Working Paper 96-7 / Document de travail 96-7

An Econometric Examination of the Trend Unemployment Rate in Canada

by
Denise Côté and Doug Hostland

Bank of Canada   Banque du Canada

May 1996

AN ECONOMETRIC EXAMINATION OF THE TREND UNEMPLOYMENT RATE IN CANADA

by

Denise Côté and Doug Hostland

Research Department

E-mail: dcote@bank-banque-canada.ca
Hostland.Douglas@fin.gc.ca

ACKNOWLEDGMENTS

The authors would like to thank Pierre Duguay, Irene Ip, Paul Jenkins, David Longworth, Tiff Macklem, Brian O'Reilly, Ron Parker, David Rose and Steve Poloz for many helpful comments and suggestions, and Sébastien Sherman for his assistance in preparing the graphs. We would also like to thank the participants of a joint Research Department / UQAM Macro-Labour Workshop for their comments and Helen Meubus for her editorial suggestions.

ISSN 1192-5434
ISBN 0-662-24596-2
Printed in Canada on recycled paper

ABSTRACT

This paper attempts to identify the trend unemployment rate, an empirical concept, using cointegration theory. The authors examine whether there is a cointegrating relationship between the observed unemployment rate and various structural factors, focussing neither on the non-accelerating-inflation rate of unemployment (NAIRU) nor on the natural rate of unemployment, but rather on the trend unemployment rate, which they define in terms of cointegration. They show that, given the non stationary nature of the data, cointegration represents a necessary condition for analysing the NAIRU and the natural rate but not a sufficient condition for defining them.

The main finding of the study is that two structural factors — the degree of unionization of the labour force and payroll taxes — can best account for the stochastic trend in the Canadian unemployment rate from 1955 to 1994. Accordingly, deviations of the observed unemployment rate from the trend unemployment rate during that period are treated as containing information relevant for measuring the output gap within the multivariate filter.

RÉSUMÉ

L'objectif des auteurs est de parvenir à identifier le taux de chômage tendanciel, concept de nature empirique, au moyen de la théorie de la cointégration. Les auteurs cherchent à établir s’il existe une relation de cointégration entre le taux de chômage et les différents déterminants structurels qui sont mis en avant dans la littérature. Leur recherche ne porte pas sur le taux de chômage non accélérationniste (TCNA) ni sur le taux de chômage naturel mais plutôt sur le chômage tendanciel défini dans le cadre de la théorie de la cointégration. Étant donné le caractère non stationnaire des données, l’hypothèse de cointégration représente une condition nécessaire à l’analyse du TCNA et du taux de chômage naturel mais non suffisante.

Le résultat principal de l’étude indique que les déterminants structurels qui expliquent le mieux la tendance stochastique du taux de chômage au Canada au cours de la période 1955-1994 sont le taux de syndicalisation de la population active et les cotisations sociales des employeurs en proportion du total des salaires et traitements. Ainsi, les écarts du taux de chômage observé par rapport au taux tendanciel renferment de l'information susceptible de servir à mesurer l'écart de production dans le cadre du filtre à plusieurs variables ou filtre multivarié.

CONTENTS

1 INTRODUCTION AND SUMMARY

In supporting the conduct of monetary policy, the staff of the Bank of Canada continually assess the significance of current economic, monetary and financial developments. A full range of key macroeconomic data are monitored and interpreted in the context of quarterly projections for the Canadian economy (Duguay and Poloz 1994). These projections are based on judgmental use of the model QPM (Quarterly Projection Model).[1] The model is used to generate scenarios for interest rates and exchange rates that, given various underlying assumptions (e.g., foreign output, commodity prices, etc.), are believed to be consistent with the overall economic outlook, including its starting point, and most importantly, with the achievement of the Bank's inflation targets. In this framework, for example, an unexpected increase in aggregate demand relative to potential output (or the opening of a positive output gap), will lead to higher expected inflation, and because the model embodies a policy reaction function, the model solutions for interest rates and exchange rates will adjust to the shock in order to return predicted inflation to the Bank's inflation targets gradually.

