Regional Disparities and the Italian NAIRU[1]

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Abstract

In this paper we estimate the Italian NAIRU using annual data for the period 1951-1996. We find evidence consistent with aggregate wage setting in Italy depending only on the rate of unemployment prevailing in the Northern and Central areas of the country. There is evidence supporting the presence of a long-run cointegrating relationship among unemployment in the Northern and Central areas, the tax wedge, the real interest rate and a measure of union power. The response of unemployment to exogenous shocks is sluggish, suggesting that persistence is an important feature of the Italian labor market.

1 Introduction

Unemployment dynamics are often explained as the combined outcome of short-run and equilibrium unemployment dynamics. Defining equilibrium unemployment as the rate of unemployment that is consistent with constant inflation (NAIRU), the common view is that this rate can vary over time if exogenous variables affecting wage and price setting behavior vary (see Layard *et al.,* 1991).

Short-run adjustment interacts with the NAIRU in at least two ways. First, with heterogeneous agents and nonlinearities, equilibrium needs not be unique. When there are multiple equilibria, demand management policies affect the location of the NAIRU and equilibrium unemployment depends on the path followed by actual unemployment (see the papers in Cross, 1995). Second, in the presence of exogenous growing variables, the long run equilibrium rate depends on the lagged adjustment process (see Karanassou and Snower, 1997a).

In the literature, there are different approaches to estimating the time-varying NAIRU. One approach consists of adding to a standard Phillips curve augmented with supply side variables an additional equation that models the NAIRU as a time varying variable (see Gordon, 1997; Staiger *et al.* 1997). Another approach models labor market equilibrium and obtains the NAIRU as the unemployment rate that reconciles aggregate wage and price setting behavior when expectations are fully realized (see Layard *et al.* 1991).

In this paper, we follow the latter approach to study Italian unemployment and to provide an estimate of the Italian NAIRU for the period 1954-94. As emphasized in the title, we believe that a critical dimension of Italian unemployment, at least compared to other experiences, is the presence of persistent regional disparities between the industrialized North and Central areas and the under-developed South.

To preview our main results, we find evidence that aggregate wage setting in Italy depends only on the rate of unemployment prevailing in the Northern and Central areas of the country. Increases in Southern unemployment do not seem to affect aggregate wage pressure. This result is reminiscent of the distinction between short-term and long-term unemployment often found in the literature on unemployment in Europe (see Layard *et al.* 1991) and has two relevant implications. First, a reduction in the rate of unemployment in the South, given unemployment in the rest of the country, does not increase wage pressure and the rate of inflation. Second, aggregate unemployment in the long run can be reduced by a re-distribution of unemployment among regions, because temporary increases in the unemployment rate in the Northern and Central areas is absorbed over time by falling wage and price inflation.

There is also evidence supporting the presence of a long-run cointegrating relationship among unemployment in the Northern and Central areas, the tax wedge, the real interest rate and a measure of union power. As it happens in other studies (see e.g. Staiger *et al.*, 1997), however, our estimates of the NAIRU are not very precise and the width of the 95% confidence interval is close to 2 percentage points. The response of unemployment to exogenous shocks is sluggish, suggesting that persistence is an important feature of the Italian labor market.

Higher taxes, higher real interest rates and higher union power are associated to a higher NAIRU. According to our simulations, a permanent reduction in the tax wedge from its 1996 value to the value prevailing in the early 80s reduces the unemployment rate in the NC areas by 21.7% and the unemployment rate in the South by 6.2%. Hence, a permanent cut in the tax wedge is more effective in cutting unemployment in the NC areas than in the SO areas. While aggregate unemployment declines, regional disparities increase.

The paper is organized as follows. Section 2 presents the main stylized facts on Italian unemployment and labor market institutions. Section 3 introduces the theoretical model. Section 4 presents our empirical findings and the discussion. Conclusions follow.

2 Italian Unemployment: Stylized Facts

The Italian official unemployment rate was 12.8 percent of the labor force in the summer of 1997, about the same level reached in France (12.5%) and lower only than in Belgium (14.1%) and Spain (20.8%) within the European Union. With the exception of the late eighties, the rate has increased steadily from 1975. Time series analysis of unemployment in Italy is somewhat complicated by occasional changes in the definition of unemployment and in the design of the quarterly labor force survey.

Up to the major revision of October 1992, for instance, the official definition included among the unemployed jobless individuals who had either looked for a job in the previous six months, or were enrolled in the local public employment offices or finally had been listed to participate in a competition for public employment. This wider definition adds about three percentage points to the official unemployment rate.[1] Since the gap between the two measures exhibits a mild positive trend, there is evidence that discourage-

[1] See Casavola and Sestito (1994) for further details and Faini *et al.* (1996) for an interesting discussion.

ment could have increased over the years.[2]

Murphy and Topel (1997) argue that official unemployment is not necessarily the best measure of joblessness due to economic conditions and suggest nonemployment (unemployment plus out of the labor force) as a better measure. Following this line, we prefer to use in this paper the wider definition of unemployment, that includes a share of those out of the labor force for economic reasons.

While the official unemployment rate is likely to understate joblessness, both the official and the wider measure could overstate it by treating individuals with irregular jobs who declare to be unemployed and looking for a job as unemployed. Given the relative importance of the irregular economy in Italy, the upward bias could be significant. According to a recent study by CENSIS (1996), official unemployment in 1995 declines from 12 to 10.2 percent if we exclude from the unemployed 410 thousand individuals who declare to be unemployed but hold an irregular job. The official rate falls further to 8.9% if we include in the labor force about 3 million individuals who are classified as out of the labor force but are actually working. While this study suggests that the official rate could be about 25 percent higher than the effective unemployment rate in 1995, we have no information on whether and how this percentage has changed over time.

Figure 1 shows both the average unemployment rate for the period 1951-1996 (u) and the unemployment rates in the Northern and Central (u_{NC}) and in Southern Italy (u_{SO}).[3] While the regional rates were more or less equal in the mid-sixties, unemployment in the South increased faster than unemployment in the North from the late sixties to the early seventies and from the mid-eighties onwards. In 1965, unemployment in the two areas was respectively 6 and 5.6 percent; in 1996, unemployment in the South was equal to 28.8 percent, about three times as large as unemployment in the rest of the country, that reached 9.5 percent.

The gap between unemployment rates in the South and in the rest of the country is particularly severe among younger cohorts (age group: 15-24) and relatively low among individuals older than 35. In 1995, this gap was equal to about 35 percentage points in the former group and to about 5 percentage points in the latter group (see Faini *et al.*, 1996). Figure 2 relates unemployment mismatch, defined as $mm = \log(u) - \log(u_{NC})$, to relative labor costs (Northern gross wages over Southern gross wages) and to migration outflows from the South during the years 1960-90. The overall

[2] See the discussion in Alogoskoufis *et al.* (1995).

[3] The unemployment series that we use in this paper have been re-evaluated by the Bank of Italy in order to avoid the problems of change of definitions discussed in the main text. We would like to thank Paolo Sestito for having provided the data.

Figure 1: unemployment rates 1951-96

picture is that of an increasing trend in mismatch and of a decreasing trend both in population outflows from the South and in relative labor costs in the North.

There is a growing consensus among Italian economists that the reduction in relative wage differentials, together with substantial government transfers to Southern households and with the widespread presence of the irregular economy, has reduced the individual incentive to migrate from the South to the North.[4] Another important factor is public employment offered by the national and local government. Household support and the significant probability of finding a lifetime job with the government might have induced young outsiders with limited previous labor market experience, the bulk of unemployment in the South, to prefer wait unemployment to migration.[5]

Persistent regional disparities in unemployment could have important implications for the Italian Nairu. To see why, define the Nairu as the unemployment rate that reconciles wage setting and price setting in the steady state[6] and assume that the price setting equation is not sensitive to

[4] See Attanasio and Padoa Schioppa (1991) and Bodo and Sestito (1991). Faini (1997) is a recent review of the main issues.

[5] See Attanasio and Padoa Schioppa (1991) and Brunello (1992).

[6] See Layard *et al.* (1991).

Figure 2: regional data 1960-90

unemployment.[7] If unemployment in the South does not affect average wage pressure in the private sector, both because wage setting is dominated by insiders employed in the industrialized North and because of limited competition by Southern unemployed, the relevant Nairu is defined with reference to unemployment in the Northern and Central areas of the country. In such a case, variations in Southern unemployment, given unemployment elsewhere, have no significant impact on the Nairu. As long as unemployment increases only in the South, there is no pressure for aggregate wages to fall and for inflation to decline.[8]

Italian unions are a key actor in wage determination. Union membership has declined significantly since the end of the war from a peak of about 55 percent of total employment to the current 30 percent (see Figure 3, panel a). The decline was temporarily interrupted and even reversed in the late sixties and early seventies, an exceptional period of union power and influence on Italian economic and social life. While membership is not high by European standards, unions wield significant power, partly because of the coverage of

[7]See Blanchard and Katz (1997) and Nickell (1997).

[8]Empirical evidence in Bodo and Sestito (1991) and in Casavola *et al.* (1995) supports the view that unemployment in the South has limited impact on aggregate wage pressure in the Italian private sector. See also the review of the empirical literature presented in Lucifora and Origo (1997).

Figure 3: membership, taxes, public employment and relative wages

collective contracts, that extend to non-union members.

The Italian bargaining structure can be broadly described as a three-tier system, with some overlap between the different levels of wage setting. Price indexation clauses, income policy and welfare benefits are negotiated at the national level. Wage increases are bargained at the sectorial level and wage drift, partially bargained between local parties, takes place at the local level. Local bargaining is more widespread in large firms and in firms operating in the Northern and Central areas of the country.[9]

The active presence of the government in the Italian labor market has steadily increased from the early fifties and public employment has risen from less than 10 to close to 20 percent of the Italian labor force (see Figure 3, panel b). The regional distribution of public employment is not homogeneous, however, and the ratio of public to private employees is much higher in the South than in the North.[10] The relevant presence of the State as employer in the South has been interpreted as an obstacle both to wage flexibility and to outward migration, with negative implication on the regional rate of unemployment.

[9]Corneo and Lucifora (1997) show that the probability of local bargaining decreases when the firm is small or localized in the South. See also Sestito (1995).

[10]See Attanasio and Padoa Schioppa (1991), Figure 6.11.

The increase in the share of public employment has been accompanied by a steady increase in the tax wedge, inclusive of social security contributions paid by employers and employees. Measured as percentage of the gross wage, the wedge has risen from about 25 percent in the early fifties to more than 50 percent in 1996 (see Figure 3, panel c).[11] While public employment has steadily increased during most of the sample period, relative wages (public versus private net wages) have fallen from the early fifties to the mid-seventies and partially bounced back afterwards (see Figure 3, panel d).

