CAPITAL STOCK AND UNEMPLOYMENT IN CANADA

By

Alfonso Palacio-Vera, Ana Rosa Martínez-Cañete, Inés Pérez-Soba Aguilar and Elena Márquez de la Cruz

Universidad Complutense de Madrid (Spain)

Abstract: The purpose of this paper is to test the proposition that capital stock relative to aggregate output was an important variable in the determination of the Canadian NAIRU in the period 1976q1-2003q4. We present new empirical evidence obtained from the use of Canadian time-series data which lends support to the claim that the aggregate capital-output ratio, the real price of imports and the degree of generosity of the unemployment subsidy system were determinants of the NAIRU in the period considered. 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 was one determinant of the NAIRU in the period considered.

Key words: Capital-output ratio, cointegration, NAIRU, aspiration wage
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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

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I INTRODUCTION

The purpose of this paper is to test the proposition that capital formation is an important variable in the determination of unemployment in the Canadian economy. Most of the literature on Canadian unemployment has focused either on labour market issues such as the role of union density, minimum wages, unemployment insurance (UI) benefits, taxation, the demographic composition of the labour force or the role of aggregate demand and sectoral shocks. Nonetheless, there is one aspect to the Canadian unemployment problem which has been neglected. This is the relationship between unemployment and the capital stock. A number of economists have emphasized the need to analyse the role of the capital stock when it comes to explaining unemployment in OECD countries. Some examples are the studies by Malinvaud (1986), Sneessens and Dréze (1986), Modigliani et al. (1987), Burda (1988), Bean (1989), Sarantis (1993), Rowthorn (1999), and Sawyer (2002).[1] The link between the (physical) capital stock and unemployment has been the focus of attention of several empirical studies which find that the former (or its rate of growth) is a significant determinant of unemployment in the long run (Gordon, 1997; Arestis and Biefang-Frisancho Mariscal, 1998, 2000; Miaouli, 2001; Malley and Moutos, 2001; Stockhammer, 2004; Arestis et al., 2007; Karanassou et al., 2008a, 2008b; Palacio-Vera et al., 2008). But these studies represent a minority view and the generally accepted view in mainstream economics is that persistent unemployment is mainly due to labour market rigidities.[2] This position is well exemplified in the very 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 and changes in the aggregate capital-output ratio. This position has become the conventional wisdom in macroeconomic analysis (Blanchard and Katz, 1997).

Whilst the literature on the relation between the capital stock and unemployment has tended to focus either on the potential role of the capital stock or its rate of growth in reducing unemployment, one source of changes in the latter which has barely been discussed is changes in the aggregate capital-output ratio. Changes in the former 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 capital-output ratio for reasons other than (i) and (ii). A fall (rise) in the latter will transitorily reduce (increase) the rate of growth of labour productivity thus causing a mismatch between workers′ wage aspirations and productivity growth. In Layard et al. (1991) there are two reasons why changes in the aggregate capital-output ratio cannot 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 the capital stock and technical progress.[3] But, for this to be the case one needs to assume that (real) wage aspirations rapidly adapt to changes in productivity growth. Contrastingly, if workers demand real-wage increases based on their previous experience and (i) adjust slowly their target real wage to the new trend rate of productivity growth or (ii) simply do not adjust it, then there is the possibility that shifts in the capital-output ratio and, hence, in productivity growth affect the NAIRU (non-accelerating inflation rate of unemployment) in the long run.[4] To the best of our knowledge this hypothesis has only been tested for the U.S. economy in Palacio-Vera et al. (2008) where the authors find that the aggregate capital-output ratio is a significant determinant of the U.S. NAIRU. Our objective in this study is to test this hypothesis for the Canadian economy for the period 1976q1-2003q4.

Hence, we set ourselves the task of exploring the possibility that changes in labour productivity brought about by changes in the aggregate capital-output ratio are *not* fully translated into changes in real wages for a time span long enough as to lead to changes in equilibrium unemployment. For that purpose, we assemble a simple model characterising a closed economy where imperfect competition prevails in the product market and money wage increases are determined through collective bargaining. We come up with an expression for the equilibrium rate of unemployment or NAIRU that is subsequently estimated using co-integration techniques in the context of the Phillip-Hansen Fully-Modified Ordinary Least Squares (FMOLS) cointegration technique. The model is then estimated using quarterly data obtained mostly from Statistics Canada for the period considered.[5] The empirical results obtained challenge conventional wisdom in the field since, as in a previous study for the U.S. economy (Palacio-Vera et al., 2008), they lend support for the hypothesis that the aggregate capital-output ratio was one significant determinant of the NAIRU in Canada in the period 1976q1-2003q4. The results suggest that changes in the rate of growth of labour productivity brought about by changes in the aggregate capital-output ratio were *not* fully translated into changes in actual real wages thus affecting the NAIRU in the period considered. Results also suggest that the real price of imports and the degree of generosity of the unemployment subsidy system were significant determinants of the NAIRU whereas factors like technical change and mark-ups set by firms in the manufacturing sector did not affect the NAIRU significantly.

