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SEARCHING FOR THE MISSING LINK**

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De conformidad con la base quinta de la convocatoria del Programa de Estímulo a la Investigación, este trabajo ha sido sometido a evaluación externa anónima de especialistas cualificados a fin de contrastar su nivel técnico.

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CAPITAL STOCK AND UNEMPLOYMENT: SEARCHING FOR THE MISSING LINK

By

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Abstract: The purpose of this paper is to examine the proposition that capital stock relative to aggregate output was an important variable in the determination of the U.S. NAIRU in the last four decades. We present 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 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, then the stance of monetary policy is one determinant of the NAIRU.

Key words: Capital-output ratio, cointegrated VAR model, NAIRU, capacity utilization
JEL Classification: E22, E24, C32

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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 US economy. Most of the literature on unemployment has focused on labour market issues, such as wage-fixing institutions, the role of welfare benefits and legal firing costs, and the quality and motivation of the labour force. However, there are other aspects to 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 been recently the focus of attention of several studies that make use of modern time-series analysis (Arestis and Biefang-Frisancho Mariscal, 1998, 2000). But these authors are in a minority and the conventional wisdom seems to be that persistent unemployment is mainly due to labour market rigidities.¹ This position is well exemplified in the influential work of Layard *et al.* (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-labour ratio. This position has become the conventional wisdom in macroeconomic analysis. For instance, Blanchard and Katz (1997, p. 56) argue that ‘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 aggregate capital-output ratio. In fact, unemployment has been on the rise in several European countries for almost two decades and this is put down to a number of factors including the generosity of European welfare systems and other related labour market rigidities in a context of an increasingly competitive world economy (Siebert, 1997). We do not wish to deny that globalisation 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 labour 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 labour productivity thus causing initially a mismatch between wage aspirations and labour 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 labour *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 *et al.* (1991) there are two reasons why changes in the capital-labour 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 But 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. 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 knowledge this possibility has not been tested empirically. Our objective is to test whether changes in the former 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 labour 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 characterising 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 U.S. 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 U.S. 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 labour markets (Cross, 1995; Blanchard and Summers, 1987; Franz, 1987, 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, 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, 1999; Blanchard and Wolfers, 2000; Bibow, 2001, 2005; Hein and Truger, 2005).

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 *et al.* (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-labour 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, p. 86). As Layard *et al.* (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 *et al.*, 1991, p. 369).

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 labour) raises labour 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 *et al.* (1982), Stiglitz (1997) and Blinder (2000), they suggest there is a link between *shifts* in productivity growth and in the NAIRU that may help explain both the rising NAIRU of the 1970s and the falling NAIRU of the 1990s in the U.S. 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 favorably’ (Ball and Mankiw, 2002, pp. 129-130).

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 behaviour 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 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, 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 UK and the USA 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 *et al.* (1991) and drop the cross-equation restriction on the capital-labour ratio in the price and wage equation. They justify this departure from Layard *et al.* (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, p. 138). 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-labour ratio and technical change. Finally, the theoretical connection between the NAIRU and the aggregate capital-labour ratio is also shown in Rowthorn (1999) although his results only apply for the case of a CES production function where the elasticity of substitution between capital and labour 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 the capital stock affects unemployment mainly through its effect on the marginal product of labour. The short-run profit maximizing decision facing the typical firm i consists of maximizing profits

$$\Pi_t^i = p_t^i Y_t^i - w_t^e N_t^i \quad (1)$$

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 labour 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 labour and capital and similar technology. This assumption allows us to simplify substantially firms' price-setting equation since we only need to consider one variable production factor, i.e., labour. Hence, the first-order condition of the optimization problem faced by the typical firm yields the following economy-wide relation:

$$p_t = m_t \cdot \frac{w_t^e}{\Omega_t^N} \quad (2)$$

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

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

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:

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

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 labour or:

$$\Omega_t^N = \gamma_0 e^{\lambda_1 t} a (K_t / N_t)^{1-a} \phi 0 \quad (5)$$

Likewise, the marginal product of capital is:

$$\Omega_t^K = \gamma_0 e^{\lambda_1 t} (1 - a) \cdot (K_t / N_t)^{-a} \phi 0 \quad (6)$$

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

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

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

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

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

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 labour 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, $\lambda_2 \neq 0$ 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 $\omega_\varphi \neq 0$ and $\omega_I \neq 0$ 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 $\varphi_U < 0$, $\varphi_{LU} > 0$.

