Working Paper No. 475

Capital Stock and Unemployment: Searching for the Missing Link

by

Alfonso Palacio-Vera,
Ana Rosa Martínez-Cañete,
Elena Márquez de la Cruz,
and
Inés Pérez-Soba Aguilar,
Universidad Complutense de Madrid (Spain)

September 2006

Address for correspondence:
Alfonso Palacio-Vera (E-mail: apv@ccee.ucm.es)
Departamento de Economía Aplicada III
Facultad de Ciencias Económicas y Empresariales
Universidad Complutense de Madrid
Campus de Somosaguas
Madrid 28223
Spain

The Levy Economics Institute Working Paper Collection presents research in progress by Levy Institute scholars and conference participants. The purpose of the series is to disseminate ideas to and elicit comments from academics and professionals.

ABSTRACT

The purpose of this paper is to examine the proposition that capital stock relative to aggregate output has been an important variable in the determination of the U.S. NAIRU (Non-Accelerating Inflation Rate of Unemployment) over the last four decades. We present new empirical evidence, obtained from the application of the cointegrated VAR methodology to U.S. time-series data, that lends strong support to the claim that the aggregate capital-output ratio, the real price of imports, and aggregate capacity utilization were determinants of the NAIRU in the period considered. The same evidence also shows that technical progress and changes in long-term unemployment did not affect the NAIRU. We believe this evidence suggests that, insofar as the aggregate capital-output ratio is affected by changes in real interest rates, the stance of monetary policy is one determinant of the NAIRU.

I   INTRODUCTION

The purpose of this paper is to examine the proposition that capital formation is an important variable in the determination of unemployment in the U.S. economy. Most of the literature on unemployment has focused on labor market issues, such as wage-fixing institutions, the role of welfare benefits and legal firing costs, and the quality and motivation of the labor force. However, there are other aspects of the unemployment problem which have been rather neglected. One such area is the relationship between unemployment and capital stock. Some economists have criticized this neglect and stressed the importance of capital stock for explaining unemployment. Some examples are the studies by Malinvaud (1986), Sneessens and Dréze (1986), Modigliani et al. (1987), Bean (1989), Rowthorn (1999), Allen and Nixon (1997), Sawyer (2002), and Stockhammer (2004). The link between capital stock and unemployment has also recently been the focus of attention of several studies that make use of modern time-series analysis (Arestis and Biefang-Frisancho Mariscal 1998 and 2000). But these authors are in the minority and the conventional wisdom seems to be that persistent unemployment is mainly due to labor market rigidities.[1] This position is well exemplified in the influential work of Layard, Nickell, and Jackman (1991). In their impressive econometric study of OECD unemployment, they impose cross-equation restrictions which ensure that the rate of unemployment is unaffected by technical progress, as well as by changes in the aggregate capital-labor ratio. This position has become the conventional wisdom in macroeconomic analysis. For instance, Blanchard and Katz (1997) argue "any model should satisfy the condition that there is no long-run effect of the level of productivity on the natural rate of unemployment."[2]

But the absence of such a trend in unemployment over very long periods does not imply that unemployment may not exhibit a trend over shorter periods, like several decades, or that unemployment may not be affected by, for example, changes in the

---

[1] Notwithstanding this dominant view, there is the important issue of the blatant inconsistency between the widely recognized reduction in the number and importance of labor-market rigidities in European economies in the 80s and 90s and the observed rising trend in the rate of unemployment in many of these countries in the same period (Nickell 1997). Although some recent studies have shed some light on this issue (see for instance Blanchard and Wolfers 2000), the debate is still far from being settled.

[2] Blanchard and Katz (1997) discuss thoroughly a number of mechanisms through which this result will be accomplished.

aggregate capital-output ratio. In fact, unemployment has been on the rise in several European countries for almost two decades and this is attributed to a number of factors, including the generosity of European welfare systems and other related labor market rigidities, in a context of an increasingly competitive world economy (Siebert 1997). We do not wish to deny that globalization may, at least partially, explain the rise in European unemployment. However, we believe that another causal factor, albeit much ignored in the literature, is the role played by capital accumulation.

One possible source of changes in labor productivity which has barely been discussed in the literature is changes in the aggregate capital-output ratio. A fall (rise) in the latter will transitorily reduce (increase) the rate of growth of labor productivity, thus, initially causing a mismatch between wage aspirations and labor productivity growth. Changes in the aggregate capital-output ratio may come about as a result of: (i) changes in technology, (ii) changes in the price of labor, *vis á vis* the rental price of capital, and (iii) changes in the pace of investment in physical capital for reasons other than (i) and (ii). In Layard, Nickell, and Jackman (1991), there are two reasons why changes in the capital-labor ratio do not influence the equilibrium level of unemployment. The authors point out that *if* the production function is Cobb-Douglas *and* benefit-replacement ratios are kept stable in real terms, then unemployment in the long run is independent of capital accumulation and technical progress.[3] For this to be the case, one has to assume that real wage aspirations adapt rapidly to changes in productivity growth. However, if workers demand real-wage increases based on their previous experience and only gradually adjust their demand to the new trend rate of productivity growth, then there is the possibility that shifts in productivity growth affect the NAIRU (Non-Accelerating Inflation Rate of Unemployment). Therefore, as a minimum, changes in the aggregate capital-output ratio are a potential source of changes in the NAIRU. To the best of our

---

[3] They nonetheless mention in passing that "if, however, the elasticity of substitution is less than one, capital accumulation (with no technical progress) raises the share of labor and reduces unemployment" (Layard, Nickel, and Jackman 1991). Interestingly enough, Rowthorn (1999) shows that the bulk of the empirical estimates of the elasticity of substitution between capital and labor are well below unity and concludes that the assumption of a Cobb-Douglas production function (for which the elasticity of substitution is equal to unity) for most economies is inappropriate. Furthermore, and importantly, using a simple model based on the work of Layard, Nickell, and Jackman (1991), Rowthorn (1999) shows that, with an elasticity of substitution lower than unity, the equilibrium rate of unemployment is affected by technical progress, labor force expansion and capital accumulation. Unfortunately, a number of technical difficulties prevented us from utilizing a more general formulation, like a CES production function, rather than a Cobb-Douglas function. Use of a CES production function would have compelled us to implement non-linear estimation methods, thus making the empirical investigation much more complex. We nevertheless consider this an avenue for future research.

knowledge, this possibility has not been tested empirically. Our objective is to test whether changes in the former has had a significant impact upon the latter in the U.S. economy in the last four decades.

Hence, we set ourselves the task of exploring the (unexplored) possibility that changes in labor productivity, brought about by changes in the amount of physical capital relative to output, are *not* fully translated into changes in real wages for a long enough time span so they may lead to changes in the equilibrium rate of unemployment. For that purpose, we present a simple model characterizing an economy where imperfect competition prevails in the product market and wage increases are bargained over by firms and workers. We obtain an expression for the equilibrium rate of unemployment, or NAIRU, that is then estimated using co-integration techniques in the context of the VAR methodology. The model is estimated using quarterly data obtained mostly from the EcoWin Pro database for the period 1964q2 to 2003q1 for the United States economy. [4] Our empirical results challenge conventional wisdom in the field, for they lend strong support for the hypothesis that the aggregate capital-output ratio was one significant determinant of the NAIRU in the United States in the last four decades. Results also suggest that the real price of imports and capacity utilization were significant determinants of the NAIRU, whereas factors like technical change and long-term unemployment did not affect the NAIRU significantly.

Since we obtained that the U.S. NAIRU depends negatively on the aggregate capital-output ratio and, hence, positively on the level of real interest rates, these results identify a link between capital stock and unemployment that has been unexplored empirically to this date. We believe our findings shed new light and qualify previous results in the literature on the presence of hysteresis in labor markets (Cross 1995; Blanchard and Summers 1987; Franz 1987 and 1990; Sachs 1987) and are of relevance to, among other topics, the effects of disinflation episodes on equilibrium unemployment (Fortin 1996; Romer and Romer 1989; Ball 1997 and 1999), the choice of identification restrictions in VAR models (Blanchard and Quah 1989; Galí 1992) and the debate on the causes of European unemployment (Ball 1997 and 1999; Blanchard and Wolfers 2000; Bibow 2001 and 2005; Hein and Truger 2005).

---

[4] EcoWin Database consists of economic and financial data collected from more than seven hundred primary sources as OECD, IMF, and Bureau of Labor Statistics, among others.

The content of the paper is as follows. Section II reviews the empirical literature on this topic. Section III displays a simple theoretical model that yields an expression of the NAIRU that is subsequently used for estimation purposes. The empirical investigation appears in section IV. Section V summarizes our findings and concludes.

II   REVIEW OF THE EMPIRICAL LITERATURE

As argued above, a number of authors have discussed the role of capital in explaining unemployment. However, recent discussions on this topic have tended to adopt the influential work of Layard, Nickell, and Jackman (1991) as the starting point for formal discussions. These authors impose restrictions on the price and wage equation such that the rate of unemployment is unaffected by technical progress and changes in the aggregate capital-labor ratio. In particular, they impose an equality of coefficients on the trend productivity term in the price and wage equation to the effect that any shift in the real product wage-employment relationship derived from pricing considerations generates a corresponding shift in the wage-setting equation, such that the equilibrium level of unemployment does not change so the benefits of higher productivity always feed through into higher real wages (Sawyer 2002). As Layard, Nickell, and Jackman (1991) recognize, were these coefficients to differ, then unemployment would either rise or fall continuously with trend productivity growth. They argue that the absence of such a trend in unemployment over centuries is consistent with their framework (Layard, Nickell, and Jackman 1991).