Arguably, since we find that the Canadian NAIRU depends negatively on the aggregate capital-output ratio and, hence, positively on the level of real interest rates, these results clearly identify a potential link between monetary policy and equilibrium unemployment which adds to a long list of studies which either postulate or find that real interest rates and unemployment are positively related (Phelps, 1994; IMF, 1999; Blanchard and Wolfers, 2000; Fortin, 2001; Baccaro and Rei, 2007; Schettkat and Sun, 2008).[6] In retrospect, our results also identify a channel through which the alleged restrictive monetary policy implemented by the Bank of Canada in the early nineties could have resulted in an increase in the Canadian NAIRU and thus complements the more conventional view based on the convexity of the long-run Phillips curve (Fortin, 1996, 1999).[7]

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 for 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 noted above, a number of authors have discussed the role of the capital stock 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.[8] 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-output 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 countries as well as Canada in the 1980s and 1990s, the unemployment rate may exhibit a trend for periods spanning well beyond a decade. One such source of persistent changes in the unemployment rate is changes in the aggregate capital-output ratio. This is because, as noted in Blanchard and Katz (1997), the equilibrium rate of unemployment depends on the level of productivity in relation to the reservation wage, as well as on many other factors. As a result of it, a higher capital stock relative to output raises labour productivity thereby mitigating inflationary pressures and allowing the economy to operate with a lower rate of unemployment. But real wage aspirations will eventually adjust upwards. 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, the rate of unemployment will exhibit a (downward) trend for some time.[9]

A similar argument is put forward in Stiglitz (1997, p. 7) and Ball and Mankiw (2002). For instance, the latter argue that in a neoclassical world, a rise in productivity growth has no obvious effect on inflation because higher productivity is fully reflected in higher real (product) wages. Nonetheless, 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.[10] According to them, all we need to accept is that "wage aspirations" adjust slowly enough to shifts in productivity growth.[11]

The predominant view is that workers´ demands for increased real wages depend mainly on their past rate of change. For instance, Ball and Moffitt (2001) estimate a

Phillips curve for the U.S. economy and allow for the possibility that wage setters have a target for real-wage growth that depends on an average of past productivity growth. Their Phillips curve model nests two competing hypotheses: (i) a neoclassical scenario where changes in trend productivity growth do not affect the Phillips curve and (ii) a non-neoclassical scenario where changes in trend productivity growth do affect the Phillips curve at least transitorily. They estimate the model using U.S. time series data for the period 1962-2000 and find strong support for the non-neoclassical version of the model.

Their results are qualified in Skott (2005) and Setterfield and Lovejoy (2006). Skott (2005) explores analytically the impact on the NAIRU of different formal specifications of the process of adjustment of wage aspirations. He defines the case discussed above and originally associated to Stiglitz (1997) as the "non-hysteretic" version of the theory of wage aspirations in which changes in trend productivity growth affect the NAIRU *only* transitorily. Next, he identifies a "hysteretic" version in which changes in productivity growth affect the NAIRU *permanently*. In this second scenario, the level at which the NAIRU settles in the long run depends crucially on the behaviour of aggregate demand. In Setterfield and Lovejoy (2006) it is assumed that both workers´ aspirations and bargaining power are jointly determined by the degree of worker insecurity. The latter is measured by a number of institutional variables. They estimate a Phillips curve model for the U.S. economy that is similar to that in Ball and Moffitt (2001) but which incorporates the institutional structure alluded to above. They then test which specification yields better empirical results and find that their own specification fares slightly better. They interpret their results as suggesting that the determinants of workers´ wage aspirations and bargaining power, i.e. the degree of worker insecurity, affect the NAIRU. Crucially, when the model specification includes all the institutional variables, the past rate of growth of real wages becomes insignificant so they conclude that changes in productivity growth may affect equilibrium unemployment *permanently*.

The theoretical model displayed in Section III takes inspiration from the work of Allen and Nixon (1997) who develop a model of the NAIRU similar to the model in Layard *et al*. (1991) in which they 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 noting that setting such restriction 'obscures the fact that there are actually two different processes taking place: capital accumulation and technological progress' (Allen and Nixon, 1997, p. 138). They show that the NAIRU is a function of relative factor prices and of the rate of accumulation so there is no unique short-run NAIRU. Notwithstanding, they 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-output ratio barring the study in Palacio-Vera *et al*. (2008) where, using U.S. data, the authors find that the NAIRU is a negative function of the aggregate capital-output ratio.

As for Canada, the marked rise in the Canadian unemployment rate vis-á-vis the United States in the 1980s and, especially, in the 1990s inspired a number of studies that attempted to quantify the contribution to the Canada-US unemployment gap of supply-side factors like (i) sectoral and technological shocks (Fortin and Araar, 1997; Lu, 1997; Sargent, 2000), (ii) adverse labour demand shifts (Kuhn, 2000), (iii) changes in the unemployment insurance (UI) system (Hornstein and Yuan, 1999), and (iv) changes in participation rates (Ip et al., 1999). Additionally, some studies have discussed and tested several competing hypotheses (Riddell and Sharpe, 1998; Riddell, 1999; Fortin et al., 2001). Importantly, some studies have emphasized the role of aggregate demand shocks and, particularly, of high real interest rates, in raising Canadian unemployment in the