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 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 labour market, and knowing that finding another job is likely to be difficult, workers 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 labour 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 will vary with labour 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 labour 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 search¹² (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, 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 will be determined, at least partly, by a wide range of idiosyncratic socio-economic 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 roughly in line with labour 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 labour net of income transfers from the government to workers and the discrepancy between the *GDP* deflator and the consumer price index. For reasons we 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. If we insert (11) into (9), assume that function ω is log-linear and take logarithms in (9) we get:

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

where the logarithms of the arguments in ω are denoted by lower-case letters, $\bar{\omega}$ can be interpreted as the ‘aspiration’ (real) wage at the initial period and $\omega_u \geq 0$ and $\omega_{lu} \leq 0$.

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 cyclicity of mark-ups. In particular, we hypothesize that mark-ups are set according to:

$$m_t = \bar{m} \cdot C_t^\phi \quad (13)$$

where C_t is aggregate capacity utilization, \bar{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 $\phi > 0$ then increases in capacity utilisation will raise the NAIRU and vice-versa. If we take logarithms in (13) we get:

$$\ln m_t = \ln \bar{m} + \phi \cdot c_t \quad (14)$$

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

$$\ln p_t - \ln w_t^e = \ln m_t - \ln \Omega_t^N \quad (15)$$

and
$$\ln \Omega_t^N = \Omega_0 + \Omega_1 \cdot t + \Omega_2 k_t \quad (16)$$

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

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

$$\ln p_t - \ln w_t^e = \ln \bar{m} + \phi c_t - \Omega_0 - \Omega_1 t - \Omega_2 k_t \quad (17)$$

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 $p_t = p_t^e$ and hence (12) and (17) become respectively:

$$\ln w_t - \ln p_t = \bar{\omega} + \omega_u u_t + \omega_{lu} l u_t + \omega_i i_t + \lambda_2 t \quad (18)$$

$$\ln p_t - \ln w_t = \ln \bar{m} + \phi c_t - \Omega_0 - \Omega_1 t - \Omega_2 k_t \quad (19)$$

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:

$$u_t^* = d_0 + d_1 t + d_2 l u_t + d_3 i_t + d_4 k_t + d_5 c_t \quad (20)$$

where $d_0 = \frac{\Omega_0 - \ln \bar{m} - \bar{\omega}}{\omega_u}$, $d_1 = \frac{\Omega_1 - \lambda_2}{\omega_u}$, $d_2 = \frac{-\omega_{lu}}{\omega_u} \phi$, $d_3 = \frac{-\omega_i}{\omega_u} \phi$, $d_4 = \frac{\Omega_2}{\omega_u} \pi$, $d_5 = \frac{-\phi}{\omega_u}$ 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 hence 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 utilised. 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 U.S. 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 U.S. 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 utilising the facility incorporated in the Ecwin 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 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 Ecwin 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:

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

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 $(p \times 1)$ time series. This VAR(k) model can be expressed as an error correction model or:

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

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 $r < p$, 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 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 labour. 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:

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

A partial model for ΔY_t can be defined as:

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

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

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

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

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 than 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 Perron¹⁴ (2001) and Elliott *et al.* (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 process¹⁵.

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 coefficients¹⁶. The number of lags of the VAR model that minimises the Schwarz and Hannan-Quinn criteria is three. In turn, the cointegration rank appears to be equal to two¹⁷ 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 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 to 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 analysed is stationary²⁰ 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 or (t-ratios are shown in parentheses):

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

(7.94) (9.16) (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 behaviour 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 behaviour 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 it 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, seem to explain properly the evolution on 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 overtime roughly in line with total factor productivity but, crucially, they did not adjust fast enough to offset the impact on the marginal productivity of labour 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:

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

(8.75) (2.69) (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 is 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 monetary policy that will, in turn, reduce aggregate capacity utilization and vice-versa.

To analyse 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 analyse 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 we have that 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, higher 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 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 increase and vice-versa.

V SUMMARY AND CONCLUSIONS

The purpose of this paper was to examine the proposition that capital formation relative to aggregate output has been an important variable in the determination of unemployment in the US 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 labour market rigidities. This position is well exemplified in the influential work of Layard *et al.* (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-labour ratio. The crucial assumption lying at the core of the conventional wisdom 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 has 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 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 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 significant 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 real

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.

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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 $k_{\max} = \text{int}(12(T/100)^{1/4}) = 13$.

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 $k_{\max} = \text{int}(12(T/100)^{1/4}) = 13$.