However, as has been the case of some European economies, the NAIRU may well exhibit a trend over several decades. One possible source of persistent changes in the NAIRU is changes in the aggregate capital-output ratio. In particular, a higher capital stock relative to output (or labor) raises labor productivity, thereby mitigating inflationary pressures and allowing the economy to operate with a lower rate of unemployment. But real wage aspirations will tend to adjust upwards. As noted in Blanchard and Katz (1997), the equilibrium rate of unemployment or NAIRU, i.e., the rate of unemployment compatible with a constant rate of inflation in the absence of transitory supply shocks, depends on the level of productivity in relation to the reservation wage, as well as on many other factors. Accepting this then, if real wage aspirations do not *fully* offset the effects of capital accumulation and/or technical

progress for a period spanning, say one decade or more, then the rate of unemployment will exhibit a (downward) trend for some time.[5]

A similar argument has been put by Ball and Mankiw (2002). They argue that in a neoclassical world, a rise in productivity growth has no obvious effect on inflation because higher productivity is reflected fully in higher real wages. Nonetheless, and building on the work of Grubb, Jackman, and Layard (1982), Stiglitz (1997), and Blinder (2000), they suggest there is a link between *shifts* in productivity growth and the NAIRU that may help explain both the rising NAIRU of the 1970s and the falling NAIRU of the 1990s in the United States economy.[6] According to them, all we need to accept is that "wage aspirations" adjust slowly to shifts in productivity growth.[7] In particular, Ball and Mankiw argue that:

> "If productivity growth falls, as in the 1970s, fundamentals dictate that real-wage growth must fall as well. Workers resist this decrease, however; they try to maintain the wage increases to which they are accustomed. To the extent that workers have some influence over wages, this means that wage setters will try to achieve real-wage increases above the level that can be sustained by productivity growth. This mismatch between real-wage aspirations and productivity growth worsens the inflation-unemployment tradeoff. In other words, the NAIRU rises… Today´s version of the story reverses the signs. Productivity has accelerated but workers have become accustomed to the slow wage growth since the 1970s. A mismatch of productivity and wage aspirations in this direction shifted the Phillips curve favourably" (Ball and Mankiw 2002).

There is at least one more channel through which capital stock may affect the equilibrium rate of unemployment. In the work by Rowthorn (1977), and in several subsequent studies, wage and price setting behavior depend on the state of demand in the particular market. Unemployment is a disciplinary device that reduces the ability of workers to push up wages, while excess productive capacity limits the ability of firms to raise prices. Therefore, additional productive capacity reduces the inflationary conflict

---

[5] Blanchard and Katz (1997) argue that the conditions that deliver long-run neutrality of the natural rate to changes in productivity need not hold as a matter of logic, but they all appear plausible. Reversing their argument, we may argue that, even if long-run neutrality may appear plausible, it does not hold as a matter of logic and so there is no need to impose it *a priori*.

[6] In the pioneering work by Grubb, Jackman, and Layard (1982), the stagflation suffered by OECD economies since 1975 was attributed partly to rising relative import prices and partly to the fall in the rate of productivity growth relative to workers' target rate of growth of real wages.

[7] They also mention that Alan Greenspan (1997) "has suggested that workers, cowed by job insecurity, lacked aggressiveness in wage negotiations" (Ball and Mankiw 2002). However, they point out that what matters is aggressiveness of workers relative to productivity so that failure to impose on firms a higher rate of growth of real (product) wages when productivity accelerates will have the same effect on the NAIRU as an exogenous decrease in workers' degree of aggressiveness.

over income distribution by preventing firms from raising profit margins and, thus, reducing the wage share. The impact of capital accumulation on equilibrium unemployment through this channel has been explored for the case of Germany and the U.K. in Arestis and Biefang-Frisancho Mariscal (1998 and 2000). Using cointegration techniques, their empirical results suggest that unemployment rose in the 80s and 90s due to insufficient investment. However, we believe their results are subject to the objection that the theoretical model they estimate predicts that firms' profit margins are pro-cyclical. Yet, the empirical evidence on the relationship between profit margins and capacity utilization (or, more generally, between market power and business fluctuations is mixed).

The link between unemployment and capital accumulation has also been recently explored empirically in Stockhammer (2004). He compares the NAIRU hypothesis regarding European unemployment and a Keynesian approach. The theories are tested using time series data for Germany, France, Italy, the U.K., and the United States using the seemingly unrelated regression method. He finds that a simple NAIRU specification performs poorly and that the Keynesian approach is more successful with the rate of capital accumulation being a significant determinant of employment growth in all countries. However, the simple NAIRU specification he tests empirically does not address the potential effect of technical change and changes in the aggregate capital-output ratio on the NAIRU.

The theoretical model we display in Section III below takes inspiration from the work of Allen and Nixon (1997). The authors present a model of the NAIRU similar to the model in Layard, Nickell, and Jackman (1991) and drop the cross-equation restriction on the capital-labor ratio in the price and wage equation. They justify this departure from Layard, Nickell, and Jackman (1991) by arguing that setting such restriction "obscures the fact that there are actually two processes taking place: capital accumulation and technological progress" (Allen and Nixon 1997). As a result of dropping this assumption, they show that the NAIRU is a function of relative factor prices and of the rate of accumulation so that there will be no unique short-run NAIRU. Besides, they argue that the level of interest rates and the rate of growth of aggregate demand determine a whole locus of equilibrium unemployment rates. However, the authors do not estimate their model and, to the best of our knowledge, no attempt has been made to this date to test empirically for the effect on the NAIRU of changes in the aggregate capital-labor ratio and technical change. Finally, the theoretical connection

6

between the NAIRU and the aggregate capital-labor ratio is also addressed in Rowthorn (1999), although his results only apply to the case of a CES production function, where the elasticity of substitution between capital and labor is less than unity.

III   A THEORETICAL MODEL OF UNEMPLOYMENT DETERMINATION AND THE ROLE OF CAPITAL STOCK

In this section, we put forward a model that aims to explain the determination of the rate of unemployment in the presence of trade unions and imperfectly-competitive product markets. Central to the model is the effect of capital stock on unemployment. Unlike some studies that postulate that the long-run impact of capital accumulation on the rate of unemployment occurs through the impact of the former on firms' mark-ups, we postulate a different framework in which capital stock affects unemployment mainly through its effect on the marginal product of labor. The short-run profit maximizing decision facing the typical firm, i, consists of maximizing profits

$$\Pi_t^i = p_t^i Y_t^i - w_t^e N_t^i \tag{1}$$

where p_t^i is the price charged for its output Y_t^i by firm i, w_t^e is the expectation of future labor costs per employee, N_t^i is employment and the subscript t denotes time. The decision variables for the firm are taken to be output (and thus, employment) and the price charged for it. For the sake of convenience, we assume that the economy consists of a number, n, of identical and fully integrated firms using equal amounts of labor and capital and similar technology. This assumption allows us to simplify substantially the firms' price-setting equation since we only need to consider one variable production factor, i.e., labor. Hence, the first-order condition of the optimization problem faced by the typical firm yields the following economy-wide relation:

$$p_t = m_t \cdot \frac{w_t^e}{\Omega_t^N} \tag{2}$$

where Ω_t^N is the marginal product of labor and m_t is equal to one plus the average mark-up set by the typical firm. Under imperfect competition we have that:

$$m_t = \left(\frac{1}{1 - 1/\varepsilon_t^d} \right) \succ 1 \tag{3}$$

Hence, the size of the (average) mark-up depends inversely on the price-elasticity of demand ε_t^d and, as we argue below, the latter may well vary cyclically.

Next, we assume that the technology available to firms is represented by a Cobb-Douglas production function so that the economy-wide output is equal to:

$$Y_t = \gamma_0 \cdot e^{\lambda_1 t} \cdot N_t^a \cdot K_t^{1-a} \tag{4}$$

where K_t is the amount of physical capital available to firms, a is a distribution parameter, γ_0 is a constant, and λ_1 is the rate of technical change.[8] Differentiating (4) with respect to employment, we can obtain the marginal product of labor or:

$$\Omega_t^N = \gamma_0 e^{\lambda_1 t} a \left(K_t / N_t \right)^{1-a} \succ 0 \tag{5}$$

Likewise, the marginal product of capital is:

$$\Omega_t^K = \gamma_0 e^{\lambda_1 t} (1-a) \cdot \left(K_t / N_t \right)^{-a} \succ 0 \tag{6}$$

Insofar as the typical firm combines capital and labor so as to minimize costs, we have that:

$$\frac{dK}{dN} = \frac{-\Omega_t^N}{\Omega_t^K} = \frac{-w/p}{r} \tag{7}$$

where w/p is the real price of labor and r is the rental price of capital services.