1980s (McCallum, 1987; Fortin, 1989) and the 1990s (Fortin, 1996, 2001), albeit some authors have criticised this hypothesis (Freedman and Macklem, 1998). Finally, some studies have focused on the behaviour of the Canadian Beveridge curve to explain the behaviour of unemployment (Fortin, 1999; Osberg and Lin, 2000; Archambault and Fortin, 2001). The picture that emerges from our reading of this literature is that the factors leading to a rise in Canadian unemployment in the 1980s differ markedly from those in the 1990s. In particular, supply-side factors like changes in the labour force attachment of the non-employed associated to changes in the UI system is the most popular explanation of the emergence of the positive Canada-US unemployment gap in the 1980s whereas adverse macroeconomic shocks (particularly high real interest rates) appear to be the key factor in the early 1990s (Riddell and Sharpe, 1998; Stanford, 2005). For instance, in an influential work, Fortin (1996) dismisses some popular explanations of the rise in Canadian unemployment in the early 1990s and puts it down to a combination of restrictive monetary and fiscal policies implemented in Canada since the late 1980s. In particular, he insists that the rise in equilibrium unemployment in the early 1990s was the ultimate result of the attempt by the Bank of Canada to press the inflation rate below three percent in a context of downward nominal wage rigidity.[12]

III A THEORETICAL MODEL OF UNEMPLOYMENT DETERMINATION AND THE ROLE OF THE 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 workings of the model is the impact of the capital stock on unemployment. We postulate a simple framework in which the capital stock affects unemployment indirectly through its effect on the marginal product of labour. The short-run profit maximizing decision facing the typical firm i consists of maximizing profits

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

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 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.[13] 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:

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

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:

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

Hence the size of the (average) mark-up depends inversely on the price-elasticity of demand ε_t^d and, as we note below, the latter may 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.[14] Differentiating (4) with respect to employment, we obtain the marginal product of labour 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 labour 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 wage and *r* is the rental price of capital services. Combining expressions (4) through (7) we can express the marginal product of labour as:

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

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 drives 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. We postulate the following actual money wage equation:

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

where w_t is the actual (bargained) money wage, p_t^e is the expected price level, φ_t denotes 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. We define the real price of imports as:

$$I = \frac{E \cdot p_m}{p} \tag{10}$$

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 the probability of finding another job which, in turn, depends on (i) the rate of unemployment (*U*), and (ii) a composite index of "unemployment insurance generosity" (*UIG*).[15] We define φ_t as:

$$\varphi_t = \varphi(U_t, UIG_t) \tag{11}$$

where we assume that φ_t is log-linear, $\varphi_U \succ 0$, and $\varphi_{UIG} \prec 0$.[16]

Expressions (9) and (11) require some explanation. They make workers´ target real wage depend on the unemployment rate, the degree of generosity of the UI system and on real wage resistance factors such as the real price of imports. An increase in the unemployment rate increases 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.[17] Conversely, in tight labour markets the target real wage will be higher than the reservation wage. Hence, even if the reservation wage were constant overtime, the target real wage would vary with labour market conditions. The presence of the variable *UIG* in expression (11) is due to the fact that several empirical studies for Canada find that changes in the Canadian UI system have had a significant and permanent effect on the Canadian unemployment rate.

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” (Blanchard and Katz, 1997). The assumption that wages depend on what workers consider “fair” represents a departure from neoclassical economics, but it receives strong empirical support. For instance, Akerlof and Yellen (1990) argue that workers reduce their effort if they perceive wages as “unfair”, making it in firms´ interests to pay “fair” wages. Perhaps a better word than reservation wage in this context is “aspiration” wage (Oswald, 1986). By “aspiration” wage we mean the real wage that workers consider “fair” (Ball and Moffitt, 2001). We assume it is determined by all the factors that the literature has identified as affecting the reservation wage plus the aspiration to a steady improvement in living standards. As such, the rate of increase of the “aspiration” real wage is likely to be determined by a wide range of idiosyncratic institutional and historical factors including past wage increases. As noted above, the notion that the “aspiration” wage of workers is tied, at least partly, to past wage increases receives strong support in the studies by Ball and Moffitt (2001) and Setterfield and Lovejoy (2006) for the U.S. economy. Thus, the (positive) time trend on equation (9) can be envisaged as capturing workers´ aspiration to ever-rising living standards.

Insofar as workers bargain over a target (consumption) real wage we may also need to take into account the wedge between real product wages and real consumption wages. This wedge includes taxes on labour net of income transfers from the government to workers and the discrepancy between the *GDP* deflator and the consumer price index (CPI). Unfortunately, we could not obtain quarterly data to construct a precise measure of the tax wedge.[18] We measure the discrepancy between the GDP deflator and the CPI by the real price of imports (*I*). Thus, to the extent that workers exhibit some degree of real wage resistance, increases in *I* will induce upward pressure on the target real product wage and vice-versa. Next, if we insert (11) into (9), assume that function ω is log-linear and take logarithms, we get:

$$\ln w_t - \ln p_t^e = \overline{\omega} + \omega_u u_t + \omega_{uig} uig_t + \omega_i i_t + \lambda_2 t \tag{12}$$

where the logarithms of the arguments in ω are denoted by lower-case letters, $\overline{\omega}$ is the 'aspiration' (real) wage at the initial period, $\omega_u \prec 0$, $\omega_{uig} \succ 0$, and $\omega_i \succ 0$.