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 [0.09]	30.05 [0.00]	6.11 [0.00]	0.09 [0.76]	1.76 [0.18]
$r=2$	7.79 [0.00]	30.36 [0.00]	8.29 [0.00]	0.22 [0.89]	1.84 [0.40]
$r=3$	24.71 [0.00]	46.16 [0.00]	27.83 [0.00]	8.28 [0.04]	5.32 [0.15]
$r=4$	26.04 [0.00]	48.50 [0.00]	29.11 [0.00]	10.44 [0.03]	5.54 [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^*	$\Delta l u_t$	Δc_t
Δu_{t-1}^*	0.62 (7.84)	-7.03 (-1.58)	1.14 (2.21)
Δu_{t-2}^*	0.02 (0.30)	-0.27 (-0.06)	-0.38 (-0.76)
$\Delta l u_{t-1}$	-0.00 (-0.88)	-0.08 (-1.15)	-0.00 (-0.51)
$\Delta l u_{t-2}$	-0.00 (-1.12)	-0.49 (-7.54)	-0.02 (-3.10)
Δc_{t-1}	-0.01 (-0.44)	-0.22 (-0.34)	0.37 (4.92)
Δc_{t-2}	-0.01 (-0.46)	-2.45 (-3.51)	-0.09 (-1.07)
Δk_t	-0.02 (-0.75)	-1.21 (-1.06)	-0.79 (-6.00)
Δk_{t-1}	-0.04 (-1.87)	0.95 (0.77)	0.16 (1.08)
Δk_{t-2}	-0.05 (-2.18)	-1.06 (-0.85)	-0.26 (-1.83)
Δi_t	0.02 (2.17)	0.45 (0.86)	0.07 (1.20)
Δi_{t-1}	0.06 (0.55)	0.13 (0.22)	0.14 (2.01)
Δi_{t-2}	-0.03 (-2.89)	0.60 (1.11)	-0.07 (-1.11)
$ecm1_{t-1}$	-0.03 (-5.44)	-0.95 (-3.55)	0.04 (1.19)
$ecm2_{t-1}$	0.04 (3.80)	-0.76 (-1.44)	-0.27 (-4.41)
Constant	-0.01 (-0.25)	8.64 (5.64)	0.95 (5.37)

Where

$$ecm1 = u_t^* + 1.18c_t + 0.46y_t - 0.24i_t$$

$$ecm2 = c_t + 0.80y_t + 0.05i_t - 0.00t$$

t-values are in brackets.

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

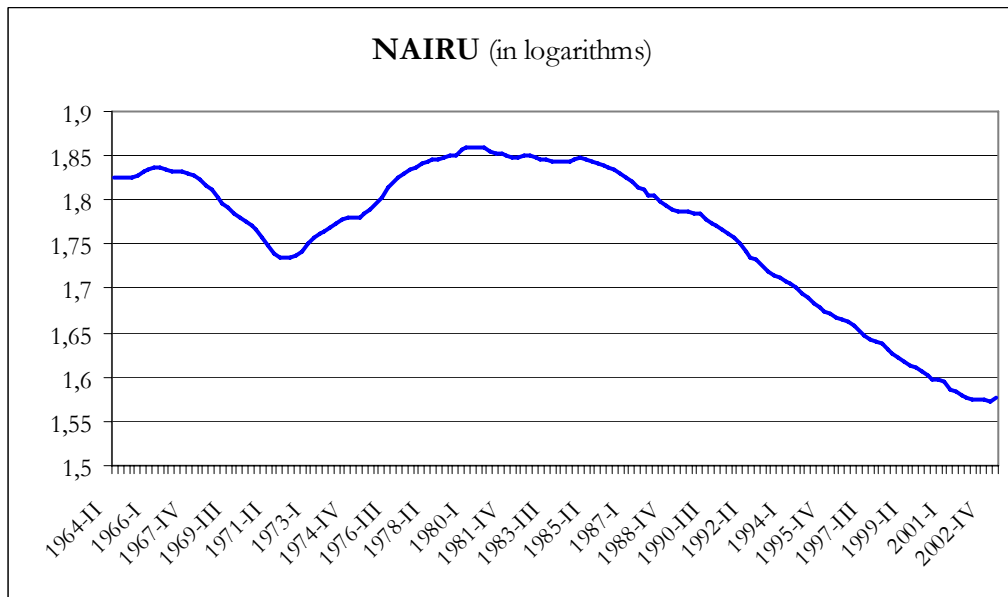
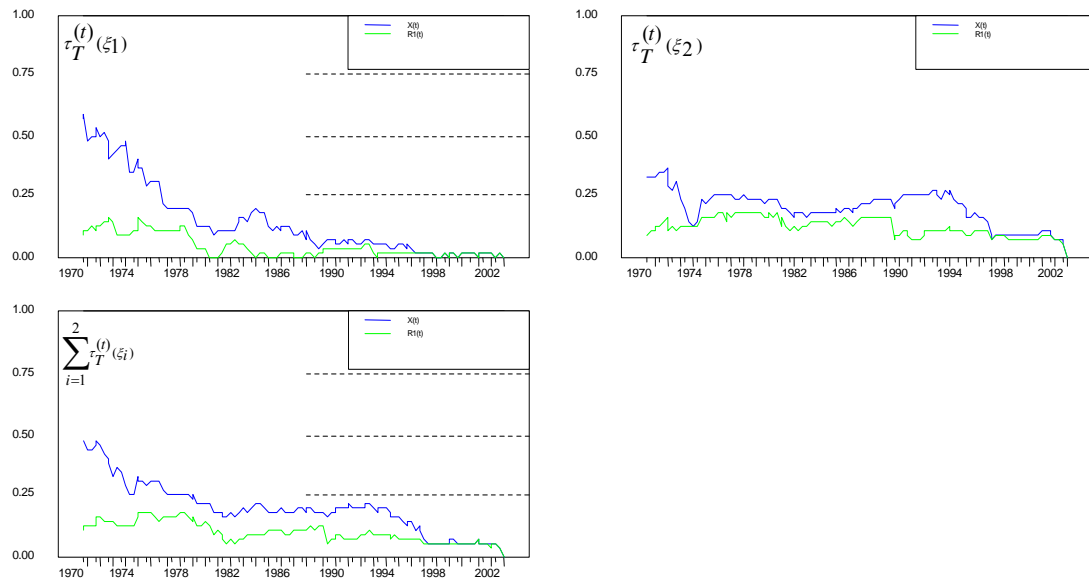
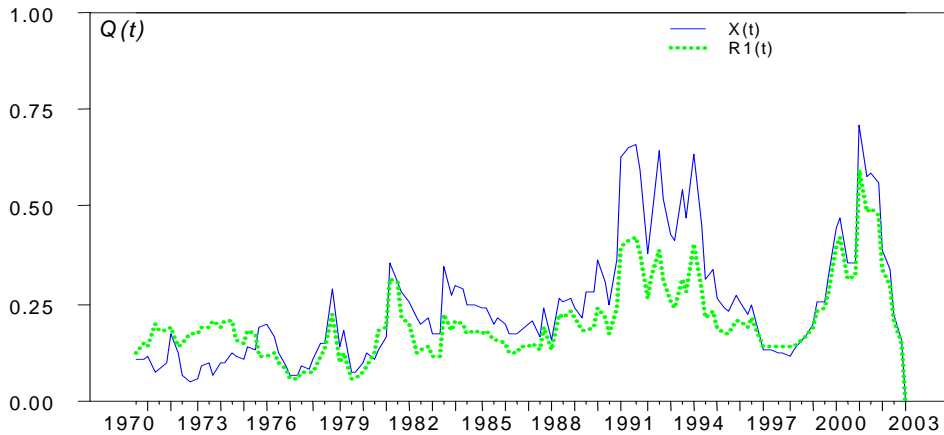