Compared to other European countries, the replacement ratio between unemployment benefits and average wages has been very low for most of the sample period and has increased only recently. Since 1991, there are two types of benefits, ordinary benefits and mobility benefits. From 1974 to 1987, ordinary benefits consisted of a constant daily amount and could be drawn for at most 6 months. After a sentence by the Supreme Court in 1988, benefits became proportional to previous earnings and the proportion increased from the original 7.5 to the current 35 percent.[12] From 1991, employees made redundant by large firms are enrolled in regional mobility lists and draw substantially higher benefits for a period ranging from one to three years.[13] In 1993, the percentage of the unemployed with previous labor market experience enrolled in these lists was close to 20 percent, and only a fraction of them drew benefits. Given this heterogeneity of treatments, it is quite difficult to construct a reasonable estimate of the replacement ratio. Since this ratio has been very small during most of the sample, we shall ignore it in the rest of the analysis.

Figure 4 plots the inflation rate, measured as the rate of change in the consumer price index. This rate has been under 5 percent from the early 50s to the late 60s and close to 5 percent from the second part of the 80s to the end of the sample period. In the interval between these two periods of moderate inflation, the rate of change of the consumer price index jumped to about 20 percent in the early 70s and was still close to that level in the late 70s. Hence, the sample includes a period of sustained increase of inflation (from 1972 to 1974) and a period of rapid decline (from 1980 to 1985). According to Ball (1996), both the size and the duration of disinflation in the early 80s were high in Italy by international standards. Larger and slower disinflations

[11]See Padoa Schioppa (1990) for an analysis of the tax wedge in Italy. Brunello (1996) presents a discussion of the relationship between wages and employment in the private and in the public sector.

[12]See Carinci *et al.* (1992).

[13]See Brunello and Miniaci (1997) for details on Italian mobility lists. According to an estimate of the Dutch Planning Bureau (1995), the average replacement ratio for workers in these lists is about 64 percent.

Figure 4: the inflation rate

are expected to produce larger cyclical increases in unemployment. With hysteresis, these increases could raise the NAIRU.

Inflation dynamics is correlated both with the dynamics of the price of imported materials (Figure 5, panel a) and with the short-term real interest rate (Figure 5, panel b). With the exception of the early 60s, the real interest rate was positive and mildly declining from the early 50s to the early 70s, negative during most of the 70s and positive again during the high interest rate period in the early 80s,[14] when it reached a plateau close to 5 percent.

Finally, Figure 6 plots the inflation rate both versus average unemployment (panel a) and versus unemployment in the North and Central areas (panel b). Ignoring short-run loops, the locus of the relationship shifted up significantly in the early 70s and again in the late 70s, more or less in line with the two supply shocks. The negative relationship between inflation and unemployment is particularly clear during the disinflation of the early 80s. From the mi-80s onwards, inflation remained more or less constant while unemployment significantly increased. Inflation in 1987 was 5.4 percent, aggregate unemployment was 12.8 percent and regional rates where respectively 8.8 and 21.1 percent; in 1995, inflation was 5.7 percent and aggregate unemployment was up to 15.8 (+23.4 percent). The bulk of this

[14]See Fitoussi and Phelps (1990) for a discussion of the high real interest rate period.

Figure 5: real price of imports and real interest rate

increase, however, was registered in the South, where unemployment rose by more than 30 percent, compared to "only" 11.3 percent in the North.

3 Labor Market Equilibrium and the Nairu

In this Section, we present a simple labor market model of the Italian NAIRU. This will be used as a guideline in the empirical analysis. Given that our focus is on the long run, we consider mainly the long run equilibrium and explicitly exclude from most of the discussion both the short-term dynamics and the characterization of how the economy behaves when it is out of equilibrium.[15]

Consider an open economy populated by a given number of identical firms, that operate either in the Northern and Central areas (NC from now on) or in the Southern areas (SO from now on). Let N and S be the number of firms located respectively in the NC and in the SO regions.

Each firm shares the following Cobb-Douglas technology

$$Y_i = L_i^a K_i^{1-a} \tag{1}$$

where Y is output, L is labor, K is capital and $i = 1, \ldots, M$, where M is the total number of firms. Following Layard *et al.* (1991), the demand curve

[15] See Nickell (1997) for a similar strategy.

Figure 6: inflation-unemployment trade-off

faced by each firm is

$$\frac{P_i}{P} = \left[\frac{Y_i}{Y}\right]^{-(1-1/\mu)} \tag{2}$$

where P_i is the price set by firm i, P is the average price, Y is average output and μ is the price mark-up. According to Phelps (1994), in a "customer market" the price mark-up is likely to increase with the real interest rate r, because of the reduced incentive to build up market share. In an open economy, μ falls when international competition increases, either because of structural reform that increases openness or because of an appreciation of the real exchange rate q.[16] Finally, the price mark-up can vary with the level of economic activity and with the overall unemployment rate u.

Each firm sets employment (and prices) by taking the relevant wage as given. The capital stock is fixed at the beginning of each period and there is no factor mobility between the NC and the SO regions. Profit maximization yields

$$L_i = K_i \left[\frac{\mu\,(r,q,u)}{a}\frac{W_i}{P_i}\right]^{-1/(1-a)}. \tag{3}$$

The national labor market has the following stylized features. First, unions influence wage decisions at the national level by bargaining over the

[16]See the discussion in Wes (1996).

"tariff wage" W_c. Second, there is wage drift and local wages are set by local bargains between firms and local unions. Let W_{co} be the wage inclusive both of the tariff wage and of wage drift.[17] Third, local bargaining occurs only in the covered sector. We capture these features by adopting a two-step bargaining process. In the first step, the national union sets the tariff wage by taking employment outcomes into account.[18] In the second step, local unions and firms bargain over the drift only in the covered sector and the tariff wage prevails in the uncovered sector.

In the Nash symmetric equilibrium, each firm sells at the same price in the national market. With equal price mark-ups, different wages in the covered and uncovered sectors imply that firms must have different capital-labor ratios to satisfy the following condition

$$\frac{W_{co}}{x_{co}^{1-a}} = \frac{W_c}{x_{un}^{1-a}} \tag{4}$$

where $x = K/L$ and the subscripts *co* and *un* refer to the covered and the uncovered sector. Since the wage inclusive of drift is a mark-up of the tariff wage, σW_c, where $\sigma > 1$, the capital-labor ratio must be higher in firms operating in the covered sector.

3.1 Wage Drift

Corneo and Lucifora (1997) find that local bargaining in Italy occurs mainly in medium and large firms operating in the NC areas of the country. Based upon this evidence, we shall assume that the covered sector is located in the NC regions and the uncovered sector in the SO regions. While this is clearly a simplification, we believe that it captures in a stylized way an important feature of wage determination in Italy.

Notice first that the Cobb Douglas technology implies that

$$\frac{\Pi_i}{W_i L_i} = \frac{\mu - a}{a} = \gamma. \tag{5}$$

Defining P^e as the expected average price, the wage drift is set in firm i operating in the covered sector by maximizing the following objective function

$$\max_{W_{co}} \left\{ \left[\frac{L_i \left(W_{co} - W_c \right)}{P^e} \right]^{\beta} \Pi_i^{1-\beta} \right\} \tag{6}$$

[17] We ignore here the part of the wage drift that is determined unilaterally by the firm. Ferner and Hyman (1992) is a detailed discussion of the structure of bargaining in Italy. Lucifora (1991), Brunello (1994) and Ordine (1994) (1996) study wage drift in Italy.

[18] This is clearly a simplification. In practice, national bargains involve mainly indexation rules, pension schemes, rules of the game and economic policy. Wages are set at the sectorial level.

where β is the relative bargaining power of the local union. In the local bargain, unions try to increase the (real) wage relative to the tariff wage. The latter is a reasonable proxy of the income received by employees during the bargain.

Both parties in the bargain take aggregate prices as given. Letting $P^e = P$ and defining λ as the absolute value of the wage elasticity of labor demand, we obtain

$$W_{co} = \frac{1 - \beta + \beta\gamma\lambda}{1 - \beta + \beta\gamma\lambda - \beta\gamma} W_c. \tag{7}$$

An increase in local unemployment u_{NC} reduces the relative bargaining power of the local union if union members either face a higher risk of losing their job or have a lower probability of finding a new job in the event that they lose the current one. In this case, $\beta = \beta(u_{NC})$ and the wage in the covered sector is

$$W_{co} = \sigma(u_{NC}) W_c. \tag{8}$$

Hence, the mark-up of the covered wage over the tariff wage, that measures local wage drift, declines when the local unemployment rate increases[19].

3.2 Central Wage Setting

Total employment in the economy is equal to

$$Z = NL_{co} + SL_{un}. \tag{9}$$

We characterize wage setting in the private sector by assuming that the central union minimizes a loss function that depends on wages and employment. We assume that the central union solves the following problem

$$\min_{\ln \frac{W_c}{P}} \left\{ \frac{1}{2} \left[\ln \frac{W_c}{P} - \ln A \right]^2 + \frac{\xi}{2} \left[\ln E - \ln \overline{E} \right]^2 \right\} \tag{10}$$

where A is the real wage target[20] and E and $\overline{E}$ are respectively the values of actual and target employment that affect union preferences. If the union cares about total employment, $E = Z$ and $\overline{E}$ is the labor force net of public employment. If it cares only about employment in the North and Central Italy, $E = NL_{co}$ and $\overline{E}$ is the labor force located in the same area. Finally, if it cares differently about employment in the South and in the North, these variables will enter with different weights.

[19]Ordine (1996) finds that wage drift in Italy varies negatively with the rate of unemployment.

[20]See Alogoskoufis and Manning (1987).

Notice that the central union sets the wage in the private sector. Assuming that $\overline{E}$ is the labor force in the private sector, F_p, and recalling that $u = (F - N_G - Z)/F$, where F is the total labor force and N_G is public employment, we have that

$$\ln F - \ln Z - s_G = u = \ln F_p - \ln Z \tag{11}$$

where $s_G = N_G/F$.

Next, the target wage A increases in the alternative wage $\overline{W}$, that measures the wage available outside the private sector. Ignoring self-employment, alternatives outside the private sector are unemployment and public employment. Hence,

$$\overline{W} = V_u \left(\frac{u}{u + s_G}\right) + \frac{W_G}{P}\left(\frac{s_G}{u + s_G}\right) \tag{12}$$

where W_G is the public sector wage and V_u is real income from unemployment. With the exception of the most recent period, unemployment benefits in Italy have been very low by international standards[21] and can be ignored in a first approximation. An implication of equation (12) is that the target wage is increasing both in the public sector wage and in public employment.[22]

The target wage A is also affected by real labor productivity, x, and by the price mark-up μ, that are both related to the size of the rents shared by the parties, and by union strength UP. This wage is also higher the higher is either skill or regional mismatch. As argued by Nickell (1997), if most of the vacancies are in the North and most of the unemployed in the South, as it happens in Italy, this has a significant impact on wage pressure for any level of unemployment. It follows that migrations from the South to the North, by reducing the mismatch between unemployment and vacancies, could reduce the target wage. In summary, we have that

$$A = A\left[u, mm, \frac{W_G}{P}, s_G, x, \mu, UP\right]. \tag{13}$$

Suppose that the central union cares about total (private) employment and assume that the impact of the tariff wage on the mark-up σ is so small that it can be ignored in a first approximation. This is equivalent to assuming that changes in the average wage do not affect wage differentials. Suppose

[21] See Emerson (1990).