Estimates of potential output are, therefore, central to the structure of the QPM model. The Bank's staff have adopted a generalized method based on the Hodrick-Prescott univariate filter, called the multivariate filter, for the purpose of estimating potential output (see Laxton and Tetlow 1992 and Butler 1996). The multivariate filter methodology occupies a middle ground between the astructural Hodrick-Prescott method and a purely structural approach to estimating potential, treating the two as complementary. When a univariate filter such as the Hodrick-Prescott is used, the underlying trend in a time series is determined solely from the information contained in the series itself. The multivariate filter methodology, howevere, makes use of information about other variables often modelled in a structural manner, to give a better estimate of potential output.

1. See Black et al. (1994); Armstrong et al. (1995); and Coletti et al. (1996). See also Poloz, Rose and Tetlow (1994) for a summary of the model and its use.

By way of illustration, a rise in inflation is generally associated with the opening of a positive output gap. With the multivariate filter, the estimate of trend output can be adjusted to reflect this notion in such a way as to increase the likelihood that the sustainable level of output is, on average, lower than actual output when inflation is accelerating, and higher when inflation decelerates. By contrast, the univariate filter produces, by construction, a time series where the mean deviation from trend over the sample period is zero, regardless of the trend in inflation.

Other general elements derived from economic theory that are built into the current version of the multivariate filter are: *an unemployment equation that attempts to explain the trend unemployment rate on the basis of structural factors*; a long-term relationship between the level of real wages and labour productivity; and Okun's relationship linking the labour market gap to the output gap. Thus, in contrast with traditional methods, the multivariate filter approach allows a broader range of information to be taken into account in the estimation of potential output. Many more extensions are possible.

As mentioned above, the multivariate filter attaches some weight to information from the labour market in generating its estimates of economic potential. Specifically, it takes departures of the unemployment rate from the underlying filtered trend as providing some information about the deviation of output from trend. Ultimately, it would be desirable to evaluate a given movement in the unemployment rate with reference to the non-accelerating-inflation rate of unemployment (NAIRU). However, a full understanding of the NAIRU has to date proved to be elusive. This is particularly reflected in the existing empirical work on the NAIRU; Rose provides a survey of empirical studies on the NAIRU for Canada that were performed up until the mid-1980s. He concludes that "the results are quite sensitive to methodology, to measurement of variables and to the estimation sample period." (Rose 1988, 43). In a more recent analysis, Setterfield, Gordon and Osberg (1992) largely corroborate Rose's results and come to much the same conclusion. Taken together, the surveys by Rose and by Setterfield, Gordon and Osberg show that the various estimates of the NAIRU proposed by numerous empirical studies can best be described as fragile.

Instead of attempting to isolate a robust empirical NAIRU relationship, a less ambitious approach is currently used: the trend in unemployment is defined in statistical terms, using cointegration theory, and deviations from that trend unemployment rate are treated as containing information relevant to measuring the output gap.

The present paper deals with the empirical estimation of the trend unemployment rate for use within the multivariate filter. As such, it does not deal explicitly with either the natural rate of unemployment or the NAIRU. In this paper, we make a conceptual distinction between the natural rate of unemployment, the NAIRU and trend unemployment rate. It is shown that, given the nonstationary nature of the data, cointegration represents a necessary condition for analysing the natural rate and the NAIRU but not a sufficient condition for defining either one. The natural rate can be thought of as that rate of unemployment that the economy produces in steady-state equilibrium at which the flows in and out of the pool of unemployment equilibrate on average and at which economic agents are not fooled by accelerating or decelerating inflation. The NAIRU, in contrast, is the rate of unemployment at which there is no tendency for inflation to accelerate or decelerate. It is generally defined more narrowly in relation to the linkage between the level of excess demand/supply for labour and short- to medium-term pressures on wage and price inflation.

It is believed likely that the trend unemployment rate developed here is in some sense subsumed within a more general specification of the NAIRU, while being in all likelihood more variable than the underlying natural rate. Accordingly, we think of the trend unemployment rate conceptually as lying between the natural rate and the NAIRU. The research undertaken here is a step forward from more mechanical approaches to estimating trend unemployment (e.g., Côté and Hostland 1994)[2] and points the way for further work on the NAIRU.