Combining expressions (4), (5), (6), and (7) we can express the marginal product of labor as:

$$\Omega_t^N = \gamma_0^{1/a} \cdot e^{(\lambda_1/a)t} \cdot a \cdot (K/Y)^{\frac{(1-a)}{a}} \tag{8}$$

Our approach to wage determination starts from the notion that both workers and firms typically have some bargaining power. The bargaining power of workers stems from the fact that they cannot be costlessly replaced. Likewise, the bargaining power of firms arises because workers cannot costlessly locate an equivalent job. We assume that workers bargain with a view to obtaining a target real (consumption) wage. This implies that they will resist any attempt by firms to reduce real (product) wages, as well as any loss of purchasing power stemming from any factor that may drive a wedge between the former and workers' purchasing power, e.g., taxes levied on consumption goods and/or

---

[8] Technological progress comes with structural change. Increases in the pace of technological progress are likely to come with a higher pace of reallocation across jobs and, to the extent that this leads to larger flows of workers in the labor market, these will lead to a higher unemployment rate. However, and for simplicity, we consider a form of technological progress that affects total factor productivity but does *not* affect the organization of production in any other way.

labor income and variations in the real price of imported consumption goods. As a result of it, the actual wage outcome is given by:

$$w_t = \omega(\varphi_t, I_t) \cdot e^{\lambda_2 t} \cdot p_t^e \quad (9)$$

where w_t is the actual money wage, p_t^e is the expected price level, φ_t is the cost to workers of losing their job, I is the real price of imports, $\lambda_2 \succ 0$ is the (trend) rate of growth of workers' real reservation wage, i.e., the wage that makes a worker indifferent to being employed or unemployed, and $\omega_\varphi \prec 0$ and $\omega_I \succ 0$ are partial derivatives. In turn, the real price of imports can be expressed as:

$$I = \frac{E \cdot p_m}{p} \quad (10)$$

where E is the price of foreign currency, p is the domestic price level, and p_m is the price of imports in terms of foreign currency.

Next, we assume that the cost to workers of losing their job, φ, depends inversely on: (i) the probability of finding another job, which in turn depends inversely on the current rate of unemployment (U) and (ii) the ratio of long-term unemployed to total unemployed (LU).[9] Therefore, we define φ_t as:

$$\varphi_t = \varphi(U_t, LU_t) \quad (11)$$

where we assume that φ_t is log-linear and $\varphi_U \succ 0$, $\varphi_{LU} \prec 0$.

Expressions (9) and (11) require some explanation. They make workers' target real wage depend inversely on the cost of job loss, as well as on a real wage resistance factor, such as the real price of imports.[10] An increase in the unemployment rate raises

---

[9] The cost to workers of losing their job may also be affected by changes in the replacement ratio. However, as noted in Blanchard and Katz (1997), in the United States, the ratio of average weekly unemployment benefits to average weekly wages for workers covered by unemployment insurance has been remarkably stable over the last 50 years, despite the peak in 1976 due to a transitory effect of a reform and enlargement of the American Unemployment Insurance. This stability was confirmed by the unit root tests we ran on a measure of the replacement ratio, which showed that the latter is I(0). If we add to this the fact that the replacement ratio is an exogenous variable from a theoretical perspective, we have that it will be part of the short-term adjustment dynamics but will *not* be part of the long-term relationships. Hence, and given the difficulty of estimating a VAR model containing many variables, we opted to leave the replacement ratio out of the theoretical model.

[10] Another possible resistance factor that could possibly affect the target real wage is taxation. Yet, the empirical literature on the NAIRU using time-series data has not attributed a significant effect of taxation on the equilibrium rate of unemployment. In addition, the analysis of several measures of the percentage of workers' income that they regularly pay as taxes and social security contributions, excluding Government transfers to households and a number of public services directly enjoyed by the latter showed they all are I(0). Insofar as this variable can also be deemed exogenous in our theoretical model, we opted, as in the case of the replacement ratio, to exclude it from the model from the start.

the cost of job loss and makes it easier for firms to replace currently employed workers. This stems from the fact that in a depressed labor market, workers know that finding another job is likely to be difficult, so they will be willing to settle for a lower wage. In this case, the target real wage will be low and close to the workers' reservation wage.[11] Conversely, in tight labor markets, the target real wage will be much higher than the reservation wage. Hence, even if the reservation wage were constant overtime, the target real wage would vary with labor market conditions.

The positive sign for the effect of the long-term unemployed (*LU*) on the target real wage may be explained by reference to membership theories. As argued above, high rates of unemployment tend to moderate wage claims; however, it is only the effective labor supply, i.e., the "insiders" that determine wage negotiations (Lindbeck and Snower 1988). If the composition of the unemployed changes in such a way that the ratio of the long-term unemployed rises, "insiders" may press for higher wages since there are fewer effective competitors for a given number of jobs. In turn, this stems from the fact that the long-term unemployed may not be perceived by employers as good substitutes for current incumbent workers, owing to the depreciation of their skills and the loss of work motivation in the unemployed or simply because the long-term unemployed may become less effective in their job search[12] (Blanchard et al. 1986; Franz 1987).

Next, models based on notions of fairness suggest that the reservation wage may depend on such factors as the level and rate of growth of wages in the past and if workers have come to consider such wage increases as fair (Akerlof and Yellen 1990). Perhaps a better word than reservation wage in this context is "aspiration" wage (Oswald 1986). The latter is determined by all the factors that the literature has identified as determining the reservation wage, plus the aspiration for a steady improvement in living standards. As such, the rate of increase of the "aspiration" wage

---

[11] The inverse relationship between the target real wage and the rate of unemployment can also be derived from efficiency wage considerations. For instance, the "shirking" model of Shapiro and Stiglitz (1984) leads to a wage relation in which the tighter the labor market is the higher is the wage that firms have to pay to prevent shirking.

[12] It has also been argued that sociological factors may increase the reservation wage of the unemployed: a long period of unemployment changes society's attitudes toward the unemployed. It becomes more socially acceptable to be unemployed and to use existing social benefits to their utmost (Lindbeck 1995). Furthermore, family insurance may also develop to help the unemployed, in particular the unemployed youth.

will be determined, at least partly, by a wide range of idiosyncratic socioeconomic and historical factors. The (positive) time trend on equation (9) can be envisaged as capturing workers' aspiration to ever-rising living standards. An additional rationale for the time trend is the notion that the reservation wage (regardless of wage aspiration considerations) depends on the various types of income support that the unemployed can expect to receive if unemployed from unemployment benefits to other social insurance programs. If all these increase over time and are roughly in line with labor productivity, then the reservation wage will exhibit a time trend.

Insofar as workers bargain over a target (consumption) real wage, we also need to take into account the wedge between product wages and consumption wages. This wedge includes not only taxes on labor net of income transfers from the government to workers, but also includes the discrepancy between the GDP deflator and the consumer price index. For reasons explained above, the former were not considered in the theoretical model. As for the latter, it depends mainly on the real price of imports (I).[13] Therefore, to the extent that workers exhibit some degree of real wage resistance, increases in the real price of imports will induce upward pressure on the real product wage and vice-versa. Next, if we insert (11) into (9), assume that function ω is log-linear, and take logarithms in (9) we get:

$$\ln w_t - \ln p_t^e = \overline{\omega} + \omega_u u_t + \omega_{lu} lu_t + \omega_i i_t + \lambda_2 t \qquad (12)$$

where the logarithms of the arguments in ω are denoted by lower-case letters, $\overline{\omega}$ can be interpreted as the "aspiration" (real) wage at the initial period, $\omega_U \prec 0$ and $\omega_{LU} \succ 0$.

Some studies find that firms' mark-ups exhibit cyclical patterns. However, there is no consensus as to whether they are pro-cyclical or counter-cyclical (see, for instance, Hall 1986; Bils 1987; Rotemberg and Woodford 1991; Chirinko and Fazzari 1994; and Galeotti and Schiantarelli 1998). A discussion of the vast literature on the relation between firms' market power and macroeconomic fluctuations exceeds the aims of this study. Hence, we adopt an agnostic approach and propose an indirect way of capturing the possible cyclicality of mark-ups. In particular, we hypothesize that mark-ups are set according to:

$$m_t = \overline{m} \cdot C_t^\phi \qquad (13)$$

---

[13] It also depends on excise taxes, but we decided to leave aside this variable due to data availability problems.

where C_t is aggregate capacity utilization, $\overline{m}$ is the average mark-up set by firms when $, C_t = 1$ and the sign of ϕ is *a priori* uncertain. Expression (13) means that the price-elasticity of demand varies according to the phase of the business cycle the economy is currently at, where the latter is captured by aggregate capacity utilisation. Insofar as firms' mark-ups are determined by capacity utilization, the latter will affect the NAIRU, e.g., if $\phi \succ 0$, then increases in capacity utilization will raise the NAIRU and vice-versa. If we take logarithms in (13) we get:

$$\ln m_t = \ln \overline{m} + \phi \cdot c_t \tag{14}$$

where $c_t = \ln C_t$ and taking logarithms in (2) and (8) yields, respectively:

$$\ln p_t - \ln w_t^e = \ln m_t - \ln \Omega_t^N \tag{15}$$

and
$$\ln \Omega_t^N = \Omega_0 + \Omega_1 \cdot t + \Omega_2 k_t \tag{16}$$

where $k_t = \ln(K_t / Y_t)$, $\Omega_0 = \ln a + (1/a) \ln \gamma_0$, $\Omega_1 = (1/a) \lambda_1$ and $\Omega_2 = \left(\dfrac{1-a}{a} \right)$.

