocation across sectors are a large component of long-run unemployment rate fluctuations. This finding, as emphasized by Abraham and Katz (1986), Campbell and Kuttner (1996), David (1987b), Palley (1992) and many others, highlights the importance of the central controversy in sectoral shift literature: is cyclical unemployment determined by aggregate demand or should it be related to sectoral shifts explanations? A measure of sectoral shifts should be independent of past unemployment; the unemployment rate should not Granger-cause sectoral shifts. The model and the impulse response functions reported in Figs. 3–5 confirm this requirement.
Consistent with a key prediction of the sectoral shifts literature, we find that an unfavourable allocative disturbance increases aggregate unemployment in the short run, but it does reduce employment persistently, shifting industry employment shares in the long run.9

These findings imply a low rate of job reallocation and indicate a difficult reshuffling of employment opportunities across sectors, with persistent sectoral-level employment changes. This may indicate that job substitution is typically associated with long-term joblessness.

It is interesting to note that a transitory wage shock is estimated to account for most of the long-run behaviour of unemployment, giving rise to a permanent shift in the composition of labour demand. On the other hand, there is no significant short-run or permanent effect on real wage due to an innovation in the shift variable, although the latter leads to a long-term increase in unemployment. Real wage rigidity and a more rapid structural shift tends to introduce inertia into the firms’ employment decision, raising the flows of people into unemployment and increasing the pool of those who are unemployed between jobs.

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9It is reasonable to interpret the reactions of the system as dynamic responses. The shocks in different variables seem to be independent. In fact, the FIML residual correlations matrix is nearly diagonal (there seem to be some simultaneous effects between \( \Delta(p - q) \) and \( \Delta h \)). Similar results (graphs) are obtained carrying out impulse response analysis based on the Choleski decomposition of the residual covariance matrix. For impulse response analysis with cointegrated systems see Lütkepohl and Reimers (1992).
5. POLICY IMPLICATIONS

From the policy perspective there are two important questions that arise from the empirical evidence produced in this paper. First, our results suggest that fluctuations due to intersectoral job allocations are important for the NAIRU. However, stylized versions of the textbook approach to the NAIRU, whereby unemployment must move to equilibrate the demands of employers and the bargaining real wage (as in the battle of the mark-ups framework) are strongly deceptive. NAIRU estimates, obtained by using the conventional reduced form (or Phillips curve) models, are incorrect since the natural rate of unemployment cannot considered as a fixed point, calling for the ceteris paribus assumption for all the data involved. This implies that the NAIRU estimates are unable to provide valuable information on the labour market status. In fact, these models do not perform any analysis of the joint distribution of the set of random variables involved (that is, an adequate description of the data generation process). As a result, point estimates vary greatly (see Staiger, Stock, & Watson, 1996) and the NAIRU indicator cannot be considered a useful concept in policy formulation. Thus, the empirical evidence based on these solutions are seriously misspecified and the policy implications erroneous.

Many authors (see also OECD, 1995) incorrectly argue that the identification problem is answerable for the lack of precision in NAIRU estimates. What is really wrong in these models is the
fact that the reduced form (the statistical model that comprises the probability model and the sampling model and that forms the basis of the parametric inference) is not estimated explicitly (see, Chiarini & Piselli, 2000; Spanos, 1989, 1990). We have shown in the paper that, considering a VAR representation as form of the joint data density function, the use of multivariate cointegration approach provides a new role for the identification problem, leading to a greatly complicated and enriched NAIRU model. The simplified version of the geometrical representation of the NAIRU reported in Figs. A1 and A2, shows that the NAIRU is not a fixed point but a solution in a stationary space. The MA representation (Eq. (5)) of the VAR model emphasizes how a random shock to an equation of the model produces random short-run effects (via \( C(L)\epsilon_t \)) and long-run effects (via \( C(1)\epsilon_t \)), engendering a new position in the space spanned by the attractor set. A shock to one variable implies a shock to all the variables in the long run: in this context, the ceteris paribus assumption is not allowed, the only way of analyzing the economic relationships remains the impulse response functions. This framework brings out a key policy point: claims such as,

in the long-run, unemployment is determined entirely by long-run supply factors and equals the NAIRU; in the short-run, unemployment is determined by the interaction of aggregate demand and short-run aggregate supply (Layard et al., 1991)

are misleading.

The position of the price–wage curve and of the equilibrium wage curve is allowed to shift in response to changes in the variables contained in vectors \( X_t \) of the bargaining model (6) and (7). The basic point here is that when government is included (e.g., real tax revenue or the ratio of public spending on welfare programmes to GDP, see Layard et al., 1991; Phelps & Zoega, 1998), a change in the government’s claim cannot shift (ceteris paribus) the bargained real wage or the price-determined real wage curve to achieve a new equilibrium rate of unemployment. In fact, we cannot hold \((n - 1)\) independent variables constant. The variables are not independent of one another both in the short and long run, so that each variable cannot vary by itself without affecting the others. What constitutes a statistical or a theoretical assumption is not apparent in the conventional NAIRU approach. This provides us with an erroneous framework for policy analysis.
As an example, we show in Fig. 5 a result that emerges from our study: the striking effect of the price wedge changes on dispersion measure and, therefore, on unemployment. An increase of a one-time (one standard deviation) impulse in the price wedge is seen to have a lasting effect on sectoral mismatch and unemployment. Although unemployment changes taper off to zero quite rapidly, the unemployment level does not die out asymptotically. Real wage and, therefore, unemployment responses will result in lower equilibrium values.