Our contribution is to identify the trend unemployment rate, an empirical concept, which we define in terms of cointegration, for use

2. Côté and Hostland (1994) derive measures of potential output and the NAIRU as unobserved variables within a system of equations using the Hodrick-Prescott filter method.

within the multivariate filter. We examine which structural factors best account for the stochastic trend in unemployment, using a cointegration approach that explicitly takes into account econometric complications that arise from the nonstationary nature of the data in our sample. This basically involves determining whether the unemployment rate is cointegrated with various structural factors that have been proposed in the literature. Our analysis is limited to the structural factors proposed by Rose (1988) and Coe (1990). These factors include changes in the demographic composition of the labour force, the replacement rate associated with the Unemployment Insurance (UI) Program, the minimum wage rate, the proportion of the labour force that is unionized, and payroll taxes. We also examine Sargent's (1995) new measure of UI generosity.

The main result of this study indicates that there is evidence of a stable cointegrating relationship between the unemployment rate and two of the structural factors listed above: 1) the proportion of the labour force that is unionized and 2) payroll taxes. Due to the high degree of collinearity between these two factors, we are unable, based on the cointegration tests, to isolate a unique cointegrating vector. Consequently, we cannot determine whether the trend unemployment rate depends upon one or both of these two factors, nor can we precisely estimate the relative contribution of the two factors with a reasonable degree of confidence. Indeed, collinearity among the variables tested makes choosing between them quite difficult. However, we find little statistical support for a cointegrating relationship between the unemployment rate and the other structural factors considered — the minimum wage rate, the UI replacement rate, Sargent's index of UI generosity and the demographic composition of the labour force. In short, we conclude that unionization and payroll taxes are the most important structural determinants of the trend unemployment rate in Canada over the 1955-94 period, although other variables may have been important episodically and thus may have played an important role in determining the natural rate in certain periods.

It is worth noting that our inferences appear to be robust with respect to many things. For example, our main results are qualitatively robust with respect to different estimation procedures, to different estimation sample periods and to several specification issues, including stability. In addition, the finding of cointegration between unemployment and the rates of unionization and payroll taxation is robust to several different approaches to testing for common stochastic trends.

Nonetheless, our study is subject to at least three caveats. First, while our main results are qualitatively robust with respect to a variety of issues, they are, however, quantitatively sensitive to the inclusion of some cyclical variables as well as to alternative estimation procedures. Second, measurement error may play an important role in our analysis, particularly with regard to the variables used to proxy the various structural factors. In particular, we suspect that the lack of evidence supporting a long-run relationship between the unemployment rate and the UI variables tested here is due to the fact that they are poor proxies for the complexity of the UI program. Third, the problem of multicollinearity cannot be underplayed. Although the trend unemployment rate is best captured by the proportion of the labour force that is unionized and by payroll taxes, there is a theoretical case for thinking that the other variables explain trend unemployment as well. It might be easily the case that with the accumulation of more data, the trend would be best captured by another combination of cointegrated variables. In any case, if one (or more) of the structural factors that are not part of the cointegration vector move considerably (e.g., the UI variable), the multivariate filter methodology for estimating potential output allows the user to insert judgment to take them into account, especially over the projection period, so one is not tied mechanically to the simple specification developed in this paper.

The plan of the paper is as follows. The next section outlines the macro framework in which we define trend unemployment and discusses econometric issues relating to estimation and inference when the unemployment rate and its structural determinants are nonstationary. The third section of the paper examines whether the unemployment rate is cointegrated with the various structural factors. This is performed within a single-equation framework using an unemployment equation. In the fourth section an extensive sensitivity analysis is conducted in order to examine the robustness of our results. The final section of the paper summarizes our main results and comments on some areas for future research.

2 ESTIMATING TREND UNEMPLOYMENT

2.1 Conceptual Framework

A number of empirical studies have examined the equilibrium or the "natural" rate of unemployment.[3] We find it useful to make a conceptual distinction between the natural rate of unemployment and the NAIRU. The natural rate is conventionally used as an equilibrium concept, where equilibrium can be defined with reference to various shocks of interest as well as over different time horizons (Friedman 1968). In contrast, the NAIRU is generally defined more narrowly in relation to the linkage between the level of excess demand / supply for labour and inflation dynamics (as summarized by a Phillips curve). One could imagine shocks that require an adjustment in the labour market with no excess demand / supply for labour throughout the adjustment process. For example, consider large sectoral shifts that require a long period of time for workers to be retrained and / or migrate between regions. There could be no excess demand / supply for labour at the aggregate level throughout the adjustment process and hence, no inflationary pressure. In this case, the unemployment rate would be at the NAIRU but would be temporarily above its long-run equilibrium or natural rate.