Inserting (14) and (16) into (15) yields:

$$\ln p_t - \ln w_t^e = \ln \overline{m} + \phi c_t - \Omega_0 - \Omega_1 t - \Omega_2 k_t \tag{17}$$

We define long-run equilibrium as a situation where exogenous factors are kept fixed and expectations are fulfilled. If expectations are fulfilled, we have that $w_t = w_t^e$ and $p_t = p_t^e$, hence, (12) and (17) become respectively:

$$\ln w_t - \ln p_t = \overline{\omega} + \omega_u u_t + \omega_{lu} lu_t + \omega_i i_t + \lambda_2 t \tag{18}$$

$$\ln p_t - \ln w_t = \ln \overline{m} + \phi c_t - \Omega_0 - \Omega_1 t - \Omega_2 k_t \tag{19}$$

Hence, when there is long-run equilibrium in the sense defined above, the current rate of unemployment is equal to the NAIRU. Thus, we can solve for the equilibrium level of unemployment u^* as:

$$u_t^* = d_0 + d_1 t + d_2 lu_t + d_3 i_t + d_4 k_t + d_5 c_t \tag{20}$$

where $d_0 = \dfrac{\Omega_0 - \ln \overline{m} - \overline{\omega}}{\omega_u}$, $d_1 = \dfrac{\Omega_1 - \lambda_2}{\omega_u}$, $d_2 = \dfrac{-\omega_{lu}}{\omega_u} \succ 0$, $d_3 = \dfrac{-\omega_I}{\omega_u} \succ 0$, $d_4 = \dfrac{\Omega_2}{\omega_u} \prec 0$,

$d_5 = \dfrac{-\phi}{\omega_u}$ and the sign of d_0, d_1, and d_5 is *a priori* ambiguous.

Expression (20) above is subsequently used as a benchmark for estimation purposes in the empirical work presented below. It tells us that the NAIRU depends positively on long-term unemployment and the real price of imports and negatively on

the capital-output ratio. It also tells us that it will shift over time in a way determined by the sign of d_1 if wage aspirations do not grow in line with total factor productivity and that changes in capacity utilization will affect the NAIRU in an unknown way *a priori* (it will depend on the actual sign of d_5). Finally, it tells us that the NAIRU depends on all the factors embedded in d_0, such as the parameters of the production function, the wage, and the mark-up equation. To the extent that the NAIRU is a negative function of the capital-output ratio and the latter is itself a positive function of the ratio of the real wage to the rental price of capital services, the model predicts a positive relationship between real interest rates and the NAIRU. We turn to the testing of this hypothesis.

IV   EMPIRICAL INVESTIGATION

The first part of this section discusses the data set utilized. This is followed by the presentation of the results derived from the estimation of the cointegrating relationships.

Variable Definition and Data

The model is estimated using quarterly and seasonally adjusted data for the period 1964q2 to 2003q1 for the United States. This was dictated by data availability, since data on the NAIRU are only available from 1964q2 onwards. Most data was obtained from the EcoWin Pro database, which collects data from the U.S. Bureau of Economic Analysis (BEA) NIPA Tables, the Bureau of Labor Statistics (BLS), and the U.S. Federal Reserve (Fed). Data on the NAIRU for the United States was obtained from the OECD. Next, we define and discuss briefly each variable used in the estimations. Capital stock (K) is defined as the stock of private non-residential fixed assets at current prices (BEA, NIPA Table 4.1) deflated by a price index for gross private non-residential domestic investment (BEA, NIPA Table 1.6.4, Line 23). Annual data were converted into quarterly data utilizing the facility incorporated in the EcoWin database. Real output (Y) is defined as Real Gross Domestic Product (BEA, NIPA Table 1.1.3). Capacity utilization (C) is defined as capacity utilization for the manufacturing sector and the original source is the U.S. Fed. We used the manufacturing sector index rather than the general one because the latter is not available before 1967. We define long-term unemployment (LU) as the ratio of long-term unemployed to the total number of unemployed. In turn, the BLS defines long-term unemployed as those workers who

13

have been unemployed for at least 27 weeks. The real price of imports (I) is calculated as the ratio of a price index of import goods measured in domestic currency divided by the consumer price index. Both price indexes were obtained from the EcoWin database. Finally, data on the NAIRU (see Figure 1 below) were obtained from the OECD. Details on the construction of the time series can be found in Richardson et al. (2000).

Econometric Specification

The aim of the empirical part of the paper is to test whether there is a long-term equilibrium relationship between the NAIRU and the aggregate capital-output ratio once we control for the rest of variables embedded in the theoretical model. For that purpose, we use a cointegrated VAR model (for a presentation of this methodology, see Johansen 1995). This maximum likelihood methodology uses as a benchmark a VAR of order k containing p variables or:

$$X_t = \phi_1 X_{t-1} + \phi_2 X_{t-2} + ... + \phi_p X_{t-k} + \Phi D_t + \varepsilon_t \qquad t = 1,...,T \qquad (21)$$

where ε_t are independently distributed Gaussian errors with zero mean and variance Ω and D_t captures the deterministic terms. Matrix X_t contains (px1) time series. This $VAR(k)$ model can be expressed as an error correction model or:

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + ... + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-1} + \Phi D_t + \varepsilon_t \qquad t = 1,...,T \qquad (22)$$

where $\Pi = \sum_{i=1}^{k} \phi_i - I$, $\Gamma_i = -\sum_{j=i+1}^{k} \phi_j$.

If $X_t \sim I(1)$ and $\Delta X_t \sim I(0)$, matrix Π cannot have full rank. If rank r is equal to zero, then this means that there is no cointegration relationship among the variables. If $r \prec p$, then matrix Π can be decomposed as $\Pi = \alpha\beta'$ where α and β are each a $(p \times r)$ matrix containing parameters. In particular, matrix β contains the r cointegration relationships and matrix α contains the adjustment parameters. As for the specification of the deterministic terms, we assume the existence of a linear time trend that is restricted to the cointegration space, i.e., if $\Phi D_t = \mu_0 + \mu_1 t$ and the parameters μ_i can themselves be decomposed as $\mu_i = \alpha\rho_i + \alpha_\perp\gamma_i$ for $i = 0,1$, we then assume that $\Phi D_t = \alpha\rho_0 + \alpha_\perp\gamma_0 + \alpha\rho_1 t$.

The choice of this specification stems from the fact that the theoretical model expounded above allows for the possibility that such a time trend exists whenever the wage aspirations of workers do not grow over time in line with the growth of total factor productivity. In addition, a preliminary inspection of the time path of variables

such as the NAIRU, the aggregate capital-output ratio, and the real price of imports suggests that they may, indeed, possess a time trend.

The real price of imports is an exogenous variable in our analysis. Likewise, our theoretical model suggests that the aggregate capital-output ratio is, at least partly, exogenous. This stems from the fact that the latter is determined by technology, expectations, and the ratio of the rental price of capital to the price of labor. In turn, the rental price of capital is strongly influenced by the stance of monetary policy. This allows for the possibility that the capital-output ratio exhibits some degree of endogeneity insofar as changes in the NAIRU brought about by changes in any variable of the model other than the capital-output ratio may lead to changes in the stance of monetary policy provided the central bank targets inflation, which has arguably been the case of the U.S. Fed in the period considered. However, the exogeneity tests performed showed unambiguously that the aggregate capital-output ratio is exogenous (see Table 5 below).

Using a partial model, i.e., a model which imposes from the start the condition that some variables are weakly exogenous, has a number of advantages. First, it allows us to reduce the dimension of the system. Second, it makes the interpretation of results easier. Following Harbo et al. (1998), let us assume that X_t can be decomposed into Y_t and Z_t, and that the following matrices can, in turn, be decomposed as:

$$\alpha = \begin{bmatrix} \alpha_1 \\ \alpha_2 \end{bmatrix} \quad , \quad \Gamma_j = \begin{bmatrix} \Gamma_{1j} \\ \Gamma_{2j} \end{bmatrix}, \quad \Phi = \begin{bmatrix} \Phi_1 \\ \Phi_2 \end{bmatrix} \tag{23}$$

A partial model for ΔY_t can be defined as:

$$\Delta Y_t = \omega \Delta Z_t + (\alpha_1 - \omega \alpha_2)\beta^{'} X_{t-1} + \sum_{i=1}^{k-1} \widetilde{\Gamma}_{1i} \Delta X_{t-i} + \widetilde{\Phi}_1 D_t + \widetilde{\varepsilon}_{1t} \tag{24}$$

where $\omega = \Omega_{12}\Omega_{22}^{-1}$, $\widetilde{\Gamma}_{1i} = \Gamma_{1i} - \omega \Gamma_{2i}$, $\widetilde{\Phi}_1 = \Phi_1 - \omega \Phi_2$, $\widetilde{\varepsilon}_{1t} = \varepsilon_{1t} - \omega \varepsilon_{2t}$.

In turn, the marginal model for ΔZ_t is represented by:

$$\Delta Z_t = \alpha_2 \beta^{'} X_{t-1} + \sum_{i=1}^{k-1} \Gamma_{2i} \Delta X_{t-i} + \Phi_2 D_t + \varepsilon_{2t} \tag{25}$$

If $\alpha_2 = 0$, then the variables in Z_t are weakly exogenous. This means that they do not possess an error correction term or, alternatively, that there is no information about β in the marginal model. Partial models pose some problems when trying to make inferences about the cointegrating rank. This is because the limit distribution depends on both the deterministic terms and the nuisance parameters. Notwithstanding

this problem, Harbo et al. (1998) show that when we consider a time trend restricted to the cointegration vector, the limit distribution is free of nuisance parameters which, in turn, makes inferences about the cointegrating rank possible as long as: (i) the system is generated by a VAR model so variables do not exhibit an order of integration higher than one and, (ii) variables in Z_t are weakly exogenous so that the cointegrating rank is equal or less that the dimension of Y_t.