Some studies find that firms′ mark-ups exhibit cyclical patterns. However, there is no consensus in the literature as to whether they are pro-cyclical or counter-cyclical (Bils, 1987; Rotemberg and Woodford, 1991; Chirinko and Fazzari, 1994; Galeotti and Schiantarelli, 1998).[19] We hypothesize that mark-ups are set according to:

$$m_t = \overline{m}_t \cdot \left(\frac{C_t}{C_t^d} \right)^{\phi} \tag{13}$$

where C_t is aggregate capacity utilization, $\overline{m}$ is the average mark-up set by firms when $C_t = C_t^d$, $0 \prec C_t^d \prec 1$ is firms′ "desired" rate of capacity utilization, and the sign of ϕ is uncertain. Expression (13) implies that the size of the average mark-up may vary according to the phase of the business cycle the economy is where the latter is captured by aggregate actual capacity utilisation relative to "desired" capacity utilization. If we take logarithms in (13) we get:

$$\ln m_t = \ln \overline{m}_t + \phi c_t - \phi c_t^d \tag{14}$$

where $c_t = \ln C_t$, $c_t^d = \ln C_t^d$, 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:

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

We define long-run equilibrium as a situation where exogenous factors are kept fixed, short-run adjustments have already worked themselves out, and expectations are fulfilled. If so, we then have that $C_t = C_t^d$, $w_t = w_t^e$, and $p_t = p_t^e$. Hence, (12) and (17) become respectively:

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

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

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

$$u_t^* = d_0 + d_1 t + d_{uig}\, uig_t + d_i i_t + d_k k_t + d_{mu}\, mu_t \tag{20}$$

where $d_0 = \dfrac{\Omega_0 - \overline{\omega}}{\omega_u}$, $d_1 = \dfrac{\Omega_1 - \lambda_2}{\omega_u}$, $d_{uig} = \dfrac{-\omega_{uig}}{\omega_u} \succ 0$, $d_i = \dfrac{-\omega_i}{\omega_u} \succ 0$, $d_k = \dfrac{\Omega_2}{\omega_u} \prec 0$,

$d_{mu} = \dfrac{-1}{\omega_u} \succ 0$, $mu_t = \ln\overline{m}_t$, and the sign of d_0 and d_1 is *a priori* ambiguous.

Expression (20) is subsequently used as a benchmark for estimation purposes in the empirical work presented below. It tells us that the NAIRU depends positively on the real price of imports and the degree of generosity of the UI system and negatively on the aggregate capital-output ratio. It also tells us that it may shift over time in a way determined by the sign of d_1 if (real) wage aspirations do not grow in line with total factor productivity. Finally, it tells us that the NAIRU also depends on all the factors embedded in d_0 such as the parameters of the production function and the wage equation. To the extent that the NAIRU is a negative function of the aggregate 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 utilised. This is followed by the presentation of the results derived from the estimation of the cointegrating relationship.

Variable definition and data

The model is estimated using quarterly and seasonally adjusted data for the period 1976q1 to 2003q4 for Canada. The choice of the sample period was determined by data availability. In particular, employment data needed for the construction of the variable *UIG* is only available from 1976. Most data were obtained from Statistics Canada. Several additional variables were dropped from the econometric analysis because it was not possible to obtain continuous time series for a long enough period. This was the case with variables like the federal relative minimum wage, taxes, union density, union coverage and union membership. The NAIRU data was obtained from the Ecowin database which, in turn, takes it from the OECD. Details on the construction of the NAIRU time series can be found in Richardson et al. (2000).[20] Data on the mark-up for the Canadian business sector was obtained from Leung (2008).[21] Finally, information for the construction of the time series on *UIG* was obtained from Statistics

Canada, Fortin et al. (2000), Lin (1998), Bertrand and Bédard (2002) and ServiceCanada (2004).

Next, we define and discuss briefly each variable used in the econometric analysis. In order to obtain the *capital-output ratio* we divide the real capital stock (K) by aggregate real output (Y). The former is defined as the stock of private non-residential fixed assets at current prices deflated by a price index for gross private non-residential domestic investment. The latter is defined as Real Gross Domestic Product. The *real price of imports* (I) is calculated as the ratio of a price index of import goods measured in Canadian dollars divided by the CPI. Mark-up data was obtained from Leung (2008) who provides estimates for the Canadian business sector as a whole. The measure of unemployment insurance generosity UIG is equal to the product of three variables:

$$UIG = COV * RR * (Max / Min)$$

where COV denotes the coverage ratio or proportion of the labour force insured, RR denotes the replacement rate for workers, Max is the maximum duration of benefits for a person who is minimally qualified, and Min is the minimum qualifying period giving access to UI benefits (Fortin, 1989). COV is equal to one after the passing in 1971 of the UI Act which provided nearly universal coverage (Lin, 1998). The replacement rate for workers is defined as:

$$RR = \frac{\text{REGULAR UI BENEFIT PAYMENTS / REGULAR UI BENEFICIARIES}}{\text{WAGES, SALARIES AND OTHER LABOUR INCOME / EMPLOYMENT}}$$

The time series for "total regular UI benefit payments", "total regular UI income beneficiaries", and "employment" exhibit monthly periodicity and thus were converted into equivalent quarterly data by averaging monthly data for periods of three months.[22]

As mentioned above, values for *Max* and *Min* were calculated by combining information from different sources. Table 5 below shows the average value for Canada of the ratio Max/Min for different sub-periods.