Much of the existing research on measuring the natural rate of unemployment essentially attempts to isolate various structural factors that can account for long-run movements in the unemployment rate. This is typically performed in the context of an unemployment equation having the general form:

$$C^*(L)u_t = D^*(L)S_t + E^*(L)Z_t + \eta_t \qquad (2.1)$$

3. Recent empirical studies estimating the natural rate for Canada include Fougère (1995); Fortin, Keil and Symons (1995); Van Rijckeghem (1993); Milbourne, Purvis and Scoones (1991); Coe (1990); and Burns (1990).

where u_t is the observed unemployment rate, Z_t is a vector of variables that are intended to capture unemployment dynamics arising from factors other than the structural factors, S_t, and the random error term, η_t, while $C^*(L)$, $D^*(L)$ and $E^*(L)$ are polynomial lag operators. The dynamic relationship between the unemployment rate and the explanatory variables in equation (2.1) is modelled using an unrestricted autoregressive distributed-lag specification. The variables comprising Z_t are assumed to be of a cyclical nature and have no permanent effect on u_t. The natural rate of unemployment, u^*_t, could be derived from equation (2.1) by calculating the long-run effects of the structural factors S_t. This would entail first estimating equation (2.1) and then calculating the long-run parameters Φ, where $\Phi = D^*(1)C^*(1)^{-1}$ and $D^*(1)$ and $C^*(1)$ denote the sum of the coefficients in polynomial lag operators $D^*(L)$ and $C^*(L)$, respectively.

$$u^*_t = \Phi S_t \tag{2.2}$$

where the vectors S_t and Φ are partitioned into two subvectors as follows:

$$\Phi = [\Phi^1\ \Phi^2] \text{ and } S_t = [S^T{}_t\ S^O{}_t] \tag{2.3}$$

The subvector, S^T, embodies a class of structural factors that are considered for estimating the trend unemployment rate, while the subvector S^O contains other factors that are not part of the trend but which could nevertheless affect the natural rate. Similarly, Φ, is a vector of two subvectors of parameters corresponding to each subvector of S.

In the context of an unemployment equation, one must first identify the trend unemployment rate, $\Phi^1 S^T$, in order to estimate the natural rate of unemployment, defined as $\Phi^1 S^T + \Phi^2 S^O$. Thus, identifying the trend unemployment rate, $\Phi^1 S^T$, is a necessary condition for estimating the natural rate.[4]

An alternative way to estimate the parameters characterizing the natural rate is to express equation (2.1) in error-correction form as follows:

4. It can be shown, in a similar way, that identifying the trend unemployment rate is also a necessary condition for estimating the NAIRU in the context of a reduced-form Phillips curve.

$$C(L)\Delta u_t = D(L)\Delta S_t + E(L)Z_t - \gamma[u_{t-1} - \Phi^1 S^T{}_{t-1} - \Phi^2 S^O{}_{t-1}] + \nu_t \qquad (2.4)$$

In the context of the error-correction model (2.4), one can directly estimate the vector, Φ, using non-linear least squares. This results in estimates of Φ that are identical to the long-run elasticities obtained from estimating equation (2.1) and making the calculations: $\Phi = D^*(1)C^*(1)^{-1}$. The error-correction model (2.4) simply represents an alternative parameterization of the autoregressive distributed lag specification represented by (2.1). We focus on the error-correction form of the model in order to estimate the trend unemployment, $\Phi^1 S^T$, because it enables us to obtain estimates of Φ^1 in a straightforward manner.

In this paper, our goal is to measure trend unemployment as opposed to the natural rate or the NAIRU. The focus of the empirical work is therefore on estimating Φ^1, the parameters on the structural determinants of trend unemployment (S^T) in the unemployment rate equation. Since the remainder of the analysis focusses on Φ^1 and S^T and says nothing about Φ^2 and S^O, to simplify the notation we drop the superscripts. Thus, hereafter Φ denotes Φ^1 and S denotes S^T.

2.2 The Spurious Regression Problem and the Cointegration Approach

Casual observation of Figures 1 to 7 suggests that the unemployment rate and the various structural factors that have been proposed to explain the trend in unemployment are nonstationary over the time period shown. Although it is theoretically inappropriate for a bounded variable such as the unemployment rate to be truly nonstationary, it might nevertheless be so, in a statistical sense, over a given sample period. We examine this issue more formally in the following section of the paper.