Tables 1 to 3 below show the tests utilized to determine the order of integration of time series. In addition to ADF unit root tests, we implemented the tests proposed in Ng and Perron[14] (2001) and Elliott, Rothenberg, and Stock (1996), as well as the stationarity tests elaborated in Kwiatkowski et al. (1992). Results strongly suggest that the NAIRU, the aggregate capital-output ratio, and the real price of imports are all integrated or order 1 or I(1). By contrast, results for long-term unemployment and capacity utilization suggest that they both could be I(0), albeit these results depend, to some extent, on the type of test implemented and the specification chosen for the data generating process.[15]

Table 4 below provides the information related to the determination of the cointegration rank for the whole system. In addition to the trace test, we used information provided by the roots of the companion matrix and the t statistics associated to the adjustment coefficients[16]. The number of lags of the VAR model that minimizes the Schwarz and Hannan-Quinn criteria is three. In turn, the cointegration rank appears to be equal to two[17] and, for this cointegration rank, the tests implemented indicate that we cannot reject the hypothesis that the real price of imports and the capital-output ratio are weakly exogenous in the complete model (see Table 5 below). As a result, we opted

---

[14] These authors develop the *M tests* to allow for GLS detrending of the data and, in addition, propose the use of a Modified Information Criteria for selecting the number of lags in the context of unit root tests. In particular, they argue that commonly used information criteria tend to select a low number of lags which, according to them, is inappropriate when the errors contain a moving-average root that is close to –1. As a result, in this study we used the Modified Akaike Information Criteria (MAIC).

[15] We also ran unit root tests with the series expressed in first differences and in all cases we found that we could reject the hypothesis that they were I(2).

[16] Of course, we also took into account the figures displaying the cointegration relations. In addition, we also provide the trace test statistics after having implemented the Barlett corrections (see Johansen 2002).

[17] If r = 2 we have that the result of the trace test after implementing the Barlett´s corrections is slightly lower than the critical value at the 5% significance level. Nevertheless, the visual analysis of the figure associated to the second cointegration relation and the fact that the adjustment coefficients of both the NAIRU and capacity utilization are significant led us to assume that r = 2.

to consider them as such from the start and performed the empirical analysis using a partial system.

The Estimated Long-Run Relations

Table 6 below presents results related to the determination of the cointegrating rank of the partial system $Y_t^{'} = (u_t^{*}, lu_t, c_t)$, $Z_t^{'} = (k_t, i_t)$. As pointed out above, our VAR(3) model contains a linear time trend that is restricted to the cointegration space. Results in Table 6 suggest that the cointegrating rank for the partial system is two as for the full system.[18] We implemented a number of tests in order to detect problems associated with possible model misspecification. Results suggest that there is neither autocorrelation in the residuals of the unrestricted VAR, nor heterokedasticity. Unfortunately, residuals do not seem to follow a normal distribution. Notwithstanding this normality problem, Gonzalo (1994) argues that the maximum likelihood principle in an error correction model exhibits better properties than other methodologies when used for estimating cointegration relations, even if errors are non-normally distributed.[19]

Table 7 below shows results for the exclusion and stationarity tests. We performed an analysis of their sensitivity with respect to the cointegrating rank for *r=1* and *r=2*. Results lend strong support to the claim that none of the variables analyzed is stationary[20] and, in addition, long-term unemployment is not included in any of the two cointegration relations. Table 8 presents results for the weak exogeneity tests. It can be seen that none of the variables can be deemed exogenous. Table 9 shows the identified structure for the long-run relations. The overidentifying restrictions are accepted with a p-value equal to 2.16 according to the test $\chi^2(2) = 0.34$. The first relation—see expression (26) below—tells us that there exists a cointegration relation encompassing the NAIRU, the aggregate capital-output ratio, capacity utilization, and the real price of imports (t-ratios are shown in parentheses):

---

[18] The visual analysis of the figures associated with the cointegration relations in the partial model made us discard the possibility that the cointegration rank be equal to 3.

[19] This author argues that the methodology proposed by Johansen represents a particular case of reduced rank simultaneous least squares where the condition that residuals must follow a given type of distribution is not imposed.

[20] The stationarity tests inserted in that figure were carried out under the assumption that the cointegration vector includes a time trend. However, even if the time trend is assumed away, we can still reject the hypothesis that the variables are stationary.

$$u_t^* = -1.18c_t - 0.46k_t + 0.24i_t \qquad (26)$$

$$(7.94) \quad (9.16) \quad (7.42)$$

The negative coefficient of the aggregate capital-output ratio means that an increase in its value will lead to a fall in equilibrium unemployment and vice-versa. Likewise, the negative coefficient attached to capacity utilization can be interpreted in the context of our model as the outcome of the counter-cyclical behavior of firms' average mark-ups. However, other possibilities cannot be discarded. As an example, this negative coefficient may be the result of the counter-cyclical behavior of firms' total average costs or of the impact on the NAIRU of changes in actual unemployment occurring through hysteresis effects unrelated to long-term unemployment, as we measured above. By contrast, an increase in the real price of imports raises equilibrium unemployment and vice-versa. Figure 1 (below) illustrates how the two oil crises (1973 and 1979), especially the first one, affected the evolution of the NAIRU. As for the time trend, we cannot reject the hypothesis that it is not part of the cointegration relation. This suggests that, at least for the period considered, wage aspirations grew over time, roughly in line with total factor productivity but, crucially, they did not adjust fast enough to offset the impact on the marginal productivity of labor of changes in the aggregate capital-output ratio. The second relation tells us that capacity utilization is cointegrated with the capital-output ratio and the real price of imports or:

$$c_t = -0.80k_t - 0.05i_t + 0.00t \qquad (27)$$

$$(8.75) \quad (2.69) \quad (6.98)$$

The negative sign of the aggregate capital-output ratio stems from the fact that aggregate output and capital appear respectively in the numerator and denominator of capacity utilization. The presence of a significant, albeit negligible, linear time trend may be due to the presence of a time trend in the aggregate capital-output ratio in the period considered. Finally, the negative relation between capacity utilization and the real price of imports stems from the fact that an increase in the latter brings about a rise in the rate of inflation and, consequently, a tightening of the stance of monetary policy that will, in turn, reduce aggregate capacity utilization and vice-versa.

To analyze the constancy of the estimated long-run parameters and the possible presence of structural change in the estimated relationships, we used the fluctuation test

of the eigenvalues and the LM-type test developed in Hansen and Johansen (1999).
Since there exists a relationship among the eigenvalues, the adjustment coefficients, and
the cointegrating vectors, the analysis of the time path of the estimated eigenvalues can
be used to test for the constancy of the estimated long-run parameters. As can be
observed in Figure 2, the null hypothesis of constancy cannot be rejected either for any
of the two eigenvalues individually or for their sum because the values of the tests are
below the 5% critical value. Figure 3 shows the LM-type test, according to which the
estimated long-run parameters are constant over the sample period since the maximum
value of the Q_t test is lower than the 5% critical value. We also used the test developed
in Hansen and Johansen (1993) to test the hypothesis that $\tilde{\beta} \in sp(\beta_\tau)$ for
$\tau = 1970:3 - 2003:1$, where $\tilde{\beta}$ is the estimate of the long-run parameters for the full
sample 1964:2-2003:1 when over-identifying restrictions are imposed. In Figure 4, we
show the test statistic we obtained. This test has been scaled using their 5% critical
value, so that the values exceeding 1 indicate the inconstancy of $\tilde{\beta}$. According to this
test, we cannot reject the null hypothesis of constancy under the *R* representation (i.e.,
the short run parameters are fixed at their full sample values), but not under the *Z*
representation (when the short run parameters are estimated recursively). However,
Hansen and Johansen (1993) point out that the *R* representation is more appropriate than
the *Z* one when the purpose is to analyze long-run stability.

The Short-Run Adjustment

Table 10 below presents results relating to the short-run adjustment dynamics of the
model. Those coefficients that are significantly different from zero are heavily outlined,
i.e., those for which the t-ratio is larger than two. We can see that the NAIRU is
strongly equilibrium correcting to *ecm1* and significantly overshoots *vis á vis* capacity
utilization (*ecm2*), i.e., the NAIRU raises when capacity utilization is above its
stationary state value and vice-versa. We may note that the dynamics of the NAIRU in a
given period are positively affected by its evolution in the previous period which
imbues this variable with a good deal of inertia. By contrast, larger increases in the
capital-output ratio in previous periods lead to lower increases in the NAIRU in the
current period, so the link between these two variables exists in the short-run as well as
in the long-run.

As for long-term unemployment, Table 10 shows that its value falls when the NAIRU exceeds its stationary state value and vice-versa. This could be easily accounted for if, as in the theoretical model, these two variables were positively related. However, the empirical analysis yielded that long-term unemployment can be excluded from both cointegration relations. Finally, capacity utilization is equilibrium correcting to *ecm2*. We may also note that its current period dynamics are positively influenced by its evolution in the preceding period and that an increase in the capital-output ratio reduces its current increase and vice-versa.