The OECD uses three distinct concepts of the NAIRU each of them relating to the same basic idea of an unemployment rate consistent with stable inflation but differ according to the time horizon to which they refer. The NAIRU (with no qualifying adjective) is defined as the rate towards which the unemployment rate converges in the absence of temporary supply influences once the dynamic adjustment of inflation is completed (i.e., in the medium term or when the effects of temporary supply shocks dissipate). The *short-term* NAIRU is defined as that rate of unemployment consistent with stabilising inflation at its current level in the next period. It depends on the NAIRU (as defined above) but is more volatile because it is affected by all supply influences, including temporary ones, expectations, inertia in the process of dynamic adjustment and speed-limit effects. Importantly, it will also be influenced by the level of actual unemployment. Finally, the *long-term* equilibrium unemployment rate (akin to the natural rate of unemployment) corresponds to a long-term steady-state in which the NAIRU has fully adjusted to all supply and policy influences, including those having long-lasting effects. As a result, we have that the NAIRU and the short-term NAIRU are, in principle, candidates to measuring the NAIRU derived in Section III. However, we believe the short-term NAIRU concept is less appropriate than the NAIRU. This is because the former is potentially affected by (i) a large number of temporary supply shocks not captured in our theoretical model and (ii) the actual rate of unemployment.

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 expression (20). For that purpose we use the residual-based test of cointegration proposed in Shin (1994). The analysis will proceed in two stages. In the first stage, we estimate the long-term relationship by applying Phillips and Hansen (1990) Fully Modified Ordinary Least Squares (FMOLS) methodology and obtain the set of residuals. In the second stage, we apply the test proposed by Shin (1994) on the residuals in order to test whether there is a cointegrating relationship among the variables embedded in the theoretical model.

Unit root tests

In tables 1 and 2 below we show the results of the *ADF* unit root tests, the tests proposed in Elliot et al. (1996) (hereafter ERS), and the M tests proposed by Ng and Perron (2001).[23] In all these tests, the null hypothesis is the existence of a unit root. In the case of the capital-output ratio, the unit-root hypothesis can be rejected at the 5 percent significance level when we apply the *ADF* test and assume there is a constant in the data generating process (DGP). By contrast, both the *ERS* tests and the M tests of Ng and Perron (2001) do not allow us to reject the null hypothesis of a unit root. In the case of the real price of imports, both the *M* and *ERS* tests lead to the rejection of the unit root hypothesis when we assume there is a constant in the DGP, whereas the results of the *ADF* test suggest the variable is I(1). In order to elucidate the order of integration of these series we subjected them to the stationarity tests proposed in Kwiatkowski *et. al* (1992), not shown here, and the results suggest that both in the case of the capital-output ratio and the real price of imports we can reject the stationarity hypothesis. In sum, we thus conclude that both time series are I(1).[24]

Long-run relationship

Once we know the order of integration of the time series, we can estimate the long-run relationship between the NAIRU and the rest of the variables embedded in expression (20) by applying Phillips-Hansen´s FMOLS. This methodology provides a robust correction of the problem stemming from the existence of serial correlation in the error terms and potential endogeneity of the explanatory variables that typically arises when the long-run relationship is estimated by estimating it with OLS. This way, we achieve an efficient estimation of the cointegrating vector. By using the residuals obtained from the long-run relationship, Shin (1994) elaborates two *LM* statistics that consist of a multivariate extension of the stationarity test proposed by Kwiatkowski et al. (1992) for the univariate case.[25] The statistics are C_{μ} and C_{τ}. The former allows to test for the existence of a stochastic cointegration relationship (when there is a trend in the long-run relationship) whereas the latter allows to test for the existence of a deterministic cointegration relationship (when there is not a trend).

We then estimated equation (20) both with and without a trend by applying FMOLS. Since the trend was not found to be significant, we show in table 3 the results obtained from the estimation of the long-run relationship without a trend. In the table we also show the results of the test by Shin (1994) when applied to the deterministic cointegration.[26] It can be seen that the null hypothesis of the existence of a deterministic cointegration relationship cannot be rejected at the 5% significance level. The estimated cointegration relationship is:

$$u_t^* = \underset{(1.03)}{2.23} + \underset{(1.88)}{0.41}\, i_t - \underset{(2.26)}{0.18}\, k_t - \underset{(0.65)}{0.38}\, mu_t + \underset{(4.32)}{0.10}\, uig_t + \hat{\varepsilon}_t$$

where, as can be seen, all explanatory variables except the mark-up are significant and exhibit the expected sign. The coefficient of the mark-up was not found significantly different from zero. As to the error-correction mechanism, both the AIC and the SBC information criteria lead us to select one lag. The error-correction mechanism for the NAIRU is as follows (t-statistics in parentheses):[27]

$$\Delta u_t^* = \underset{(-0.17)}{-0.00} + \underset{(15.54)}{0.84}\, \Delta u_{t-1}^* - \underset{(-1.15)}{0.04}\, \Delta i_{t-1} + \underset{(0.13)}{0.04}\, \Delta mu_{t-1} - \underset{(-0.01)}{0.01}\, \Delta k_{t-1} + \underset{(0.15)}{0.00}\, \Delta uig_{t-1} - \underset{(-1.93)}{0.03}\, \hat{\varepsilon}_{t-1}$$

Again, it can be seen that the error-correction term presents the expected sign and is significant albeit only at the 10% percent level. We may also 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.