[22] Public employment could affect wage pressure through alternative routes. A well known example is discussed in Calmfors and Horn (1985): if government policy supports employment by compensating reductions of employment in the private sector with employment creation in the public sector, the perceived trade-off between wage gains and employment losses faced by the central union improves and wage pressure increases.

also that the central union does not internalize the potential effects of the tariff wage on regional mismatch. Then the first order condition associated to equation (10) is

$$\ln W_c - \ln P - \ln A + \frac{\xi\lambda}{1+\nu\lambda}u = 0. \tag{14}$$

where $\nu = -\partial \ln A/\partial u > 0$. Given that wage drift occurs only in the covered sector, the average real wage in the private sector, W, is defined by

$$\frac{W}{P} = \left[\frac{W_{co}}{P}\right]^{\rho}\left[\frac{W_c}{P}\right]^{1-\rho} = \frac{W_c}{P}\sigma^{\rho}. \tag{15}$$

where ρ is a constant weight. Letting $W/W_G = \theta$ and using the definition of mismatch to eliminate u_{NC}, equation (15) can be re-written in implicit form as follows

$$\frac{W}{P} = \frac{W}{P}\left[u, mm, \theta, s_G, x, r, q, UP\right]. \tag{16}$$

Notice that the real wage is gross of taxes and does not directly depend on the tax wedge. This is because we have assumed from the start that there is no long-run real wage resistance, in line with most of the theoretical and empirical literature.[23] Even without real wage resistance, however, real gross wages and the tax wedge could be related via the government budget constraint. Suppose that this constraint is given by

$$W_G s_G (1-\tau) + bu = \tau W (1 - u - s_G) \tag{17}$$

where b are unemployment benefits and τ is the tax wedge. It is reasonable to expect that this constraint be satisfied in the long run equilibrium, since the share of public employment on the labor force cannot be constantly risen, as it has happened in Italy over the post-war period, without a corresponding increase in taxation (see Figure 3 above). Ignoring unemployment benefits, the budget constraint implies the following relationship between s_G and τ

$$s_G = \frac{\tau(1-u)}{\tau + (1-\tau)/\theta} \tag{18}$$

where $\partial s_G/\partial\tau \mid_u > 0$.

In the log run equilibrium, we also expect that the current account deficit (surplus) accumulated in the current period be matched by a discounted current account surplus (deficit) in the future. It is standard to assume that

[23] The same argument applies to the relative price of imports. See Layard *et al.* (1991) and Daveri and Tabellini (1997) for a recent different view.

the current account as a proportion of real GNP is a negative function of the real exchange rate, q. Taking into account the inter-temporal nature of the constraint, we obtain an additional equation that relates the real exchange rate q to the real rate of interest r.[24] This equation allows us to eliminate the real exchange rate from equation (16).

To sum up, a log-linear specification of the average (real) wage in the private sector is given in the long run by

$$\ln W - \ln P = \alpha_o + \alpha_1 \ln x + \alpha_2 \ln \theta + \alpha_3 \tau + \alpha_4 r - \alpha_5 \ln u + \alpha_6 mm + \alpha_7 \ln UP. \tag{19}$$

3.3 Price setting

Let average productivity x be a geometric average of productivity in the covered and uncovered sectors

$$x = x_{co}^{\rho} x_{un}^{1-\rho} \tag{20}$$

Using equation (4) and the fact that $W_{co} = \sigma W_c$, we can express productivity in the covered sector as $x_{co} = x\sigma^{(1-\rho)/(1-\alpha)}$.

It is useful to start from the fact that equation (4) can be written as

$$\frac{\sigma}{x_{co}^{1-a}} = \frac{1}{x_{un}^{1-a}} \tag{21}$$

that yields

$$\sigma = \left\{ \frac{x_{co}}{x_{un}} \right\}^{1-a} \tag{22}$$

and $x_{co} > x_{un}$. Next, in the symmetric equilibrium equation (3) for the covered sector can be written as

$$P = \frac{\mu\left(r, q, u\right) W_{co}}{a x_{co}^{1-a}} \tag{23}$$

Using the intertemporal external constraint and equations (15) and (20) to substitute out q, x_{co} and W_{co}, we obtain

$$P = \mu\left(r, u\right) W / ax \tag{24}$$

that can be re-written as

$$\frac{P}{W} = P\left[r, u, x\right]. \tag{25}$$

[24] See Abel and Bernanke (1992) for a discussion of the inter-temporal current account equilibrium condition.

3.4 Equilibrium

Following Layard *et al.* (1991), we define the long run equilibrium rate of unemployment, u^*, as the rate that makes wage pressure and price setting behavior consistent. In this paper, we use the terms "equilibrium rate of unemployment" and "NAIRU" interchangeably. After log-linearizing equation (25) as follows

$$\ln P - \ln W = \phi_0 - \phi_1 \ln u - \phi_2 \ln x + \phi_3 r \tag{26}$$

the equilibrium rate of unemployment is given by

$$\begin{aligned} \ln u^* \;\; = \;\; & \frac{\phi_0 + \alpha_0}{\phi_1 + \alpha_5} + \frac{\alpha_3}{\phi_1 + \alpha_5}\tau + \frac{\alpha_1 - \phi_2}{\phi_1 + \alpha_5}\ln x + \frac{\alpha_4 + \phi_3}{\phi_1 + \alpha_5} r \\ & + \frac{\alpha_7}{\phi_1 + \alpha_5}\ln UP + \frac{\alpha_6}{\phi_1 + \alpha_5} mm + \frac{\alpha_2}{\phi_1 + \alpha_5}\ln\theta. \end{aligned} \tag{27}$$

It turns out that labor productivity x has no effect on the NAIRU if $\alpha_1 - \phi_2 = 0$. Given the tremendous increase in productivity experienced by most developed countries in the long run, this condition has to be satisfied by any reasonable model of the NAIRU.[25] We impose it in the ensuing empirical analysis.

Specification (27) suggests that the NAIRU depends on the real interest rate. As discussed by Phelps (1994), this rate is affected by the interaction of the domestic demand and supply of goods. Thus demand shifts, such as changes in government policy, can affect the NAIRU if they affect the real interest rate. Moreover, since credit contracts are usually not indexed, the real rate of interest is negatively correlated with inflation. Hence, higher inflation can also affect the equilibrium rate of unemployment if it reduces the real rate of interest.[26]

3.5 Dynamics

Following Nickell (1997), short term dynamics can be introduced by augmenting both the price and the wage setting equation with price surprises and by adding the effects of unemployment changes in the wage setting equation. If price surprises are equivalent to inflation changes, the unemployment-inflation trade-off is given by

$$\ln u = \gamma_0 \ln u_{-1} + \gamma_1 S - \gamma_2 \Delta^2 P \tag{28}$$

[25] See Blanchard and Katz (1997).

[26] The impact of inflation on real interest rates and the failure of the "Fisher hypothesis" are discussed in Blanchard (1993).

where S is a vector that includes all the variables on the right hand side of (19) and $\Delta^2 P$ is the change in the rate of inflation.

In a recent series of papers, Snower and his associates[27] have argued that a combination of exogenous shocks and of sluggish dynamic adjustment can provide an alternative explanation of the long term movements of unemployment, that is qualitatively different from the standard view based upon the movement of the natural rate of unemployment, driven by shifts in the wage and price setting equation. Because of this, in the empirical part of the paper we shall also look at dynamic adjustment in the Italian labor market.

4 Empirical Analysis

4.1 Univariate Properties of the Italian Unemployment Rate

In this Section we study the univariate properties of the Italian unemployment rate for the period 1951-96. This investigation is useful at least for two reasons. First, it gives useful preliminary results that can be used in later stages of the analysis. Second, economically interesting stylized facts can usually be derived from univariate analyses.

According to Karanassou and Snower (1997a), the fact that the unemployment rate u is bounded from above by $\overline{u} = 1$ and from below by $\underline{u} = 0$, by construction rules out the possibility that $\{u_t\}$ can be $\mathsf{I}(1)$. From this argument it follows that unit root testing is superfluous or even misleading when applied to unemployment rates series. We argue here that this conclusion could be too hasty. To see why, let $\{x_t\}$ be the random walk $x_t = x_{t-1} + \varepsilon_t$, with $\varepsilon_t \sim \mathsf{IID}(0, \sigma^2)$ and $x_0 = 0$. It is fairly well known[28] that if T is large, x_T is of magnitude $O_p(\sqrt{T})$. In practice this means that the random walk $\{x_t\}$ can exhibit arbitrarily large excursions from the origin. Furthermore, it can be proved that for sufficiently large T

$$\mathsf{P}(x_T > M, x_{T+1} > M, \ldots) > 1 - \delta \tag{29}$$

for arbitrarily small $\delta > 0$ and arbitrarily large values of $M > 0$. Therefore,

$$\lim_{T \to \infty} \mathsf{P}(x_T \notin [-\underline{x}, \overline{x}]) = 1 \qquad \forall \underline{x}, \overline{x} > 0. \tag{30}$$

In particular, this means that a zero mean random walk will be greater than one or smaller than zero with certainty after a sufficiently large time interval

[27] See for instance Henry and Snower (1997) and Karanassou and Snower (1997a,b).

[28] A good account of random walk properties can be found in Cox and Miller (1965).

has elapsed. Assume, however, that $\{x_t\}$ behaves like a random walk with two reflecting barriers placed at $\underline{x} = -a$ and $\overline{x} = b$. The reflecting barriers bind the series within the interval $[-a, b]$. It can be shown that in this case $\{x_t\}$ is stationary, but the expected number of steps T_B required for the barriers to be binding is

$$\mathsf{E}(T_B) \approx \frac{|\underline{x}\overline{x}|}{\sigma_v^2}. \tag{31}$$

For $t < T_B$, $\{x_t\}$ behaves as an ordinary unrestricted random walk.

To illustrate the practical implications of this result with reference to the unemployment rate, assume now for simplicity that the actual Italian unemployment rate is generated by the random walk $u_t = u_{t-1} + \upsilon_t$ with $\upsilon_t \sim \mathsf{IID}(0, \sigma_\upsilon^2)$. Even if this is a simplification, it serves as an expository device and might be not too far from reality.[29] Once the random walk simplification is imposed, we can try to estimate $\mathsf{E}(T_B)$ for the unemployment series. In order to do so, we first demean the series and consider the reflecting barriers $\underline{u} = -\hat{\mu}_u$ and $\overline{u} = 1 - \hat{\mu}_u$, with $\hat{\mu}_u$ being the sample mean of the unemployment rate series. Then we use the sample variance of the first differences of the demeaned series as an estimate of σ_υ^2 in the expression for $\mathsf{E}(T_B)$. Our calculations show that we should expect one of the barriers to be binding after approximately 1,100 years![30] During this interval the unemployment rate is exactly equivalent to a standard unrestricted random walk. This might also explain (on a statistical, not on an economic ground) why we don't know of any country in any historical epoch where $u_t = 0$ or $u_t = 1$.