V   SUMMARY AND CONCLUSIONS

The purpose of this paper was to examine the proposition that capital stock relative to aggregate output has been an important variable in the determination of unemployment in the United States economy, at least in the last four decades. This proposition runs against conventional wisdom in the field which holds that persistent unemployment is mainly the result of a number of labor market rigidities. This position is well exemplified in the influential work of Layard, Nickell, and Jackman (1991) where they impose cross-equation restrictions which ensure that the rate of unemployment is unaffected by technical progress and changes in the aggregate capital-labor ratio. The crucial assumption lying at the core of this position is that changes in productivity, whatever their cause, are automatically reflected in equivalent changes in real wages, so the NAIRU is unaffected by the former. Yet, a number of authors have recently claimed that real wage aspirations are tied down to wage and productivity growth in the past so it may take some time before they fully adjust to changes in productivity growth. If so, the NAIRU will be affected in the meantime. One possible, but still unexplored, source of changes in productivity growth is changes in the aggregate capital-output ratio. Whether or not the latter leads to significant changes in the NAIRU is ultimately an empirical question.

This paper presented new empirical evidence obtained from the application of the cointegrated VAR methodology to U.S. time-series data, which lends strong support to the claim that the aggregate capital-output ratio, the real price of imports, and capacity utilization were significant determinants of the NAIRU in the U.S. economy in the last four decades. In particular, increases in the aggregate capital-output ratio and capacity utilization and decreases in the real price of imports were found to be

20

associated with significant decreases in the NAIRU. The same evidence showed that we cannot reject the hypotheses that technical progress and changes in long-term unemployment did not affect the NAIRU in the same period. Contrary to conventional wisdom, this evidence suggests that, insofar as the aggregate capital-output ratio is affected by changes in real interest rates, then the stance of monetary policy has a considerable impact on the NAIRU. In particular, a policy of using short-term interest rates to control inflation may well be successful in the short run by reducing aggregate demand but it has also negative implications for supply, bringing closer the point where the economy runs into inflationary problems. Thus, interest rate policy in itself is no panacea. It does have supply-side consequences in that it only delivers low inflation through low growth and high unemployment. To finish off this section, we should like to conclude that, given the absence of empirical studies of this type for other OECD economies, any final verdict on the theoretical relevance of our results will have to wait for the emergence of further evidence on the empirical significance of capital deepening as a determinant of the NAIRU for other economies. Hopefully, our research agenda will find some room for this task.

REFERENCES

Akerlof, G. and J. Yellen. 1990. "The Fair Wage-Effort Hypothesis and Unemployment." *Quarterly Journal of Economics* 105(2): 255–284.

Allen, C. and J. Nixon. 1997. "Two Concepts of the NAIRU." in C. Allen and S. Hall (eds.) *Macroeconomic Modelling in a Changing World*. Chichester, U.K.: John Wiley & Sons Ltd.

Arestis, P. and I. Biefang-Frisancho Mariscal. 1998. "Capital shortages and asymmetries in UK unemployment." *Structural Change and Economic Dynamics* 9: 189–204.

————. 2000. "Capital Stock, Unemployment and Wages in the UK and Germany." *Scottish Journal of Political Economy* 47(5): 487-503.

Ball, L. 1997. "Disinflation and the NAIRU" in C.D. Romer and D.H. Romer (eds.), *Reducing Inflation: Motivation and Strategy*. Chicago: Chicago University Press.

————. 1999. "Aggregate Demand and Long-Run Unemployment." *Brookings Papers on Economic Activity* 2: 189–236.

Ball, L. and N. G. Mankiw. 2002. "The NAIRU in Theory and Practice." *Journal of Economic Perspectives* 16(4): 115–136.

Bean, C. 1989. "Capital Shortage." *Economic Policy* April: 12–53.

Bibow, J. 2001. "Making EMU work: Some lessons from the 1990s." *International Review of Applied Economics* 15(3): 233–259.

————. 2005. "Germany in Crisis: The Unification Challenge, Macroeconomic Policy Shocks and Traditions, and EMU." *International Review of Applied Economics* 19(1): 29–50.

Bils, M. 1987. "The Cyclical Behavior of Marginal Cost and Price." *American Economic Review* 77: 838–55.

Blanchard, O., R. Dornbusch, J.H. Dréze, H. Giersch, R. Layard, and M. Monti. 1986. "Employment and Growth in Europe: A Two-Handed Approach, 1985 Report of the CEPS Macroeconomic Policy Group." in O. Blanchard et al., *Restoring Europe Prosperity, Macroeconomic Papers from the Centre for European Policy Studies*. London and Cambridge (Mass.): MIT Press.

Blanchard, O.J. and L. H. Summers. 1987. "Hysteresis in Unemployment." *European Economic Review* 31(1/2): 288–295.

Blanchard, O.J. and D. Quah. 1989. "The Dynamic Effects of Aggregate Demand and Supply Disturbances." *American Economic Review* 79(4): 655–73.

Blanchard, O. and L.F. Katz. 1997. "What We Know and Do Not Know About the Natural Rate of Unemployment." *Journal of Economic Perspectives* 11(1): 51–72.

Blanchard, O. and J. Wolfers. 2000. "The role of shocks and institutions in the rise of European unemployment: The aggregate evidence." *Economic Journal* 110(462): C1–C33.

Blinder, A. 2000. "The Internet and the New Economy." *Brookings Institution Policy Brief* 60, June.

Chirinko, R.S. and S.M. Fazzari. 1994. "Economic Fluctuations, Market Power and Returns to Scale: Evidence from Firm Level Data." *Journal of Applied Econometrics* 9: 47-69.

Cross, R. (ed.). 1995. *The Natural Rate of Unemployment.* Cambridge: Cambridge University Press.

Elliott, G., T. J. Rothenberg, and J. H. Stock. 1996. "Efficient tests for an autoregressive unit root." *Econometrica* 64: 813–836.

Franz, W. 1987. "Hysteresis, Persistence, and the NAIRU: An Empirical Analysis for the Federal Republic of Germany." in Layard, R. and Calmfors, L. (eds.) *The Fight Against Unemployment: Macroeconomic Papers from the Centre for European Studies*. Cambridge, MA: MIT Press.

————. 1990. "Hysteresis in Economic Relationships: An Overview." *Empirical Economics* 15: 109–25.

Fortin, P. 1996. "The Great Canadian Slump." *Canadian Journal of Economics* 29(4): 761–787.

Galeotti, M. and F. Schiantarelli. 1998. "The Cyclicality of Markups in a Model with Adjustment Costs: Econometric Evidence for U.S. Industry." *Oxford Bulletin of Economics and Statistics* 60(2): 121–142.

Galí, J. 1992. "How Well Does the IS-LM Model Fit Postwar U.S. Data?" *Quarterly Journal of Economics* 107(2): 709–738.

Gonzalo, J. 1994. "Comparisons of five alternative methods of estimating long-run equilibrium relations." *Journal of Econometrics* 60: 203–234.

Greenspan, A. 1997. "The Federal Reserve´s Semiannual Monetary Policy Report, Testimony before the U.S. House Committee on Banking, Housing and Urban Affairs." February 26.

Grubb, D., R. Jackman and R. Layard. 1982. "Causes of the Current Stagflation." *Review of Economic Studies* Special Issue 49(5): 707–30.

Hall, R. E. 1986. "Market Structure and Macroeconomic Fluctuations." *Brookings Papers on Economic Activity* 2: 285–322.

Hansen, H. and S. Johansen. 1993. "Recursive estimation in cointegrated VAR-models." Preprint n.1, Institute of Mathematical Statistics, University of Copenhagen, Copenhagen.

————. 1999. "Some tests for parameter constancy in cointegrated VAR-models." *Econometrics Journal* 2: 306-333.

Harbo, I., S. Johansen, B. Nielsen, and A. Rahbek. 1998. "Asymptotic inference on cointegrating rank in partial systems." *Journal of Business and Economic Statistics* 16(4): 388–399.

Hein, E. and A. Truger. 2005. "What Ever Happened to Germany? Is the Decline of the Former European Key Currency Country Caused by Structural Sclerosis or by Macroeconomic Mismanagement?" *International Review of Applied Economics* 19(1): 3–28.

Johansen, S. 1995. *Likelihood-based inference in cointegrated vector autoregressive models*. Oxford: Oxford University Press.

————. 2002. "A small sample correction for the test of cointegrating rank in the Vector Autoregressive Model." *Econometrica* 70(5): 1929–1961.

Kwiatkowski, D., P. C. B. Phillips, P. Schmidt, and Y. Shin. 1992. "Testing the null hypothesis of stationary against the alternative of a unit root." *Journal of Econometrics* 54: 159–178.

Layard, R., S. Nickell, and R. Jackman. 1991. *Unemployment: Macroeconomic Performance and the Labour Market.* Oxford: Oxford University Press.

Lindbeck, A. and D. Snower. 1988. *The insider-outsider theory of employment and unemployment*. Cambridge, MA: MIT Press.

Lindbeck, A. 1995. "Hazardous Welfare-State Dynamics." *American Economic Review* 85(2): 9–15.

MacKinnon, J. G., A.A. Haug, and L. Michelis. 1999. "Numerical distribution functions of likelihood ratio tests for cointegration." *Journal of Applied Econometrics* 14: 563–577.

Malinvaud, E. 1986. "The Rise of Unemployment in France." *Economica* 53(210): S197-S217.