Finally, and given that the use of the FMOLS estimators is only legitimate when there exists one cointegration relationship among I(1) variables, we have also applied the trace test proposed by Johansen (1995) in order to test whether there is more than one cointegration relationship among the variables embedded in the theoretical model. Therefore, in table 4 we show the results obtained when we considered both a constant restricted and a constant not restricted to the cointegration vector. According to the information criteria, we used two lags in the VAR model. It can be seen that we cannot reject at the 5% level the hypothesis that there is only one cointegration relationship.

V SUMMARY AND CONCLUSIONS

The purpose of this paper was to examine the proposition that the capital stock relative to aggregate output has been an important variable in the determination of unemployment in the Canadian economy in the period 1976q1-2003q4. 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 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 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 but still relatively 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 Shin (1994) methodology to Canadian time-series data which lends support to the claim that the aggregate capital-output ratio, the real price of imports and the degree of generosity of the unemployment subsidy system (*UIG*) were significant determinants of the Canadian NAIRU in the period considered. In particular, increases in the aggregate capital-output ratio and decreases in the real price of imports and the *UIG* were found to be associated with significant decreases in the NAIRU. The same evidence showed that we cannot reject the hypothesis that technical progress 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 may only deliver low inflation through low growth and higher unemployment.

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FIGURES AND TABLES

Figure 1. Evolution of the NAIRU and the capital/output ratio in Canada over the period 1976q1-2007q7

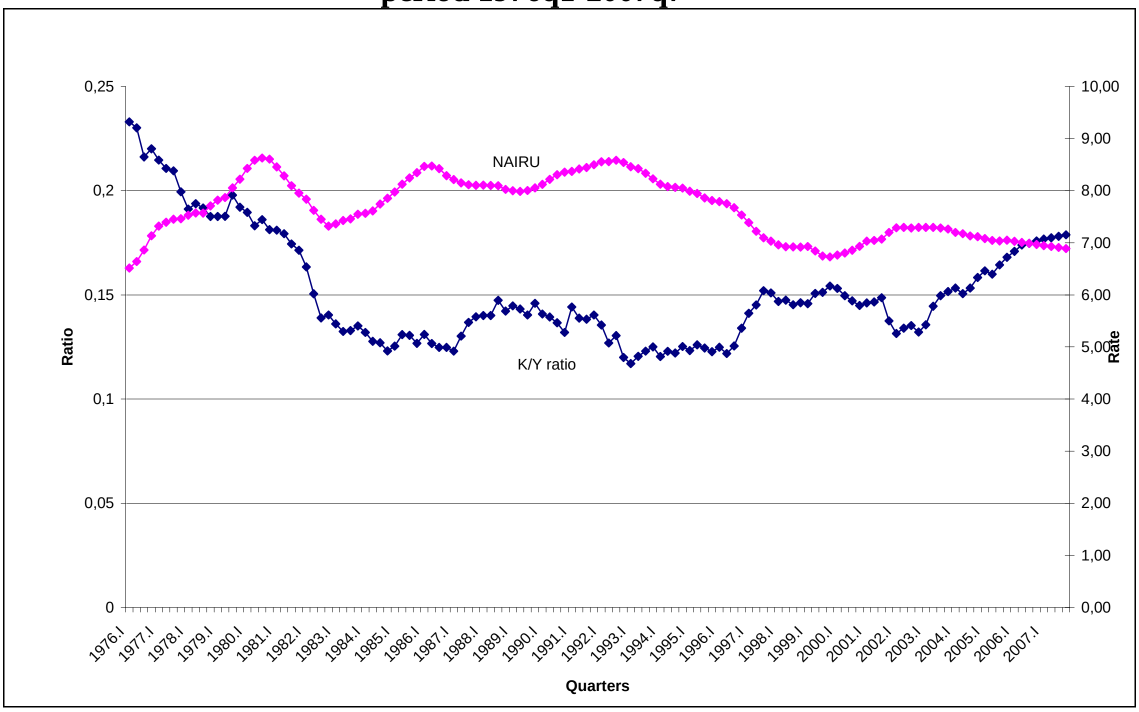

Table 1. ADF and Elliott-Rothenberg-Stock (ERS) unit root tests, 1976:1-2003:4

	ADF			ERS	
	[1]	[2]	[3]	P_T^{μ}	P_T^{τ}
u^*	-1.88	-2.61	-0.25	8.34	23.12
k	-2.95**	-1.76	-1.43	82.30	64.90
i	-1.37	-2.87	-0.50	2.59**	9.57
mu	-1.25	-1.59	-0.72	5.31	8.91
uig	-1.42	-1.86	-1.57	32.12	16.16

Notes:

[a] 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.

[b] (**) indicates that we can reject the null hypothesis of a unit root at the 5% significance level. The critical values were taken from Elliot *et al.* (1996), table 1.