In our view, the bottom line is that there is no serious conflict between the common empirical evidence that the unemployment rates of many (if not most) countries appear to be nonstationary and the fact that these rates are bounded in the interval [0, 1] by construction. At the very least, this does not seem to be a real practical problem in the observed time series samples. Rather, the important point seems to us to be that most unit root tests may have low power in small samples.

[29] Using a Cramér-von Mises test for martingality we cannot reject the null that Italian unemployment rate behaves like a random walk. This test might, however, lack power in very small samples. For the theoretical foundations of this test, see Durlauf (1991). A Monte Carlo analysis of its properties is in Lupi (1997).

[30] An even longer period should be necessary for the series to manifest its stationarity. Of course, $\mathsf{E}(T_B)$ would be reduced if the admissible range for $\{u\}$ were shrinked. In particular, it would be approximately 260 years if the upper bound were placed at $\overline{u} = 0.30$. It would be reduced further to $\mathsf{E}(T_B) \approx 110$ for $\overline{u} = 0.30$ and $\overline{u} = 0.03$. However, while the bounds $\underline{u} = 0$ and $\overline{u} = 1$ derive from the definition of u itself, the barriers selected in this example are completely arbitrary and there is no special reason to choose them instead of others.

Having cleared a possible misunderstanding about the meaning of "I(1)" referred to the unemployment rate, we start our univariate analysis by testing for the order of integration of the Italian unemployment rate. We study both the national average unemployment rate, u_t, and two regional unemployment rates, $u_{NC,t}$ and $u_{SO,t}$, the former defined for the Italian Northern-Central regions, and the latter for the Italian "Mezzogiorno", in the Southern part of the country. Given the relative shortness of the sample, and in order to make the analysis as robust as possible, we apply four different unit root tests. In particular we use the Augmented Dickey Fuller ($ADF(\tau)$; Said and Dickey, 1984), the Covariate-Augmented Dickey Fuller ($CADF(\tau)$; Hansen, 1995), the Phillips and Perron (Z_α; Phillips and Perron, 1988), and the Weighted Symmetric ($WS(\tau)$; Pantula *et al.*, 1994) tests. All these tests are performed with a constant and a trend, and the results are reported in Table 1. For the $ADF(\tau)$ test, the number of lags is selected on the basis of the properties of the residuals, starting with five lags and simplifying the model whenever possible without inducing autocorrelation and heteroskedaticity. It turns out that no lags are necessary and standard DF tests can be carried out for all the variables under study. Since, however, symptoms of residuals non-normality are apparent in the $ADF(\tau)$ for u_S, due to the presence of an outlier in 1956, the corresponding $ADF(\tau)$ test is also computed with an impulse dummy variable.[31] This model augmentation does not involve changes in the asymptotic distribution theory of the test, but can affect its size in small samples (see Franses and Haldrup, 1994).

The $CADF(\tau)$ test is computed following essentially the same procedure used for the standard $ADF(\tau)$ test, with inflation and union power as covariates. This test is expected to be significantly more powerful than the standard univariate unit root tests in finite samples. Turning to the $WS(\tau)$ test, the number of lags is chosen using the *AIC2* criterion.[32] Finally, in the Z_α test, the same number of lags used for the $WS(\tau)$ test is selected.

In no instance the unit root null can be rejected, nor is even close to being rejected (see Table 1). Therefore, in spite of the sample being rather short, these results make us fairly confident that the I(1) representation is adequate for the Italian unemployment rate over the period considered in this paper.

Asymmetry in the unemployment rates series is tested using an extension

[31] It should be noted that the same outlier was also identified using Gómez and Maravall's (1997) procedure TRAMO for automatic outlier detection. (See also Gómez and Maravall, 1994).

[32] The *AIC2* criterion is $Nlag = j + 2$ where *Nlag* is the number of lags selected by *AIC2*, and j is the number of lags that minimizes Akaike's information criterion. The *AIC2* criterion should reduce size distorsion for the $WS(\tau)$ test. (Cfr. Pantula *et al.*, 1994).

Augmented Dickey Fuller tests				
Variable	t-value	$AR(1-2)$	$ARCH(1)$	$Norm.$
$\ln u$	-1.881 (-3.511)	1.011 [0.373]	0.172 [0.681]	2.907 [0.234]
$\ln u_{NC}$	-2.019 (-3.511)	1.760 [0.185]	0.134 [0.716]	0.034 [0.983]
$\ln u_{SO}$	-1.903 (-3.511)	0.389 [0.680]	0.167 [0.685]	24.181 [0.000]
$\ln u_{SO}(+d_{56})$	-2.834 (-3.511)	0.147 [0.864]	0.088 [0.768]	4.275 [0.118]
Covariate-Augmented Dickey Fuller tests				
	t-value	ρ^2	V	
$\ln u(+\Delta INF, \Delta \ln UP)$	-1.010 (-3.262)	0.794	0.840	
$\ln u_{NC}(+\Delta INF, \Delta \ln UP)$	-1.885 (-3.209)	0.735	0.756	
$\ln u_{SO}(+\Delta INF, \Delta \ln UP)$	-0.949 (-3.308)	0.863	1.020	
Phillips-Perron and Weighted Symmetric tests				
	Z_α	$WS(\tau)$		
$\ln u$	-3.844 [0.896]	-0.737 [0.988]		
$\ln u_{NC}$	-5.742 [0.767]	-1.184 [0.955]		
$\ln u_{SO}$	-4.479 [0.857]	-0.828 [0.984]		

5% critical values are reported in parenthesis for $ADF(\tau)$ and $CADF(\tau)$ (see MacKinnon, 1991; Hansen, 1995). $AR(1-2)$, $ARCH(1)$, and $Norm.$ denote the tests for autocorrelation up to order 2, ARCH, and normality of the $ADF(\tau)$ regression residuals (P-values in brackets). ρ^2 and V are long run (squared) correlation and variance ratio, respectively (see Hansen, 1995). d_{56} is an impulse dummy corresponding to the year 1956. ΔINF and $\Delta \ln UP$ are first differences of inflation and (log) union power. $(+z)$ in the "Variable" column indicates model augmentation with variable z. Z_α and $WS(\tau)$ are the Phillips-Perron and the Weighted Symmetric tests, respectively (P-values in brackets).

Table 1: unit root tests

	sk	P_{WN}	P_S	$N_{x>\mu}$	P_N
$\Delta \ln u$	-0.368	0.157	0.134	0.489	0.311
$\Delta \ln u_{NC}$	0.093	0.399	0.385	0.467	0.165
$\Delta \ln u_{SO}$	-0.447	0.110	0.087	0.578	0.065

sk is the coefficient of skewness of the observed series. P_{WN} and P_S are the P-value under the hypothesis that the series is a Gaussian white noise and the P-value based on 10,000 simulations, respectively (see Lupi and Ordine, 1997). $N_{x>\mu}$ is the fraction of observations greater than the mean: P_N is the corresponding tail probability.

Table 2: tests for symmetry

of the approach followed in Lupi and Ordine (1997). This is a simulation-based procedure, conceptually similar to that used by De Long and Summers (1986), but not dependent on any implicit null hypothesis. Being nonparametric in nature, contrary to LM-type tests, this test does not require a parametric alternative either (see *e.g.* Luukkonen *et al.*, 1988; Luukkonen and Teräsvirta, 1991). The test estimates both the tail probability of the empirical coefficient of skewness of the observed series and that related to the number of observations greater (or smaller) than the mean. The second version of the test is less sensitive to the presence of outliers. As shown in Table 2, the tests cannot reject the lack of significant asymmetries in all the three series under study.[33]

4.2 The long run

Before embarking in any modelling effort, we first check the order of integration of the variables used in the empirical analysis.[34] In this Section, the number of lags for the ADF(τ) test is chosen on the basis of the last signifi-

[33]The tests reject for P-values less than a prespecified confidence level $\alpha/2$. The coefficient of skewness sk of n IID Gaussian observations is distributed as $\sqrt{n}sk \sim \mathsf{N}(0, 6)$. (See *e.g.*, Kendall and Stuart, 1969.). The evidence gathered in this study is partly at odd with the results presented by Lupi and Ordine (1997), who use quarterly data for the period 1980-1996.

[34]Since we plan to apply multivariate cointegration analysis, the univariate unit root tests might seem redundant. Given, however, that unit root testing in the multivariate setting is conditional upon the estimated number of cointegrating vectors, and given that tests for the cointegration rank suffer of nonsimilarity problems (Nielsen, 1997), we prefer to have also an independent univariate piece of evidence.

Variable	$ADF(\tau)$	Z_α	$WS(\tau)$
τ	-1.242	-4.087	-1.478
	(-3.519)	$[0.882]$	$[0.898]$
$\Delta\tau$	-8.307	-49.776	-3.802
	(-3.522)	$[0.000]$	$[0.008]$
r	-2.035	-7.492	-2.187
	(-3.519)	$[0.625]$	$[0.513]$
Δr	-6.241	-43.053	-4.394
	(-3.522)	$[0.000]$	$[0.001]$
$\ln\theta$	-1.007	-2.516	-1.464
	(-3.519)	$[0.955]$	$[0.902]$
$\Delta\ln\theta$	-4.843	-33.492	-3.312
	(-3.522)	$[0.004]$	$[0.035]$
$\ln UP$	-2.044	-7.742	-2.434
	(-3.519)	$[0.604]$	$[0.335]$
$\Delta\ln UP$	-3.568	-18.526	-3.723
	(-3.516)	$[0.095]$	$[0.011]$
mm	-3.097	-11.019	-2.519
	(-3.519)	$[0.371]$	$[0.281]$
Δmm	-4.574	-21.399	-4.169
	(-3.522)	$[0.053]$	$[0.003]$

Augmented Dickey-Fuller ($ADF(\tau)$), Phillips-Perron (Z_α), and Weighted-Symmetric ($WS(\tau)$) unit root tests. 5% critical values in parenthesis (MacKinnon, 1991). P-values in brackets. The $ADF(\tau)$ test for $\ln(UP)$ includes an impulse dummy variable.

Table 3: unit root tests

cant lag (at the 10% confidence level).[35] The Phillips-Perron and Weighted Symmetric tests have been carried out as in Section 4.1. The results presented in Table 3 indicate that all the variables used in this study are well represented by $\mathsf{I}(1)$ processes.

In the model discussed in Section 3, equation (27) suggests that, in the long run steady state equilibrium, the unemployment rate that reconciles wage setting with price setting should be cointegrated with union power, relative wages in the private and public sectors, regional mismatch, the tax wedge and the real interest rate. An important issue is whether the unemployment rate prevailing in the Northern and Central areas has stronger effects on aggregate wage pressure than the unemployment rate in the South. Unemployment in the South is both characterized by longer duration and concentrates more among individuals looking for their first job.[36] If labor

[35]See Ng and Perron (1995) on this selection criterion.