Modigliani, F., M. Monti, J.H. Dréze, H. Giersch, and R. Layard. 1987. "Reducing Unemployment in Europe: The Role of Capital Formation." in R. Layard and L. Calmfors (eds.) *The Fight Against Unemployment.* Cambridge, MA: MIT Press.

Ng, S. and P. Perron. 2001. "Lag length selection and the construction of unit root tests with good size and power." *Econometrica* 69(6): 1519–1554.

Nickell, S. 1997. "Unemployment and Labor Market Rigidities: Europe versus North America." *Journal of Economic Perspectives* 11(3): 55–74.

Oswald, A. 1986. "Is Wage Rigidity Caused by Layoffs by Seniority?" in Beckerman, Wilfred, (eds.) *Wage Rigidity and Unemployment.* Baltimore, MD.: John Hopkins University Press.

Richardson, P., L. Boone, C. Giorno, M. Meacci, D. Rae, and D. Turner. 2000. "The Concept, Policy Use and Measurement of Structural Unemployment: Estimating a Time Varying NAIRU Across 21 OECD Countries." OECD Economics Department Working Paper no. 250, ECO/WKP (2000)23.

Romer, C.D. and D. Romer. 1989. "Does Monetary Policy Matter? A New Test in the Spirit of Friedman and Schwartz." *NBER Macroeconomics Annual* 4: 121–170.

Rotemberg, J.J. and M. Woodford. 1991. "Markups and the Business Cycle." *NBER Macroeconomics Annual.* Cambridge, MA: MIT Press.

Rowthorn, R. 1977. "Conflict, inflation and money" *Cambridge Journal of Economics* 1(3): 215–239.

————. 1999. "Unemployment, wage bargaining and capital-labour substitution." *Cambridge Journal of Economics* 23(4): 413–26.

Sachs, J. 1987. "High Unemployment in Europe: Diagnosis and Policy Implications." in Claes-Henric Siven (ed.) *Unemployment in Europe: Analysis and Policy Issues.* Stockholm: Timbro.

Sawyer, M.C. 2002. "The NAIRU, Aggregate Demand and Investment." *Metro-economica* 53(1): 66–94.

Shapiro, C. and J. Stiglitz. 1984. "Equilibrium Unemployment as a Discipline Device." *American Economic Review* 74(June): 433–44.

Siebert, H. 1997. "Labor Market Rigidities: At the Root of Unemployment in Europe." *Journal of Economic Perspectives* 11(3): 37–54.

Sneessens, H.R. and J.H. Dreze. 1986. "A Discussion of Belgian Unemployment, Combining Traditional Concepts and Disequilibrium Econometrics." *Economica* 53(210): S89-S119.

Stiglitz, J. 1997. "Reflections on the Natural Rate Hypothesis." *Journal of Economic Perspectives* 11(1): 3–10.

Stockhammer, E. 2004. "Explaining Unemployment: Testing the NAIRU Hypothesis and a Keynesian Approach" *International Review of Applied Economics* 18(1): 3–23.

FIGURES AND TABLES

Table 1. Ng-Perron unit root tests

	MZ_α^{GLS}	MZ_t^{GLS}	MSB^{GLS}	MP_T^{GLS}
u^*	0.67	0.31	0.46	19.46
	-3.41	-1.17	0.34	24.29
k	1.84	2.20	1.19	113.59
	-7.50	-1.86	0.25	12.32
lu	-6.54*	-1.70*	0.26*	4.10*
	-14.43*	-2.67*	0.18	6.38*
c	-5.61	-1.44	0.26*	5.03
	-26.39***	-3.61***	0.14***	3.57***
i	-2.00	-0.82	0.41	10.45
	-2.53	-1.00	0.39	31.42

(***), (**) and (*) indicate respectively that we can reject the null hypothesis of a unit root at the 1%, 5% and 10% significance level. We obtained the critical values from Ng and Perron (2001, table 1). In the first row the tests include a constant and in the second row they include a constant and a trend. The criteria used to choose the number of lags was the Modified Akaike Information Criterion (MAIC), with $k_{max} = \text{int}\left(12(T/100)^{1/4}\right) = 13$.

Table 2. ADF and Elliott-Rothenberg-Stock (ERS) unit root tests

	ADF			ERS	
	[1]	[2]	[3]	P_T^μ	P_T^τ
u^*	-0.34	-1.30	-1.27	26.66	31.44
k	0.01	-1.75	-1.93*	295.24	12.25
lu	-1.83	-3.39*	-0.82	4.14*	2.63***
c	-3.23**	-3.61**	-0.91	2.57**	3.51***
i	-0.64	-1.23	-0.57	10.71	41.88

(***), (**) and (*) indicate respectively that we can reject the null hypothesis of a unit root at the 1%, 5% and 10% significance level. In the ADF tests [1] includes a constant, [2] includes a constant and a time trend, and [3] includes neither a constant nor a time trend. The number of lags was chosen according to the Modified Akaike Information Criterion (MAIC), with $k_{max} = \text{int}\left(12(T/100)^{1/4}\right) = 13$.

Table 3. KPSS stationarity tests

	η_μ	η_τ
u^*	0.92***	0.33***
k	1.43***	0.15**
lu	0.74***	0.12*
c	0.45*	0.15**
i	0.52**	0.33***

(***), (**) and (*) indicate respectively that we can reject the null hypothesis of stationarity at the 1%, 5% and 10% significance level. We used critical values provided in Kwiatowski *et al.* (1992, table 1).

Table 4. Determination of the cointegrating rank in the full system

r	$p-r$	Trace test	Trace test*	CV95
0	5	143.88	120.98	88.80
1	4	76.47	62.98	63.87
2	3	41.59	30.93	42.91
3	2	12.41	9.59	25.87
4	1	4.93	4.68	12.52

Modulus of the six largest roots of the companion matrix		The t-values of the α coefficients				
			u^*	lu	c	i	k
$r=1$	1,00; 1,00; 1.00; 1.00; 0.87; 0.79	α_1	-2.44	-7.59	-3.49	1.47	0.36
$r=2$	1,00; 1.00; 1.00; 0.94; 0.79; 0.79	α_2	4.87	-0.37	-3.27	-0.40	0.64
$r=3$	1.00; 1.00; 0.94; 0.92; 0.92; 0.79	α_3	1.81	-2.25	3.81	-2.03	-3.31
$r=4$	1.00; 0.97; 0.97; 0.92; 0.92; 0.79	α_4	0.86	-1.04	0.59	0.74	2.04
$VAR(3)$	0,98; 0,97; 0,97; 0,92; 0,92; 0.79	α_5	0.79	0.29	0.67	1.97	-0.73

Notes: Trace test* represents the values of the trace test after having implemented the Barlett corrections. CV95 represents the critical values of the trace test at the 5% significance level. We obtained these values from MacKinnon *et al.* (1999).

Table 5. Testing for weak exogeneity in the full system

	u^*	lu	c	k	i
$r=1$	2.94	30.05	6.11	0.09	1.76
	[0.09]	[0.00]	[0.00]	[0.76]	[0.18]
$r=2$	7.79	30.36	8.29	0.22	1.84
	[0.00]	[0.00]	[0.00]	[0.89]	[0.40]
$r=3$	24.71	46.16	27.83	8.28	5.32
	[0.00]	[0.00]	[0.00]	[0.04]	[0.15]
$r=4$	26.04	48.50	29.11	10.44	5.54
	[0.00]	[0.00]	[0.00]	[0.03]	[0.24]

Note: LR tests of weak exogeneity follow a $\chi^2(r)$; *p-values* are in brackets.

Table 6. Determination of the cointegrating rank in the partial model

r	$p-r$	Trace test	Trace test*	CV95
0	3	119.42	105.78	56.3
1	2	53.77	46.44	35.5
2	1	19.36	16.72	17.9

Modulus of the six largest roots of the companion matrix		The t-values of the α coefficients			
			u^*	lu	c
$r=1$	1,00; 1,00; 0,84; 0.84; 0.80	α_1	-2.53	-7.79	-3.68
$r=2$	1,00; 0,89; 0,86; 0.79; 0.79	α_2	5.22	-0.28	-3.01
$VAR(3)$	0,92; 0,91; 0,91; 0.79; 0.79	α_3	2.12	-2.30	3.52

Notes: Trace test* indicates the values of the trace test after having implemented the Barlett corrections. CV95 indicates the critical values of the trace test at the 5% significance level. We obtained these values from Harbo *et al.* (1998, table 2).

Table 7. Testing for long-run exclusion and stationarity in the partial model

	Exclusion tests						Stationarity tests		
	u^*	lu	c	k	i	$trend$	u^*	lu	c
$r=1$	5.94 [0.00	1.51 [0.22]	18.40 [0.00]	5.98 [0.00]	2.93 [0.09]	2.23 [0.14]	61.86 [0.00]	52.36 [0.00]	21.43 [0.00]
$r=2$	20.68 [0.00]	2.16 [0.34]	33.23 [0.00]	20.44 [0.00]	15.80 [0.00]	17.03 [0.00]	31.66 [0.00]	23.60 [0.00]	21.42 [0.00]

Note: the LR tests of exclusion follow a $\chi^2(r)$ and the LR tests of stationarity follow a $\chi^2(5-r)$; *p-values* are in brackets.