Developments in cointegration theory over the last decade indicate that estimation and inference are greatly complicated by the presence of nonstationary variables. In particular, failure to account for nonstationarity can result in the spurious regression problem. Consider the situation where the unemployment rate u and the structural factors S are all nonstationary but are not cointegrated. Suppose, for example, that the econometrician has mistakenly excluded one structural factor from the cointegration vector. The theoretical results obtained by Phillips (1986) imply that in this situation, estimates of Φ derived from the unemployment equation (2.4) would have no well-defined statistical interpretation. Moreover, the estimated t-statistics corresponding to the parameters of the structural factors comprising S_t are biased upward.[5] Consequently, inferences made using conventional statistical procedures would tend to indicate that the structural factors have a statistically significant long-run effect on the unemployment rate, even when there is no underlying statistical relationship. This is known as the spurious regression problem.[6]

The spurious regression problem does not arise with stationary variables. For example, if the unemployment rate and the structural factors used to model the trend unemployment rate were all stationary, then normal distribution theory would apply and hence, conventional statistical procedures for estimation and inference would be valid. The spurious regression problem is an econometric complication that arises from the nonstationary nature of the data.

The spurious regression problem also does not arise in the case where the unemployment rate u and the structural factors S are cointegrated. Under the cointegration hypothesis, the parameters Φ have well-defined statistical properties and valid inferences can be made, provided that the appropriate statistical procedures are used.[7] Thus, in order to identify the trend unemployment rate and therefore the natural rate using an equation like that given by (2.4), the unemployment rate u must be cointegrated with the structural factors S.

5. More precisely, Phillips' (1986) analysis shows that the limiting distribution of Φ converges in probability to a random variable, while the limiting distributions of the t-statistics corresponding with the elements of Φ diverge with the sample size.

6. Granger and Newbold (1974) illustrated the nature of the spurious regression problem using a simulation approach. Phillips (1986) developed the asymptotic theory underlying their results.

7. This will be discussed more fully later in the text.

From an econometric perspective, the cointegration approach contributes to the existing research in the following way. Because of the apparent nonstationary nature of the data, conventional statistical procedures do not result in asymptotically efficient estimates of the parameter Φ, nor do they lead to valid inferences regarding the parameter. In the next section of the paper, we examine recently developed estimation procedures that result in asymptotically efficient estimates of these parameters and enable us to make valid inferences.

3 DETERMINANTS OF TREND UNEMPLOYMENT

Our approach to estimating trend unemployment begins within a single-equation framework using an unemployment equation. We examine the possibility that the unemployment rate is cointegrated with one or more of the following six structural factors:[8]

uirr	= the UI replacement rate adjusted for coverage
uidx	= Sargent's (1995) index of UI generosity
minw	= the minimum wage relative to the average hourly wage
union	= the percentage of the labour force that is unionized
paytax	= the payroll tax rate
dem	= the percentage of adult women and youths in the labour force

The measures of structural factors listed above are taken from Rose (1988), Coe (1990), and Sargent (1995). Each of the six structural factors is illustrated in Figures 2 to 7. The variables "*uirr*" and "*uidx*" are intended to proxy the generosity of the UI program. The variables "*minw*", "*union*", and "*paytax*" are each intended to capture the effect of real wage disequilibrium: *minw* proxies the effect of provincial minimum wage rates; *union* proxies the effect of wage setting in unionized sectors; and *paytax* proxies the effect of workers' resistance to bear the burden of real wage declines that arise from increases in payroll taxes. The demographic variable "*dem*" is intended to capture the effect of changes in the demographic composition of the labour force.

8. See the appendix for a detailed description of the data.

The payroll tax and UI variables call for some further discussion. First, consider payroll taxes. Standard economic theory suggests that firms will pass higher payroll taxes onto workers in the form of lower wages. Thus a rise in the payroll tax should not affect unemployment, although it may affect labour-market participation, if the reduction of wages causes some workers to leave the labour force. In fact, however, it may take some time for firms to adjust to increases in payroll taxes, and over this adjustment period, the higher payroll taxes will act to increase the real wage paid by firms and thus reduce labour demand and raise unemployment. Moreover, the evidence cited by Layard, Nickell and Jackman (1991, 210), for example, suggests that the adjustment period is long and the resulting impact on unemployment can last for at least a decade. Given this very long adjustment period, we consider payroll taxes as a potential determinant of trend unemployment, since in our statistical framework, "trend" essentially refers to factors that influence unemployment for periods longer than the average business cycle.