[36]See for instance Bodo and Sestito (1991). Quarterly flow data show that the probability of remaining in the unemployment pool is significantly larger for individuals living in the South. Micro evidence is also presented in Ordine (1992).

market outsiders and the long term unemployed living in the South have limited impact on wage pressure, we should find that u_{NC} matters more than u_{SO} in equation (19).

When only the rate of unemployment in the Northern and Central areas affects wage pressure, the NAIRU is a particular value of this rate rather than of the aggregate rate. To illustrate, it is useful to re-write equations (19) and (26) as follows

$$\ln W - \ln P = \kappa_0 - \kappa_1 \ln u + \kappa_2 mm + \kappa_3' \mathbf{X} \tag{32}$$

$$\ln P - \ln W = \upsilon_0 + \upsilon_1' \mathbf{Y} \tag{33}$$

where $\mathbf{X}$ and $\mathbf{Y}$ are vectors of the other variables included in either equation. Notice that the price setting equation is assumed to be independent of the rate of unemployment. Empirical evidence suggests that this is a good approximation. Nickell (1997), for instance, argues that demand effects on price setting are small; Blanchard and Katz (1997) and Carlin and Soskice (1990) also use a flat price setting equation. Empirical evidence supporting this assumption in the Italian case is presented by Modigliani *et al.* (1986), Destefanis (1995) and Cristini (1995).

Solving for the equilibrium aggregate unemployment rate, we obtain

$$\ln u^* = \frac{\kappa_0 + \upsilon_0}{\kappa_1} + \frac{\kappa_3'}{\kappa_1}\mathbf{X} + \frac{\upsilon_1'}{\kappa_1}\mathbf{Y} + \frac{\kappa_2}{\kappa_1} mm \tag{34}$$

where u^* is equilibrium unemployment. When wage setting does not depend on mismatch but only on aggregated unemployment, $\kappa_2 = 0$ and the coefficient of mm in equation (34) is equal to zero. On the other hand, wage setting depends only on $\ln u_{NC}$ when $\kappa_2 = \kappa_1$. In this case, the coefficient of mm in (34) is equal to 1. Finally, only $\ln u_{SO}$ would be relevant if $\kappa_2/\kappa 1 = \varphi^*/(\varphi^* - 1)$, where φ^* is the long-run value of φ, with $u = u_{NC}^{\varphi} u_{SO}^{1-\varphi}$.

Based upon our realistic assumption about price setting behavior, the above discussion shows that we can gather indirect evidence on the relative importance of $\ln u_{NC}$ for aggregate wage setting without explicitly estimating an aggregate wage setting equation. Given that identification of such an equation is difficult, as discussed by Manning (1993) and Bean (1994), this is particularly appealing.

We start our empirical analysis by considering the following VAR system

$$\mathbf{\Pi}(B)\mathbf{Z}_t = \mathbf{\Phi}\mathbf{D}_t + \varepsilon_t \ , \qquad \varepsilon_t \sim \mathsf{NIID}(\mathbf{0}, \mathbf{\Omega}_\varepsilon) \tag{35}$$

where $\mathbf{\Pi}(B)$ is a k-order polynomial matrix in the backshift operator B, such that all the roots of $|\mathbf{\Pi}(z)| = 0$ are greater or equal to one in modulus; $\mathbf{Z}_t$

is a (6×1) vector that includes the variables in (27) with the exception of x as explained in Section 3.4, namely $\ln u$, $\ln UP$, r, $\ln \theta$, τ and mm; $\mathbf{D}_t$ is a vector of deterministic variables that includes the constant and a linear trend. We rule out the possibility of a deterministic quadratic trend in the levels and therefore we restrict the trend to be linear, as in model $\mathsf{H}^*(r)$ in Johansen (1994).

Dealing with a "semi-reduced" form like (35) has the merit of restricting the number of variables. Apart from the advantages in terms of efficiency, this makes it less likely that the bias problems described in Abadir *et al.* (1997) might arise. We find that a reasonable model is the simple VAR(1). The main model diagnostics are reported in Table 4. Relying on the results by Franses and Haldrup (1994), $\mathbf{D}_t$ contains an impulse dummy variable in order to correct for a large, isolated outlier.

Engle and Granger (1987) have shown that an isomorphism exists between the VAR (35) and the error correction mechanism (ECM)

$$\mathbf{\Pi}^*(B)\Delta\mathbf{Z}_t = \mathbf{\Phi}\mathbf{D}_t + \mathbf{\Pi}_0\mathbf{Z}_{t-1} + \varepsilon_t \tag{36}$$

with $\mathbf{\Pi}_0 = \alpha\beta'$. The model in the form (36) can be used to test for the presence of cointegration along the lines discussed in Johansen (1988) and Johansen (1995). As shown in Table 5 below, the tests for cointegration rank (not adjusted for the degrees of freedom)[37] indicate the presence of one cointegrating vector.

An interesting question is whether some of the variables included in the VAR can be considered as weakly exogenous with respect to the long run parameters. As shown in Table 6, it turns out that the real interest rate r, the tax wedge τ and the relative wage $\ln \theta$ are weakly exogenous for cointegration ranks $ra \leq 2$. Since we are concerned with an equilibrium concept, it is for us particularly important that the coefficient of the trend in the cointegrating vector be zero.[38] This restriction cannot be rejected by the data. Furthermore, using the weak-exogeneity constraints on α, and jointly testing for restrictions on β' along the lines exposed *e.g.* in Johansen (1995), we identify the following cointegrating vector (standard errors in parenthesis)

$$c_1 = \underset{(0.283)}{1.538}\,\tau + \underset{(0.595)}{2.216}\,r - \ln u + \underset{(0.110)}{0.715}\,\ln UP + mm. \tag{37}$$

[37]The role of degrees of freedom correction in the tests for cointegration rank is rather controversial (see *e.g.* Reimers, 1992 and Nielsen, 1997). However, Nielsen (1997) shows that they are arbitrary in general. Therefore we do not adjust for degrees of freedom but use rather tight criteria (5%) in order to minimize the risk of accepting "too many" cointegrating vectors, therefore incurring in non-similarity problems (Nielsen, 1997).

[38]For a discussion of the correspondence between equilibrium and cointegration, see Hatanaka (1996).

Equation	Test	Distribution	test value	P-value
τ	AR 1-2	F(2,34)	1.162	0.325
r	AR 1-2	F(2,34)	0.020	0.980
$\ln u$	AR 1-2	F(2,34)	0.688	0.509
$\ln \theta$	AR 1-2	F(2,34)	1.392	0.262
$\ln UP$	AR 1-2	F(2,34)	3.372	0.046
mm	AR 1-2	F(2,34)	2.007	0.150
VAR	AR 1-2	F(98,103)	1.340	0.083
τ	Normality	$\chi^2(2)$	2.031	0.362
r	Normality	$\chi^2(2)$	9.164	0.010
$\ln u$	Normality	$\chi^2(2)$	0.074	0.964
$\ln \theta$	Normality	$\chi^2(2)$	3.686	0.158
$\ln UP$	Normality	$\chi^2(2)$	2.865	0.239
mm	Normality	$\chi^2(2)$	4.687	0.096
VAR	Normality	$\chi^2(14)$	18.768	0.094
τ	ARCH(1)	F(1,34)	1.415	0.242
r	ARCH(1)	F(1,34)	0.140	0.711
$\ln u$	ARCH(1)	F(1,34)	0.230	0.634
$\ln \theta$	ARCH(1)	F(1,34)	0.005	0.944
$\ln UP$	ARCH(1)	F(1,34)	0.046	0.831
mm	ARCH(1)	F(1,34)	0.380	0.541
τ	ξ^2	F(14,21)	1.435	0.221
r	ξ^2	F(14,21)	0.470	0.925
$\ln u$	ξ^2	F(14,21)	0.613	0.825
$\ln \theta$	ξ^2	F(14,21)	0.829	0.634
$\ln UP$	ξ^2	F(14,21)	0.438	0.942
mm	ξ^2	F(14,21)	0.519	0.895
VAR	ξ^2	F(294,52)	0.431	1.000

Table 4: main VAR diagnostics

H_0: rank=ra	$\lambda_{\max}$	P-value	Trace	P-value
$ra = 0$	46.60	0.024	128.00	0.007
$ra \leq 1$	30.40	0.303	81.42	0.151
$ra \leq 2$	22.11	0.454	51.02	0.373
$ra \leq 3$	13.21	0.831	28.91	0.589
$ra \leq 4$	9.41	0.680	15.70	0.422
$ra \leq 5$	6.29	0.068	6.29	0.068
The P-values are from MacKinnon *et al.* (1996).				

Table 5: multivariate cointegration tests

ra	1	2
τ	0.088	0.192
r	0.561	0.705
$\ln u$	0.032	0.002
$\ln \theta$	0.548	0.690
$\ln UP$	0.032	0.024
mm	0.016	0.027

Table 6: tests for exogeneity: P-values

The likelihood ratio test cannot reject the restrictions both on the impact factors α and on the coefficients of the cointegrating vector β' (P-value=0.307). Importantly, we cannot reject the hypothesis that the coefficient attached to the regional mismatch is equal to 1. On the other hand, we reject the hypothesis that the mismatch coefficient is equal to 0 (P-value=0.033). Using the sample average of φ as an estimate of φ^*, we also strongly reject the restriction that the coefficient attached to mm equals $\varphi^*/(\varphi^* - 1)$ (the P-value in this case is only 0.0003).[39]

Using the definition of regional mismatch $mm = \ln u - \ln u_N$, we obtain

$$c_1 = \underset{(0.283)}{1.538}\,\tau + \underset{(0.595)}{2.216}\,r + \underset{(0.110)}{0.715}\,\ln UP - \ln u_{NC}. \tag{38}$$

that we interpret as follows: in the long-run equilibrium, a cointegrating relation exists between unemployment in the NC regions, the real interest rate, union power, and the tax wedge.

4.3 Implications

These findings have the following two important implications:

1. The Nairu, the unemployment rate that reconciles aggregate wage and price setting behavior, is not affected by Southern unemployment and depends only on unemployment in the NC regions. Hence, variations in the aggregate unemployment rate have no effect on wage and price dynamics unless they are driven by variations of u_{NC} with respect to the Nairu. Similarly, policies that reduce u_{SO} but maintain u_{NC} close to the Nairu have no effect on price and wage inflation.

[39] In the same way, using in the model $mm_{SO} = \ln(u) - \ln(u_{SO})$ instead of mm, we can reject the restriction that the coefficient of mm_{SO} is equal to 1 with a P-value 0.0000. However, the Var in this case performs slightly worse.