Table 8. Testing for weak exogeneity in the partial model

	u^*	lu	c
$r=1$	2.91 [0.09]	31.04 [0.00]	7.44 [0.00]
$r=2$	14.64 [0.00]	35.61 [0.00]	12.13 [0.00]

Note: LR tests of weak exogeneity follow a $\chi^2(r)$; *p-values* are in brackets.

Table 9. Identified β structures in the partial model

	u^*	lu	c	k	i	$trend$
$\hat{\beta}_1$	1.00 (NA)	-	1.18 (7.94)	0.46 (9.16)	-0.24 (-7.42)	-
$\hat{\beta}_2$	-	-	1.00 (NA)	0.80 (8.75)	0.05 (2.69)	-0.00 (-6.98)

The LR-test for the restricted model is $\chi^2(2) = 2.16$ with p-value = 0.34; *t-values* are in brackets.

Table 10. The short-run structure in the partial model

|	Δu_t^*	Δlu_t	Δc_t
Δu_{t-1}^*	**0.62**	-7.03	**1.14**
	(7.84)	(-1.58)	**(2.21)**
Δu_{t-2}^*	0.02	-0.27	-0.38
	(0.30)	(-0.06)	(-0.76)
Δlu_{t-1}	-0.00	-0.08	-0.00
	(-0.88)	(-1.15)	(-0.51)
Δlu_{t-2}	-0.00	**-0.49**	**-0.02**
	(-1.12)	**(-7.54)**	**(-3.10)**
Δc_{t-1}	-0.01	-0.22	**0.37**
	(-0.44)	(-0.34)	**(4.92)**
Δc_{t-2}	-0.01	**-2.45**	-0.09
	(-0.46)	**(-3.51)**	(-1.07)
Δk_t	-0.02	-1.21	**-0.79**
	(-0.75)	(-1.06)	**(-6.00)**
Δk_{t-1}	-0.04	0.95	0.16
	(-1.87)	(0.77)	(1.08)
Δk_{t-2}	**-0.05**	-1.06	-0.26
	(-2.18)	(-0.85)	(-1.83)
Δi_t	**0.02**	0.45	0.07
	(2.17)	(0.86)	(1.20)
Δi_{t-1}	0.06	0.13	**0.14**
	(0.55)	(0.22)	**(2.01)**
Δi_{t-2}	**-0.03**	0.60	-0.07
	(-2.89)	(1.11)	(-1.11)
$ecm1_{t-1}$	**-0.03**	**-0.95**	0.04
	(-5.44)	**(-3.55)**	(1.19)
$ecm2_{t-1}$	**0.04**	-0.76	**-0.27**
	(3.80)	(-1.44)	**(-4.41)**
Constant	-0.01	**8.64**	**0.95**
	(-0.25)	**(5.64)**	**(5.37)**

Where

$$ecm1 = u_t^* + 1.18c_t + 0.46k_t - 0.24i_t$$
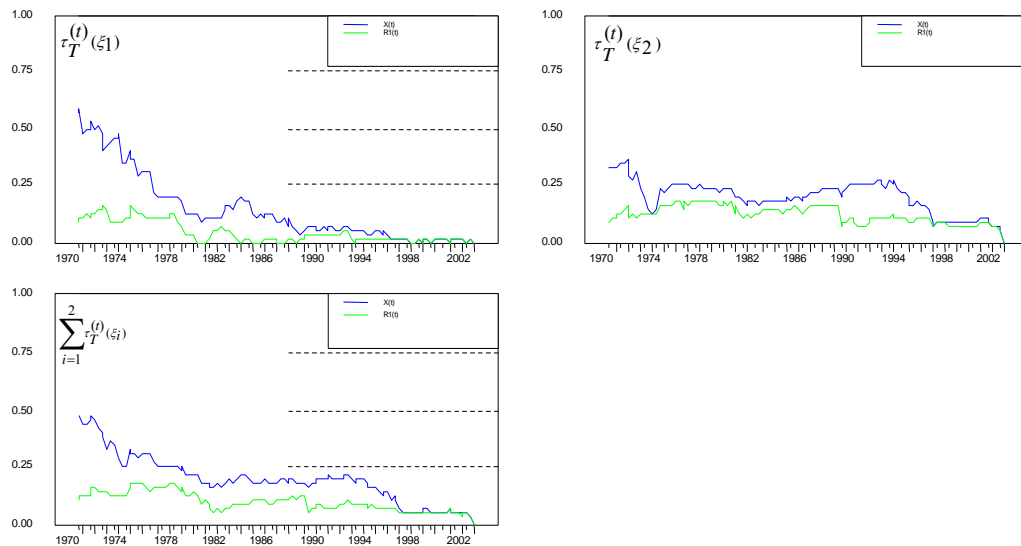
$$ecm2 = c_t + 0.80k_t + 0.05i_t - 0.00t$$

t-values are in brackets.

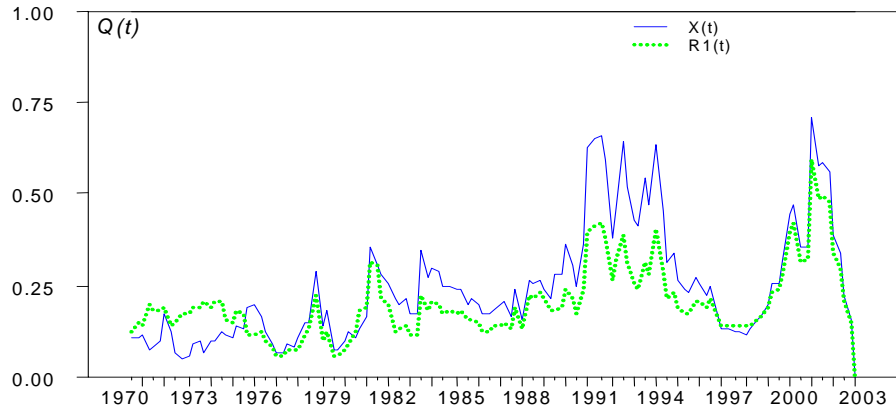**Figure 1. Evolution of NAIRU during the period 1964-2003**

NAIRU (in logarithms)
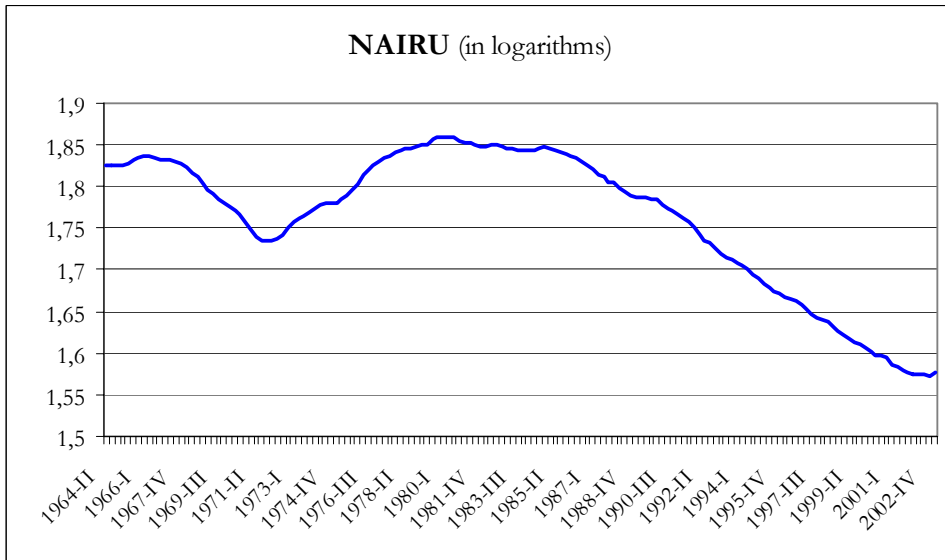
Figure 2. Eigenvalue Fluctuation Test

$\tau_T^{(t)}(\xi 1)$

$\tau_T^{(t)}(\xi 2)$

$\sum_{i=1}^{2}\tau_T^{(t)}(\xi_i)$

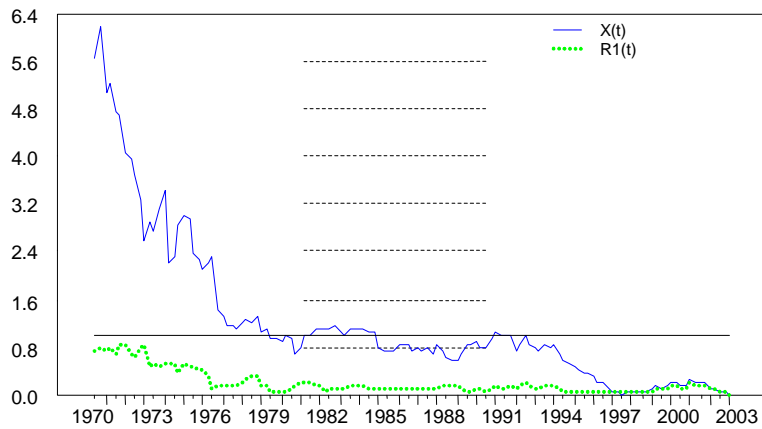Note: the critical value for sup $\tau_T^{(t)}$ at the 5% significance level is 1.36

Figure 3. LM-type test of constancy of beta.

Note: the critical value for sup $Q_T^{(t)}$ at the 5% significance level is 3.16.

Figure 4. Test of Beta(t) = 'Known Beta' (forward estimation)

Note: the critical value at the 5% significance level is 1.