With respect to the UI variables, the legislated UI replacement rate is typically not used in empirical work, because it has declined over the 1971-94 period[9] and as a result, cannot account for the trend increase in unemployment observed over this period. Instead, two measures of the UI replacement rate are typically used. One measure, represented by the variable *uirr* shown by the bold solid line in Figure 2, is constructed using maximum weekly UI benefits relative to the average weekly commercial wage. This ratio is then adjusted for the coverage of the UI program. Unpublished data by the Department of Human Resources Development Canada (HRDC) indicate that in 1991 less than 7 per cent of UI claimants received maximum weekly benefits, suggesting that the maximum benefit rate is

9. The legislated UI replacement rate was reduced four times since the 1971 UI reform (embodied in the Unemployment Insurance Act, 1971), which set the replacement rate at 75 per cent and 66.7 per cent for claimants with and without dependants, respectively. The 75 per cent replacement rate for claimants with dependants was then eliminated in 1975. The 66.7 per cent replacement rate was reduced to 60 per cent in 1978 (as part of the Act to Amend the Unemployment Insurance Act, 1971) and then to 57 per cent in 1993 (as part of the Government Expenditures Restraint Act, 1993). In July 1994, the replacement rate was reduced further to 55 per cent.

not binding for the majority of UI claimants.[10] The other measure, represented by the variable *uirra* shown by the thinner line in Figure 2, is constructed using average weekly UI benefits relative to the average weekly commercial wage, again adjusted for coverage. Since the legislated UI replacement rate has changed on only four occasions since 1971, most of the variation in average weekly UI benefits is due to changes in the salary composition of UI claimants. It is not clear that compositional changes of this nature should be systematically related to the aggregate unemployment rate. Moreover, the composition effects are quite small, so that the variable *uirra* fluctuates very little over the sample period. This series acts much like a dummy variable with a discrete shift in 1972 (arising from a large increase in coverage that reflects the beginning of universal coverage).

For the reasons outlined above, we believe that proxying the generosity of the UI program using the replacement rate is not satisfactory. An alternative approach pursued in the literature is to proxy the generosity of the UI program using the ratio of the UI benefit period to the qualification period (as in Milbourne, Purvis and Scoones 1991). However, this approach may also be questionable, because the UI benefit and qualification periods are determined in part by the unemployment rate itself.[11] The ratio of the UI benefit weeks to qualification weeks is correlated, to some degree, with the aggregate unemployment rate by construction. This would lead to a simultaneity issue in estimation. It is not evident how one can construct instruments to correct for this simultaneity problem using aggregate time-series data.

Sargent (1995) has developed a new measure for proxying the generosity of the UI program. He combines and modifies the partial equilibrium model of Fortin (1984) and the general equilibrium model of

10. Andrée Houde of HRDC provided valuable information on the Unemployment Insurance Program.

11. The 1971 UI reform act (the Unemployment Insurance Act, 1971) introduced national and regional extended UI benefits that were defined with reference to past unemployment rates. The 1977 UI reform act (the Employment and Immigration Reorganization Act) eliminated national extended UI benefits but maintained regional extended UI benefits, which were given a more important role in the overall benefit package.

Milbourne, Purvis and Scoones (1991) to better reflect the Canadian UI program. For example, he considers parameters such as the maximum insurable earnings, which is not used in Fortin (1984), and the legislated replacement rate, which is not considered in Milbourne, Purvis and Scoones (1991).

Sargent‘s index is illustrated in Figure 3. His measure, the flexible time horizon variant (*uidx*), shows that the potential disincentive from the UI program has decreased on average over the last two decades.[12] This contrasts with the profile of the UI replacement rate adjusted for coverage (*uirr*). The overall downward trend in Sargent's index is explained in part by the 1977 and the 1990 reforms, which have increased the qualification periods, therefore making the UI program less generous.[13] The downward trend in Sargent's index is also explained by a decline in the legislated UI replacement rate over the 1971-94 period. Although the UI replacement rate adjusted for coverage (*uirr*) declined between 1971 and 1981 as a result of accelerating inflation, this measure resumed its upward trend thereafter owing to the lag catch-up in the maximum weekly UI benefits and the deceleration in inflation over the same period.