2. Higher taxes increase the Nairu. In our model, this result does not rely on the presence of real wage resistance to changes in the tax-wedge, as in Daveri and Tabellini (1997), but depends on the interaction between private wages, public wages and public employment, that are linked together in the long run by the government budget constraint.

When the interaction between the aggregate price and wage setting equations determines the equilibrium rate of unemployment in the NC regions, the natural question to ask is what determines in the long run the unemployment rate in the South. To answer this question, first notice that the unemployment rate in the South is defined by

$$u_{SO} = \ln F_{SO} - \ln L_{SO} - s_{G,SO} \tag{39}$$

where F_{SO}, L_{SO} and $s_{G,SO}$ are respectively the labor force, employment and the share of public sector employment in the South. Next, using equations (4) and (8), average employment in the South is defined by

$$L_{SO} = K_{SO}\left(\frac{L_{NC}}{K_{NC}}\right)\left[\sigma\left(u_{NC}\right)\right]^{1/(1-\alpha)} \tag{40}$$

where K_{SO}, K_{NC} and L_{NC} are respectively the capital stocks in the two areas and average employment in the NC regions. This equation combines two key aspects of the model: first, wages in the NC areas are higher than in the SO areas because of the local wage drift. Second, firms sell at the same price in the national market. Finally, using stars to indicate long term values, we obtain from (39) and (40) that the long run regional unemployment differential is given by

$$u^*_{SO} - \text{Nairu} = \left[\ln\frac{F^*_{SO}}{K^*_{SO}} - \ln\frac{F^*_{NC}}{K^*_{NC}}\right] - \left(s^*_{G,SO} - s^*_{G,NC}\right) - \frac{1}{1-\alpha}\ln\sigma\left(\text{Nairu}\right) \tag{41}$$

Given the capital stock in the two areas, equation (40) shows that an increase in the Nairu affects long run unemployment in the South in two ways. First, the capital-labor ratio in the NC areas increases. Second, the wage differential between the NC and the SO areas falls, because the local wage drift is smaller. As the wage in the NC area falls relative to the wage in the SO area, it becomes relatively less convenient to employ labor in the South. The combination of these two effects implies that the capital-labor ratio in the South increases more than in the rest of the country. Hence, unemployment differentials widen as the Nairu increases.

Long run unemployment in the South increases relative to the Nairu when a) the relative share of public employment falls in the long run; b) the relative ratio of the labor force over the capital stock increases in the long run. If we approximate long run values with 5-year moving averages of actual values, we find that the labor force has increased in the South relative to the NC areas by close to 10 percentage points during the years 1975-1994. During the same period, the capital stock in the South relative to the rest of the country has increased by close to 3 percentage points and the relative share of public employment in the labor force has remained more or less constant. Hence, long run relative changes in public employment and in the capital stock in the South have not been sufficient to absorb the relatively large increase of the labor force in the area. With limited inter-regional labor flows, this increase has generated an increase in the unemployment rate in the South.

4.4 Dynamic Adjustment

We use the restrictions on the impact factors and on the cointegrating vector to estimate the Ecm model specified in equation (36) by Full Information Maximum Likelihood. Since the real rate of interest, the tax wedge and the relative wage can be taken as weakly exogenous, we condition on these three variables and estimate a three-equation model in the changes of regional mismatch, the log of union power and the log of aggregate unemployment. After sequential simplification, we end up with the estimates in Table 7.[40] These estimates suggest that a temporary increase of unemployment in the NC regions over its long run value (a reduction in c_1) induces a correcting variation in aggregate unemployment u, an increase in regional mismatch and a decline in union power.

Having estimated the vectors α and β by $\widehat{\alpha}$ and $\widehat{\beta}$, we use a version of the Granger Representation Theorem (see *e.g.*, Johansen, 1995) to derive the moving average representation of our Var in order to compute the Ma impact matrix $\mathbf{C}$ consistent with the estimated restricted cointegrating relation as $\widehat{\mathbf{C}} = \hat{\beta}_\perp(\hat{\alpha}'_\perp\hat{\beta}_\perp)^{-1}\alpha'_\perp$, where $\hat{\beta}_\perp$ and $\hat{\alpha}_\perp$ are (6×5) matrices such that $\widehat{\alpha}'\widehat{\alpha}_\perp = \mathbf{0}$ and $\widehat{\beta}'\widehat{\beta}_\perp = \mathbf{0}$.[41] An interpretation of the matrix $\mathbf{C}$ is that of giving the long-run effect of shocks to the variables of the system. For example, the first column of $\mathbf{C}$ represents the long-run effect of a shock to the first equation. Furthermore, the long-run covariance matrix is a quadratic function of $\mathbf{C}$ and is proportional to the zero-frequency spectral density matrix of $\Delta\mathbf{X}_t$, $\mathbf{F}_{\Delta\mathbf{X}}(0)$, a natural candidate as a measure of persistence (see *e.g.* Engle and Yoo, 1991; Lupi, 1992, for some criticisms). In order to obtain more eas-

[40] In the dynamic regression, each equation includes also an impulse dummy for 1956.

[41] See the discussion in Paruolo (1997) on the asymptotic inference on $\widehat{\mathbf{C}}$.

Endogenous variable: $\Delta \ln UP$	Coefficient	Standard error
Constant	-0.347	0.09
$\Delta \ln \theta$	-0.308	0.12
c_1	0.129	0.03
σ^2	0.034	
Endogenous variable: $\Delta \ln u$		
Constant	-0.622	0.25
Δr	1.331	0.63
c_1	0.243	0.09
σ^2	0.09	
Endogenous variable: Δmm		
Constant	0.281	0.11
Δr	-0.310	0.25
c_1	-0.104	0.04
σ^2	0.04	
Serial Correlation $F\,(18, 93)$	[0.24]	
Normality $\chi^2\,(6)$	[0.72]	

Table 7: FIML estimates

ily interpretable results we compute the multivariate measure of persistence proposed by Pesaran *et al.* (1993), $\widehat{\mathbf{P}} = \{P_{ij}\}$, as

$$\widehat{P}_{ij} = \frac{\mathbf{e}_i'\widehat{\mathbf{C}}\widehat{\mathbf{\Omega}}_\varepsilon\widehat{\mathbf{C}}'\mathbf{e}_j}{\mathbf{e}_j'\widehat{\mathbf{\Omega}}_\varepsilon\mathbf{e}_j}\ , \qquad i, j = 1, 2, \ldots, 6 \tag{42}$$

where $\mathbf{e}_s$ is a (6×1) selection vector with its s-th element equal to one, and all the other elements equal to zero. $\widehat{P}_{i,j}$ represents the long-run effect of a unit shock in equation j on the level of the variable i. The equation-specific measures of persistence are the square roots of the elements on the main diagonal of $\widehat{\mathbf{P}}$, *i.e.* $P_i = \sqrt{P_{ii}}$. According to our estimates

$$\widehat{\mathbf{P}} = \begin{pmatrix} 1.000 & 0.052 & 0.007 & -0.001 & -0.015 & 0.061 \\ 0.240 & 1.000 & 0.082 & -0.123 & -0.022 & -0.154 \\ 0.571 & 1.527 & 0.474 & 0.057 & 1.443 & 0.405 \\ -0.019 & -0.437 & 0.011 & 1.000 & 0.600 & -0.506 \\ -0.232 & -0.073 & 0.254 & 0.557 & 3.240 & -0.566 \\ -1.333 & -0.717 & 0.101 & -0.067 & -0.802 & 1.245 \end{pmatrix}. \tag{43}$$

The ordering of the variables in (43) is τ, r, $\ln u$, θ, $\ln UP$, mm. Our computations show substantial persistence in the system. In particular, the equation-

Years	$\Delta \ln u$	Δmm
1	0	0
2	-3.9	1.7
3	-6.9	2.9
4	-9.0	3.9
5	-10.6	4.6
10	-14.5	5.1
15	-15.2	6.2
20	-15.2	6.5

Table 8: permanent change in the tax wedge: percentage deviations from the baseline and adjustment path

specific measures are

$$\Big(\underset{\tau}{1.000}, \underset{r}{1.000}, \underset{\ln u}{0.689}, \underset{\theta}{1.000}, \underset{\ln UP}{1.800}, \underset{mm}{1.116}\Big). \tag{44}$$

Furthermore, note that $\ln u$ shows substantial persistence also with respect to shocks in the other variables (see the third row in $\widehat{\mathbf{P}}$).

To further illustrate the dynamic adjustment of unemployment and regional mismatch, we present a simple simulation exercise. After obtaining a baseline by simulating the three equations dynamic model from 1955 to 2020, under the assumption that the weakly exogenous variables remain unchanged for the period 1997-2020, we study unemployment adjustment to a shock in the tax wedge. In particular, we assume that the wedge returns from 1997 onwards to its 1984 level, falling from 0.54% to 0.44 % of the labor cost, a 18.5% permanent reduction. Table 8 shows both the percentage deviation from the baseline 20 years after the shock and the adjustment path of both unemployment and regional mismatch during the period.

According to our simulations, the permanent reduction in the tax wedge from the 1996 value to the value prevailing in the early 80s yields a 15.3% reduction in the unemployment rate, with about 70% of this reduction taking place after 5 years and about 90% occurring after 10 years. This suggests that the adjustment process to the new equilibrium is rather sluggish.

The dynamic adjustment of the three endogenous variables $\ln u, mm$ and $\ln UP$ can be written as

$$\begin{bmatrix} \Delta \ln UP \\ \Delta \ln u \\ \Delta mm \end{bmatrix} = \begin{bmatrix} \gamma_0 & \gamma_1 & 0 & \gamma_3 \\ \gamma_4 & 0 & \gamma_6 & \gamma_7 \\ \gamma_8 & 0 & \gamma_9 & \gamma_{10} \end{bmatrix} \begin{bmatrix} 1 \\ \Delta \ln \theta \\ \Delta r \\ c_1 \end{bmatrix}. \tag{45}$$

Using the definitions of u and mm, and with a constant weight φ, the last two dynamic equations can also be written as

$$\Delta \ln u_{NC} = (\gamma_4 - \gamma_8) + (\gamma_6 - \gamma_9)\,\Delta r + (\gamma_7 - \gamma_{10})\,c_1 \tag{46}$$

$$\Delta \ln u_{SO} = \left(\gamma_4 + \frac{\varphi}{1-\varphi}\gamma_8\right) + \left(\gamma_6 + \frac{\varphi}{1-\varphi}\gamma_9\right)\Delta r + \left(\gamma_7 + \frac{\varphi}{1-\varphi}\gamma_{10}\right)c_1 \tag{47}$$

that describe the dynamics of unemployment in the NC and SO regions of the country.

Based upon the previous simulation exercise and estimating $\varphi = 0.58$ in the sample period, a permanent reduction in the tax wedge yields a 21.7 percent reduction in NC unemployment and a 6.2% reduction in SO unemployment. Hence, a permanent cut in the tax wedge is more effective in cutting unemployment in the NC areas than in the SO areas. While aggregate unemployment declines, regional disparities increase.