The second issue raises the question of which policy should be attempted. The results suggest that the sectoral distribution of employment changes is important for both the short and the long-run unemployment performance. The picture that one should have in mind is that the continuous process of disturbances and sectoral job reallocation produce unemployment persistence. While it can be argued that macroeconomics policies have an important role to play in attenuating negative disturbances to the economy (we emphasize again that an uncritical bargaining framework is problematic for this purpose), it appears reasonable to state that structural policies have a central role in acting on those factors that produce and enhance mismatch problems in the economy.

Because of industry-specific skills and a labour market situation characterized by institutional rigidities, with a plethora of regulations and a climate of resistance to change (see the literature quoted in Section 3), the process of labour reallocation between different sectors of industry involves temporary mismatches but also provides an important source of the persistence of unemployment.

All this puts a premium on policies that improve the functioning and flexibility of the labour market. Recommendations in these areas have been mostly encouraged by the OECD Secretariat (see OECD, 1995, 1997) in the context of reducing unemployment. To this end, a wide range of micropolicies (measures to enhance labour mobility and part-time work and measures that aim to reforming existing regulations and restrictive work practices; retraining programmes; favourable fiscal treatment for education and training; housing policy and measures that provide a better information on the labour market based on the notification of vacancies through efficient work agencies) should be regarded as crucial for macroeconomic performance. These policies may reduce the fraction of the “temporary” (or “frictional”) unemployment arising as a result of sectoral shifts, which translates into structural unemployment.
Moreover, the impulse response functions in Fig. 4 show that not only should considerable efforts be made in employment and redundancy legislation areas and working-time arrangements, but also changes in wage-setting procedures are required to accommodate structural unemployment. Mismatching of jobs and workers may raise both vacancies and unemployment. However, if wages rise quickly in expanding demand sectors there is no reason to believe that the slow (if any) adjustment to intersectoral shift of labour demand will lead to firms with more capital-intensive.

The impulse responses suggest that a reallocation distribution variable plays a significant role in short-run adjustment and long-run consequences of unemployment to wage and price shocks: wage policy agreements between both sides of industry and government may be useful to moderate wage increases in the context of a growth policy based on investment, employment and labour market programmes.

6. CONCLUDING REMARKS

The result that is more interesting and more worthy of comment is that permanent shifts in the composition of employment lead to permanent (in contrast with transitory) increases in unemployment. Permanent sectoral shifts are significant determinants of unemployment. Given the real wage process, our results make extreme Lilien’s view that sectoral shifts are a prime factor generating long lasting (rather than cyclical fluctuations) effects in joblessness. This result seems to contradict other analysis carried out for Italian unemployment that emphasizes that the potential explanatory power of mismatch and structural imbalances by industry, relative to unemployment, is very modest.10

Finally, from the statistical point of view, one of the main effects of including $\delta_t$ in the data set is to cancel the need to utilize dummy variables in the model (see the model discussed in Chiarini & Piselli, 1997). This is not surprising since $\delta_t$ accounts for the structural change effects in the period. All this is obtained with better stochastic properties.

APPENDIX A:

The attractor set $\beta_1$ (the space within which the cointegrated variables move) is a hyperplane with dimension $n - r = 4$ (six

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10 A variety of issues related to the measurement of mismatch are reported in the papers in Padoa Schioppa (1991). See also Elmeskov (1993).
variables and two cointegration vectors). For the purposes of this exposition, a geometrical representation may be provided considering only three variables, say, \((w - p), \delta\) and \(U\). In this \(\mathbb{R}^3\) space, we have the following stationary relations: \((w - p) = -0.015\delta\) and \((w - p) = -0.06U\), that is: \(\beta_1 = (1, 0.015, 0)\) and \(\beta_2 = (1, 0, 0.06)\). It is important to stress that the variables labour productivity, price wedge and worked hours are considered only for expositional purposes: the example that is proposed for an attractor for three variables \([(w - p), \delta, U]\) is represented in Fig. A1. To obtain the attractor space \(\beta_\perp\) we have to solve the linear system:

\[
\begin{pmatrix}
1 & 0.015 & 0 \\
1 & 0 & 0.06
\end{pmatrix}
\begin{pmatrix}
x_1 \\
x_2 \\
x_3
\end{pmatrix} = 0.
\]

This system (three variables and two equations) has \(\infty\) solutions. Consequently, one can achieve only a parametric solution. Fig. A2 shows that in our simplified system, the attractor space is a straight line: the equations describe orthogonal planes, respectively, to \(\beta_1\) and \(\beta_2\) and their intersection is a line. The points along this line are vertex of all vectors with coordinates \((\lambda, -66.7\lambda, -6.7\lambda)\). In this case, the process \(x = [(w - p), \delta, U]\) is pushed along the attractor space spanned by \(\beta_\perp\) by the common trends (the economic or driving forces) \(\alpha'_i \sum_{i=1}^t \varepsilon_i\). Johansen (1995) emphasizes that the long-run relations should be considered as stable links, statistically relevant within the sample, as described by the statistical model.
Figure A2. Attractor Set in $R^3$.

REFERENCES