Given the divergence in the two measures and concerns raised earlier, our analysis will examine the UI replacement rate adjusted for coverage as well as Sargent's index for evidence of cointegration with the unemployment rate.

---

12. There are two variants of Sargent's index of UI generosity: the flexible time horizon and the 52-week time horizon. We chose the flexible time horizon, since it was available to us over the 1961Q1-95Q1 period, while the 52-week time horizon was only available from 1966Q1. In any case, the time-series behaviour of the two series is very similar; they have a correlation of the order of 0.98 over the 1966Q1-95Q1 period.

13. The increase in the qualification periods in 1990 resulted from the Act to Amend the Unemployment Insurance Act and the Employment and Immigration Department and Commission Act.

3.1 Testing the Order of Integration

A necessary condition for cointegration is that the unemployment rate and each of the structural factors be integrated of order one (I(1)).[14] We test whether this condition is consistent with the data using two basic types of tests. First we examine the null hypothesis of nonstationarity using unit-root tests. A rejection provides evidence that the series is I(0). We then examine the null hypothesis of stationarity, in which case a rejection is evidence that the series is I(1).

3.1.1 Unit-Root Tests

The unit-root tests developed by Dickey and Fuller (1979) and Said and Dickey (1984) are perhaps the most commonly used procedures for testing the nonstationarity (unit-root) hypothesis in the univariate framework. In its general form, the augmented Dickey-Fuller (ADF) test involves estimating the following regression:

$$\Delta y_t = \mu + \rho y_{t-1} + \tau t + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + \varepsilon_t \tag{3.1}$$

where Δ is the first-difference operator, μ is a constant, t is a deterministic time trend and ε_t is a random (i.i.d.) error term.[15] One can test the nonstationarity hypothesis H_0: $\rho = 1$ against the trend-stationary alternative hypothesis H_1: $\rho < 1$ using the t-statistic corresponding to ρ, which we will denote by $\hat{\tau}_\tau$. Under the null hypothesis, y_t is I(1) with drift, whereas under the alternative, y_t is I(0) around a linear deterministic trend (τt). In the case where the variable y_t does not have a significant drift component, the deterministic linear time trend (τt) is excluded from the ADF regression (3.1). In this case, the nonstationarity hypothesis H_0: $\rho = 1$ is tested against the mean-stationary alternative hypothesis H_1: $\rho < 1$ using the t-statistic corresponding to ρ, which we will denote by $\hat{\tau}_\mu$. In the “no-drift” case, y_t is

---

14. Strictly speaking, the variables that constitute the cointegration vector must be at least I(1). We confine our analysis to the special case where all variables are I(1).

15. The ADF test requires that the order of the autoregressive lag structure (k in equation 3.1) is specified appropriately. Following Ng and Perron (1995), we begin with eight autoregressive lags and sequentially reduce the number of lags until the t-statistic on the longest lag is statistically significant at the 0.10 level.

I(1) without drift under the null hypothesis and is I(0) around its mean under the alternative.

Simulation studies by Phillips and Ouliaris (1990) and Stock (1992) show that ADF t-tests have fairly reliable size properties but are less powerful than some alternative unit-root tests, such as the parameter bias test developed by Phillips (1987) and Phillips and Perron (1988). The Phillips-Perron (PP) parameter bias test involves estimating the following regression using OLS:

$$y_t = \mu + \alpha y_{t-1} + \beta t + \varepsilon_t \tag{3.2}$$

to obtain $\tilde{\alpha}$, which is used to calculate the test statistic $Z(\tilde{\alpha})$.[16] Since (3.2) includes a linear time trend (βt), y_t is I(1) with drift under the null hypothesis, whereas under the alternative, y_t is I(0) around the linear deterministic trend (βt). As in the ADF framework, we also examine the case where y_t does not have a significant drift component by excluding the deterministic linear time trend (βt) from (3.2). Following the notation used by Phillips and Perron (1988), we denote the parameter bias test statistic calculated in the no-drift case as $Z(\hat{\alpha})$.