4.5 Discussion

Following Karanassou and Snower (1997) and defining equilibrium unemployment as the rate at which "..there is no tendency...to change at any time t, given the values of the exogenous variables at that time.." (p.561) we compute the NAIRU by taking the long-run solution of the dynamic model. We smooth the computed series by using 5-years moving averages of the driving variables. Moreover, we compute 95 percent confidence intervals using the delta method. The result is presented in Figure 7, that shows both aggregate actual unemployment and the estimated NAIRU. Notice that our estimates of the NAIRU are not very precise, with the width of the 95% confidence interval close to 2 percentage points.[42]

According to our estimates, the NAIRU fell from close to 7 percent in the mid-50s to close to 5 percent in the mid-60s, when the post-war economic boom reached its peak; it increased during most of the seventies to reach 7.2 percent in the late 70s and 8.4 percent in the mid-80s, and remained anchored to the 8.3 percent plateau for the following 10 years or so.

While the NAIRU was stable at 8.3 percent during the 1986-94 period, actual aggregate unemployment increased from 11.9 percent in 1986 to 14.9 percent in 1994. The gap between actual aggregate unemployment and the NAIRU increased during the period from 3.6 to 6.6 points but inflation fell only moderately, from 6.25% in 1986 to 4.26% in 1994. The reason is that

[42] This, however, is a common and well known problem. See the discussion in Staiger *et al.* (1997).

Figure 7: the NAIRU and actual unemployment

Years	ΔNAIRU	$\Delta\tau$	Δr	ΔUP
$60 - 74$	7.45	5.75	-13.08	14.78
$75 - 94$	27.63	24.05	22.08	–18.50

Table 9: decomposition of variations in the NAIRU. Percentage changes

actual unemployment in the NC regions increased during the period "only" from 8.9 to 9.7 percent (see Figure 8), while large part of the increase in aggregate unemployment was due to higher unemployment in the South.

To conclude, the results in Table 7 can be used to decompose the variations in the NAIRU in terms of the contributions of variations in the tax wedge, union power and the real rate of interest. This is done in Table 9 for two periods, 1960-74 and 1975-94. In the first sub-period, the NAIRU increased by less than 8 percent, mainly in association with the increase in union power. In the second sub-period, it increased much more sharply, as a result of the increase both of the real interest rate and of the tax wedge, only partially compensated by the decline in union power.

5 Conclusions

In this paper, we have estimated the Italian NAIRU using annual data for the period 1951-1996. We summarize our key results as follows:

1. aggregate wage setting in Italy depends only on the rate of unemployment prevailing in the Northern and Central areas of the country.

Figure 8: the NAIRU and unemployment Un

Southern unemployment does not affect wage pressure;

2. there is evidence of a long-run cointegrating relationship between unemployment in the Northern and Central areas, the tax wedge, the real interest rate and a measure of union power;

3. higher taxes on labor increase the natural rate of unemployment. According to our simulations, the permanent reduction in the tax wedge from the 1996 value to the value prevailing in the early 80s yields a 15.3% reduction in the unemployment rate, with about 70% of this reduction taking place after 5 years and about 90% occurring after 10 years.

4. persistence is an important feature of the Italian labor market;

5. the estimates of the NAIRU are not very precise and the width of the 95% confidence interval is close to 2 percentage points;

6. since the mid 80s, inflation in Italy fell from 6.25% to 4.26%. While the NAIRU was stable at 8.3 percent, actual aggregate unemployment increased from 11.9 percent in 1986 to 14.9 percent in 1994 and actual unemployment in the NC regions increased during the same period from 8.9 to 9.7 percent. Since the bulk of the increase in unemployment occurred in the South, there was relatively little downward pressure on wages and prices;

7.

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A Data Appendix

The sources of the data used in the empirical analysis carried out in this paper are briefly listed below:

u = aggregate unemployment rate; 1959-1994: Bank of Italy; 1951-1958: Brunello-Checchi (1997); 1995-96: estimated by the authors using Bank of Italy, *Relazione Annuale*, Appendix, Table AB23.

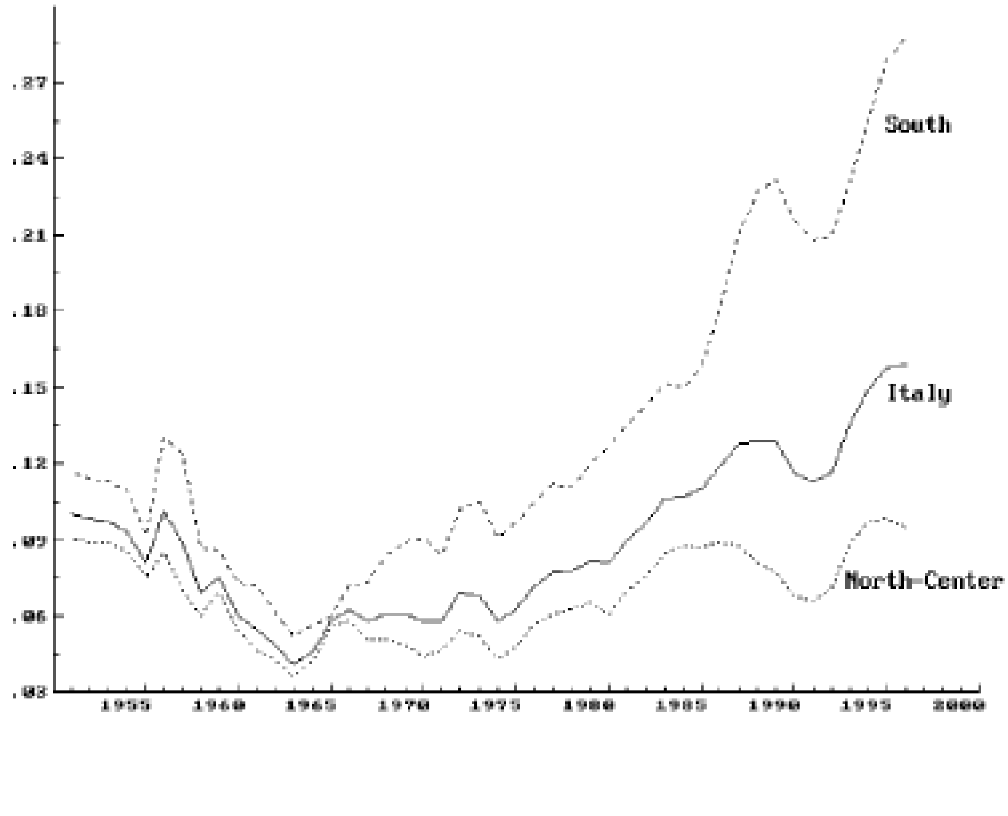
u_{NC}, u_{SO} : unemployment rates in the Northern and Central areas and in the South, respectively; 1954-1994: Bank of Italy; 1995-96: estimated by the authors using Bank of Italy, *Relazione Annuale*, Appendix, Table AB26. The data 1951-53 are estimated by the authors by applying backwards the dynamics of the aggregate series.

UP = union membership rates; see Brunello-Checchi (1997).

r = nominal short-term interest rate minus the inflation rate. Source: The European Economy.

τ = payroll and income taxes measured as percentages of gross wages. Source: Baviera-Rossi (1993), updated by the authors using Bank of Italy, *Relazione Annuale*, Appendix.

$\ln\theta$ = difference of the logs of the wages in the private and public sectors. Source: Istat, National Accounts.



South
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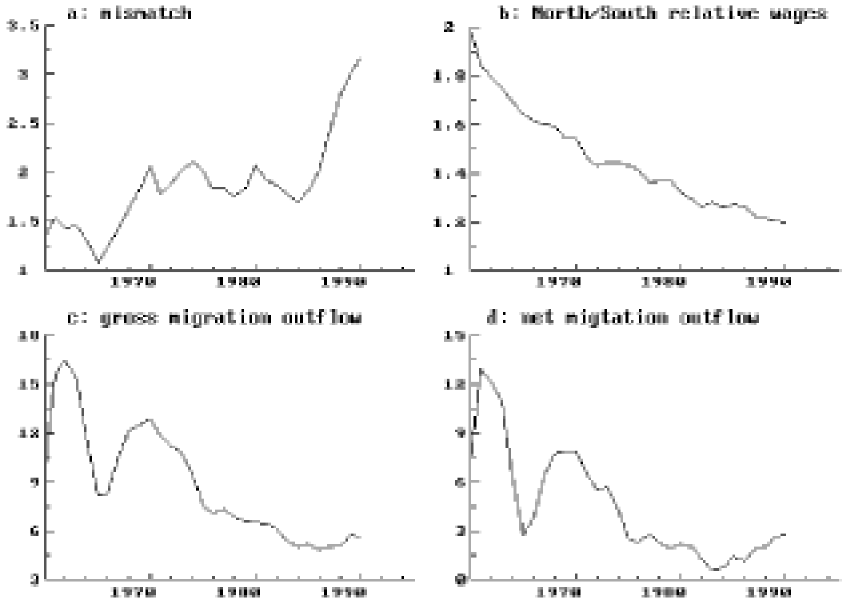