[c] The number of lags was chosen according to the Modified Akaike Information Criterion (MAIC), with $k_{\max} = \text{int}\left(12(T/100)^{1/4}\right) = 12$.

Table 2. Ng-Perron unit root tests, 1976:1-2003:4

	MZ_{α}^{GLS}	MZ_t^{GLS}	MSB^{GLS}	MP_T^{GLS}
u^*	-1.95	-0.96	0.49	12.34
	-3.94	-1.35	0.34	22.48
k	-0.18	-0.18	1.00	54.08
	-1.02	-0.50	0.49	51.52
i	-11.45**	-2.26**	0.19**	2.64**
	-15.34	-2.68	0.17	6.44
mu	-1.92	-0.87	0.45	11.57
	-5.92	-1.70	0.28	15.34
uig	0.18	0.10	0.57	23.85
	-5.75	-1.69	0.29	15.83

Notes:

[a] In the first row the tests include a constant and in the second row they include a constant and a trend.

[b] (**) indicates that we can reject the null hypothesis of a unit root at the 5% significance level. The critical values have been taken from Ng and Perron (2001), table 1.

[c] The criteria used to choose the number of lags was the Modified Akaike Information Criterion (MAIC), with $k_{\max} = \text{int}\left(12(T/100)^{1/4}\right) = 12$.

Table 3. Estimation of the long-run relationship

$u_t^* = \alpha + \beta_1 i_t + \beta_2 k_t + \beta_3 mu_t + \beta_4 uig_t + \varepsilon_t$ **and Shin (1994) cointegration test**

Variables	Coefficient
Intercept	2.23
	(1.03)
i	0.41
	(1.88)
k	-0.18
	(2.26)
mu	-0.38
	(0.65)
uig	0.10
	(4.32)
Deterministic cointegration Test C_μ	0.115

Notes:

[a] The covariance matrix robust to heteroscedasticity and serial correlation has been constructed following Andrews (1991), using a Bartlett kernel with a VAR(1) model to calculate the automatic bandwidth. The residuals are pre-whitened using a VAR(1).

[b] t-ratios of the FMOLS are in brackets.

[c] For $m = 4$ the critical value for C_μ at the 5% level of significance is 0.121 (see Shin, 1994, table 1).

Table 4. Determination of the cointegrating rank in the system

$y' = [u_t^*, i_t, k_t, mu_t, uig_t]$

Unrestricted intercept				Restricted intercept			
r	$p-r$	*Trace test*	CV95	r	$p-r$	*Trace test*	CV95
0	5	73.303	69.611	0	5	80.406	76.813
1	4	45.714	47.707	1	4	52.260	53.945
2	3	26.403	29.804	2	3	31.139	35.070
3	2	10.683	15.408	3	2	14.755	20.164
4	1	0.736	3.841	4	1	2.289	9.142

Note:
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. Average values for Canada of the ratio Max/Min for different periods

Period (monthly data)	Period (quarterly data)	Weeks	Average for Canada	**Max/Min**
1972 – Nov 1977	**1976q1-1977q3**	Max = 26-44 Min = 8	35/8	**4,37**
Dec 1977– Jun 1979	**1977q4-1979q2**	Max = 26-44 Min = 10-14	35/12	**2,91**
Jul 1979 – Jan 1990	**1979q3-1989q4**	Max = 26-44 Min = 20	35/20	**1,75**
Feb 1990 – Nov 1990	**1990q1-1990q3**	Max = 26-44 Min = 14	35/14	**2,5**
Dec 1990 – Jul 1994	**1990q4-1994q2**	Max = 26-44 Min = 10-20	35/15	**2,3**
Ag 1994 – Dec 1996	**1994q3-2007q4**	Max = 26-44 Min = 12-20	35/16	**2,18**
Jan 1997 – Dic 2007	**1997q1-2003q4**	Max = 26-44 Min = 52	35/52	**0,67**

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[1] One key aspect emphasized in earlier studies of the impact of the capital stock on unemployment was the existence of limited ex-post substitutability between capital and labour that creates an *asymmetry* in the adjustment of unemployment to shocks of similar magnitude but opposite sign. According to these studies, an adverse supply shock that led initially to an increase in equilibrium unemployment would be followed by a decrease in the capital stock thereby reducing capacity output. However, if the supply shock was subsequently reversed, the unemployment rate could not reach rapidly the initial level owing to an insufficiency of physical capital (Bean, 1989).

[2] Brief expositions of the conventional view can be found in Nickell (1997), Siebert (1997) and IMF (2003). A systematic critique of the conventional view on the causes of European unemployment is in Howell et al. (2007) and a similar critique for the Canadian case is in Kuhn (2000).

[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). In this respect, 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) is, for most economies, inappropriate. Furthermore, using a simple model based on the work of Layard *et al.* (1991), he 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.

[4] In a recent empirical study for the OECD countries covering the period 1961-1995, Malley and Moutos (2001) find a strong negative association between capital accumulation and unemployment for all OECD countries but Belgium, Ireland and the United States. The authors argue persuasively that what matters for the evolution of unemployment in an international context is not the absolute growth rate of a country´s capital stock but its evolution *relative* to other countries´ capital stock.

[5] The Ecowin Database consists of economic and financial data collected from more than seven hundred primary sources as OECD, IMF or Statistics Canada, among others.