Figure 2. Eigenvalue Fluctuation Test



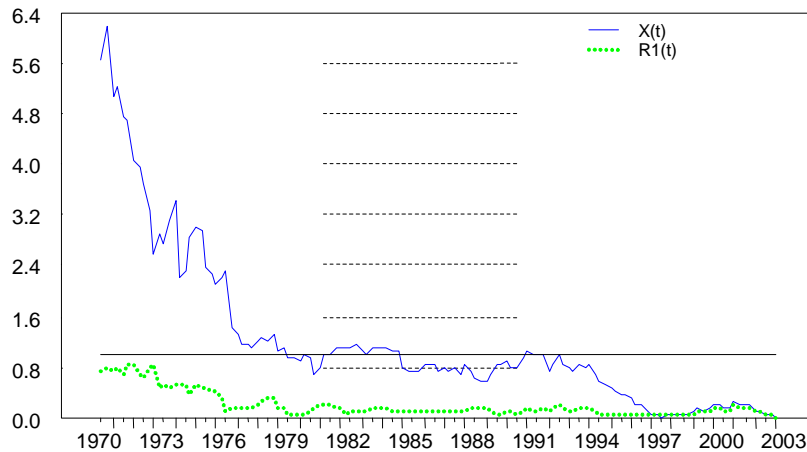
Note: the critical value for $\sup \tau_T^{(1)}$ at the 5% significance level is 1.36

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.

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 labour-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.

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 labour and reduces unemployment' (Layard *et al.*, 1991, p. 107). Interestingly, Rowthorn (1999) shows that the bulk of the empirical

estimates of the elasticity of substitution between capital and labour 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 *et al.* (1991), Rowthorn (1999) shows that, with an elasticity of substitution lower than unity, the equilibrium rate of unemployment is affected by technical progress, labour 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.

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

⁵ Blanchard and Katz (1997, p. 57) 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 *et al.* (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, p. 132). 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.

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 labour 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.

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, p. 56), in the U.S. 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.

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 labour 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.

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

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 $r = 2$.

¹⁸ The visual analysis of the figures associated to 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.

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