a: mismatch
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1980
1990
b: North/South relative wages
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1.6
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1
1970
1980
1990
c: gross migration outflow
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d: net migtation outflow
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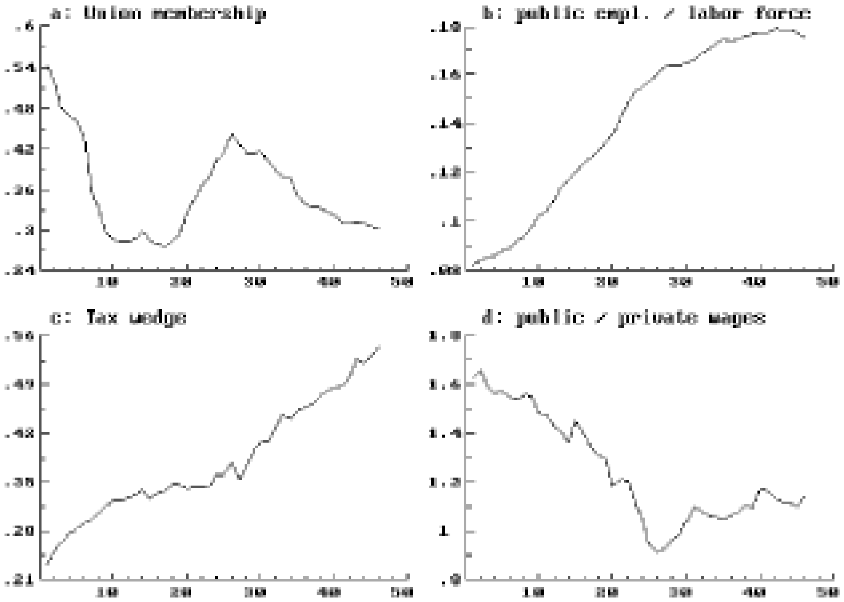
a: Union membership
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10
20
30
40
50
b: public empl. / labor force
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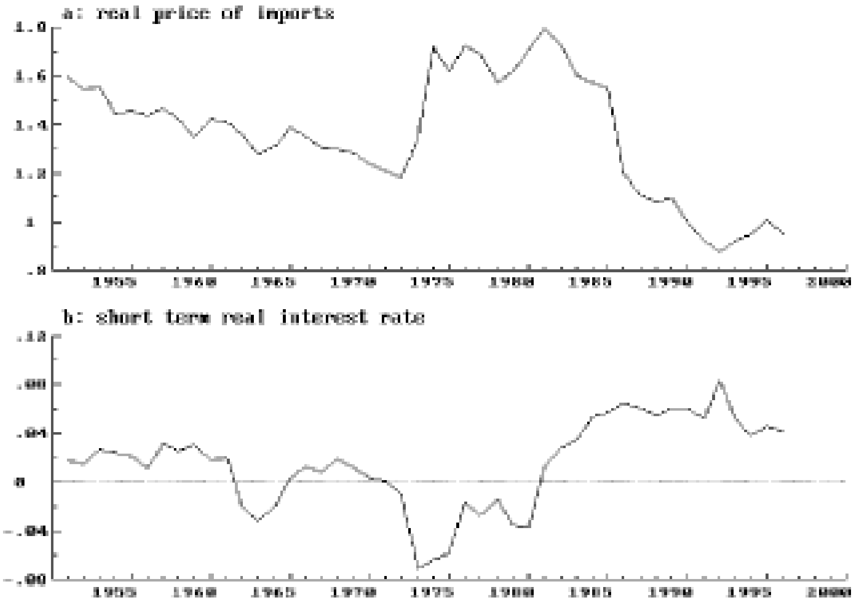
.16
.14
.12
.1
.08
10
20
30
40
50
c: Tax wedge
.56
.49
.42
.35
.28
.21
10
20
30
40
50
d: public / private wages
1.8
1.6
1.4
1.2
1
.8
10
20
30
40
50

Inflation rate
.22
.2
.18
.16
.14
.12
.1
.08
.06
.04
.02
0
-.02
1955
1960
1965
1970
1975
1980
1985
1990
1995
2000



a: real price of imports
1.8
1.6
1.4
1.2
1
.8
1955
1960
1965
1970
1975
1980
1985
1990
1995
2000
b: short term real interest rate
.12
.08
.04
0
-.04
-.08
1955
1960
1965
1970
1975
1980
1985
1990
1995
2000

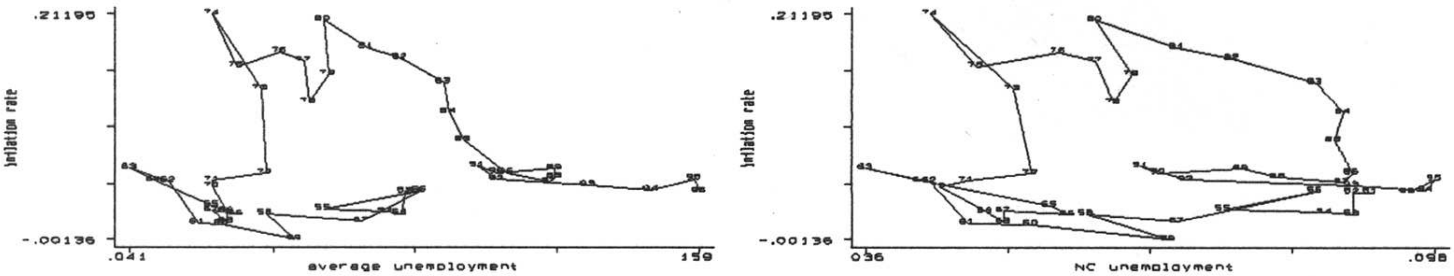


fig. 6: Inflation unemployment trade offs

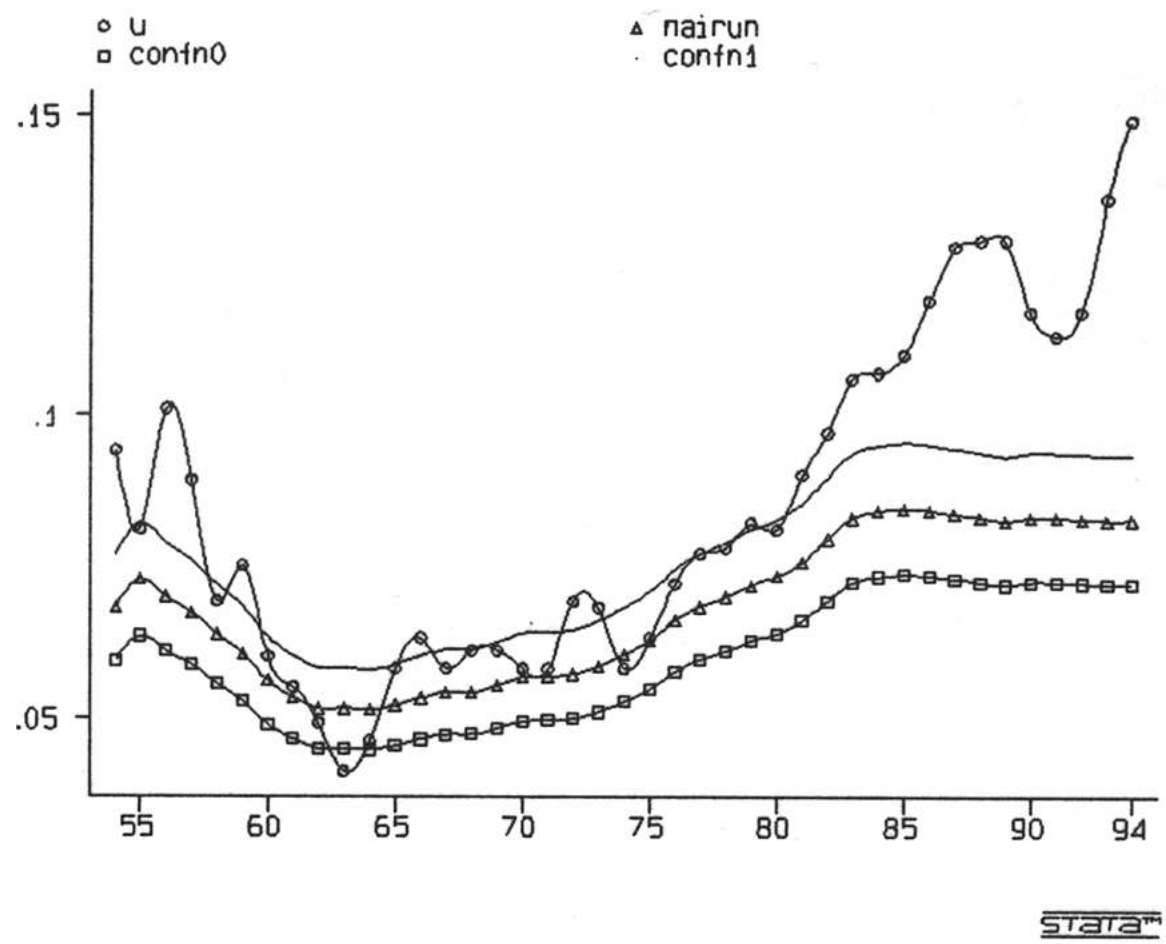


fig.7: the NAIRU and actual unemployment

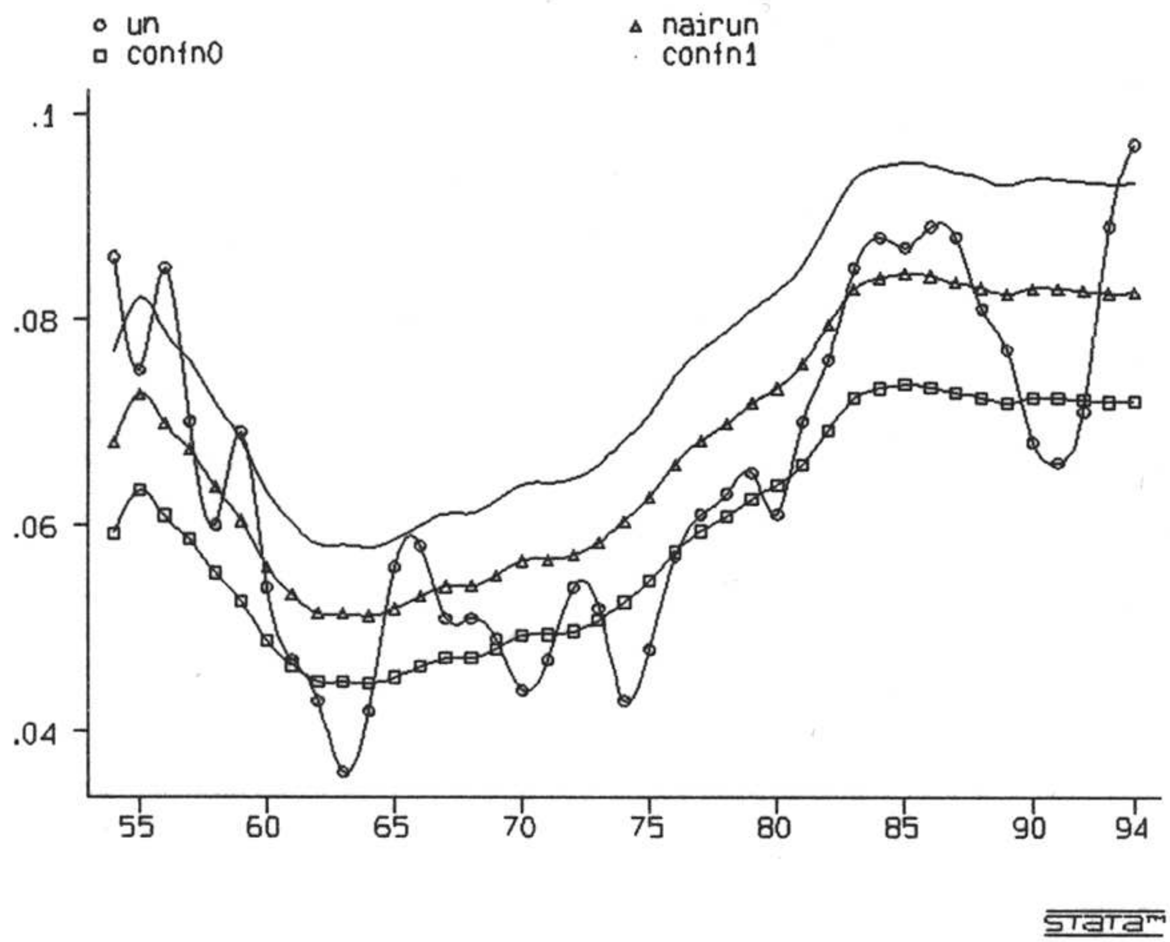


fig. 8: the NAIRU and unemployment Un