[6] Whilst not explicitly connected to the NAIRU, the results of the meta-analysis performed in De Grauwe and Costa Storti (2008) are highly relevant for our study. Using as many as 83 studies that report numbers on the effect of monetary policy on aggregate output, the authors conclude that the long-run neutrality of money has a weak empirical basis.

[7] In their influential study about Canadian unemployment, Fortin et al. (2001) find that the level of real interest rates were a significant determinant of the unemployment rate and conclude that 'future work will have to address… the nature and channels of the *permanent* effect of the real interest rate on unemployment' (Fortin et al., 2001, p. 87, emphasis added).

[8] The conventional wisdom is that the gradual adjustment of the capital stock to its optimal level tends to increase the degree of persistence of supply and demand shocks but nevertheless has no long-run effects on equilibrium unemployment (IMF, 1999; Blanchard, 2006).

[9] 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 appears plausible, it is still the case that it does not hold as a matter of logic and so there is no need to impose it *a priori*. Incidentally, in a recent study that uses U.S. data for the period 1948q4-2007q2, Schreiber (2009) finds that the unemployment rate is negatively related to productivity growth in the long run.

[10] Of course, this is not a new story. 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 decrease in the rate of productivity growth relative to workers´ target rate of growth of real wages.

[11] To support their hypothesis, they mention that Alan Greenspan '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 (beneficial) effect on the NAIRU as an exogenous decrease in workers´ aggressiveness.

[12] However, he rejects the notion of hysteresis as an explanation for the observed behaviour of Canadian unemployment (Fortin, 1989, 1996). This hypothesis has yielded mixed results when applied to Canadian data. For instance, Koustas and Veloce (1996) find that the rate of unemployment in Canada exhibits high shock persistence hence supporting the notion of hysteresis in unemployment (see also Jackson, 2000). By contrast, in a comprehensive study of Canadian labour markets, Jones (1995) concludes that there is no strong evidence in favour of hysteresis in the Canadian unemployment rate.

[13] We assume that imperfect competition results from product differentiation.

[14] 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 may lead to a higher unemployment rate. Be that as it may, Sargent (2000) finds there is in fact little evidence that technological change had an important negative impact on overall unemployment and employment rates in Canada over the 1990s.

[15] A variable that may affect the cost to firms of replacing workers is the "degree of union density" (*UD*). For instance, Fortin et al. (2001) estimate a positive impact of *UD* on unemployment albeit the variances around the estimates are large.

[16] The most important changes made to the UI Canadian system occurred in 1971 and 1972 with the approval of the 1971 UI Act and the Mackasey reform of 1972 respectively. According to Fortin et al. (2001), the implementation of the 1972 Mackasey reform —which raised the average UI wage subsidy to 242% from just 43% in previous years— contributed significantly to a rise in the unemployment rate. After 1972 there have been several partial reforms of the UI system whose effects on unemployment are more uncertain (see Fortin et al., 2001).

[17] The inverse relationship between the bargained 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 real wage relation in which the tighter is the labour market the higher is the real wage that firms have to pay to prevent shirking.

[18] In any case, the construction of a precise measure of the "tax wedge" is problematic. As argued in Blanchard (2006, p. 34) 'for each tax, we need to know whether it is levied on labour income and other benefits, or just on labour income. For each tax, we also need to know' the extent to which they entitle the taxpayer to benefits. For the case of Canada, Fortin et al. (2001) find that changes in the wedge (that includes a measure of the "regional" tax wage and the terms of trade) have a strong *but* transitory effect on unemployment.

[19] Leung (2008) provides estimates of Canadian mark-ups over the 1961-2003 period for the business sector as a whole and for 33 industries. He finds that, after rising steadily a total of 8.1 per cent over the 1961 to 1985 period, mark-ups in the Canadian business sector fell by 10 per cent over the next eight years. He also finds substantial heterogeneity in the evolution of mark-ups across industries.

[20] The OECD uses a reduced-form Phillips curve framework to measure the NAIRU (Richardson et al., 2000). The underlying structural model assumes an economy where wages are bargained between workers and firms and where firms operate in imperfectly competitive markets. This method constitutes a compromise between purely "structural methods" and the methods that attempt to estimate the NAIRU using a variety of statistical techniques to directly split the actual unemployment rate into cyclical and trend components, with the latter identified as the NAIRU.

[21] The time series does not appear explicitly in Leung (2008) so it was kindly provided by the author upon request.

[22] In particular, we obtained two time series made up of three-month averages for the variables "total regular UI income beneficiaries" and "employment" whereas, in the case of the "total regular UI benefit payments", we added the values corresponding to three consecutive months.

[23] 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).

[24] 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 are I(2).

[25] In his well-known study, Shin (1994) uses the residuals obtained from the estimation of the long-run relationship by applying dynamic OLS estimators (DOLS). Nonetheless, he argues that the LM statistics typically used to test the null hypothesis would have the same asymptotic distribution if, instead of using DOLS, we used an alternative efficient estimator as the FMOLS.

[26] The results obtained from the estimation of the model that includes a trend are available from the authors upon request.

[27] For the sake of simplicity we have only shown the error-correction mechanism for the NAIRU variable. The error-correction mechanisms for variables i_t, mu_t, k_t and uig_t are also available from the authors upon request.