In addition to the unit-root tests, we "shift the burden of the proof" by testing the stationary null hypothesis against the nonstationary alternative hypothesis (see Amano and van Norden 1992). We apply the testing procedure proposed by Kwiatkowski et al. (1992) (henceforth KPSS), which can be described with reference to the following model:

$$y_t = \beta t + r_t + \varepsilon_t \tag{3.3}$$

$$r_t = r_{t-1} + \nu_t \tag{3.4}$$

16. $Z(\tilde{\alpha})$ is calculated as: $Z(\tilde{\alpha}) = T(\tilde{\alpha} - 1) + \tilde{\lambda}/M$, where $\tilde{\lambda} = (1/2)\left(\tilde{\sigma}^2_{Tl} - \tilde{s}^2\right)$ (see Phillips and Perron 1988, 10, for the definition of M). The term $\tilde{\lambda}/M$ basically corrects the conventional OLS estimate of the variance $\tilde{s}^2 = \sum \tilde{\varepsilon}_t$ for the effect of serial correlation and / or heteroscedastic innovations using a non-parametric estimate of the variance. In our analysis, we obtain a non-parametric estimate of $\tilde{\sigma}^2_{Tl}$ using the VAR prewhitened quadratic spectral kernel estimator proposed by Andrews and Monahan (1992). This was implemented using a RATS procedure provided by Robert Amano and Simon van Norden.

where r_t is the random walk component and $\nu_t \sim (0, \sigma_\nu^2)$, that is, a stationary error. The KPSS test is based on the insight that if $\sigma_\nu^2 = 0$, then y_t is trend-stationary. Kwiatkowski et al. derive a test statistic η_τ that enables them to test the trend-stationarity hypothesis H_0: $\sigma_\nu^2 = 0$ against the alternative (nonstationary with drift) hypothesis H_1: $\sigma_\nu^2 > 0$.[17] For the "no-drift" case, the linear time trend βt is excluded from (3.3), so that one can test the mean-stationarity hypothesis H_0: $\sigma_\nu^2 = 0$ against the alternative (nonstationary without drift) hypothesis H_1: $\sigma_\nu^2 > 0$ using the KPSS test statistic η_μ.

3.1.2 Results of Unit-Root Tests

The unemployment rate is illustrated in Figure 1 over the 1953-94 period. The first difference of the unemployment rate (Δu_t) exhibits little drift over this period and hence, we test the unit root hypothesis against the mean-stationary alternative hypothesis (and vice versa for the KPSS test). The ADF and PP tests reported in the upper panel of Table 1 are able to reject a unit root in Δu_t (at the 0.025 level and 0.01 level, respectively), while the KPSS test cannot reject the mean-stationarity hypothesis. These tests indicate that Δu_t is I(0) without drift. Since the level of the unemployment rate (u_t) exhibits evidence of drift, we test the unit root with drift hypothesis against the trend-stationary alternative hypothesis (and vice versa for the KPSS test). The ADF and PP tests reported in the lower panel of Table 1 are unable to reject a unit root in u, while the KPSS is unable to reject the trend-stationary hypothesis.[18] Nevertheless, the balance of the statistical evidence indicates that the level of the unemployment rate (u) has behaved like an I(1) variable with drift over our sample period. It should be mentioned that modelling a variable, such as the unemployment rate, which is bounded between 0 and 1 as an I(1) variable cannot be literally true, since it implies that the variable has an unconditional infinite variance. On the other hand, the possibility that the unemployment rate is I(0) around a

17. η_τ is calculated as: $\eta_\tau = T^{-2}\sum S_t^2 / \tilde{\sigma}_{Tl}^2$ where $S_t^2 = \sum \hat{\varepsilon}_t$. As for the PP parameter bias test, we obtain a nonparametric estimate of $\tilde{\sigma}_{Tl}^2$ using the procedure proposed by Andrews and Monahan (1992) which is implemented using a RATS procedure provided by Robert Amano and Simon van Norden.

18. As shown by Amano and van Norden (1992), the KPSS test has questionable finite sample properties.

time trend would literally imply that its underlying trend is a deterministic process. We find this explanation (or lack thereof) unappealing from an economic perspective. On this basis, we proceed under the assumption that the unemployment rate is I(1) with drift for the sample period under study.

We next consider the six structural factors (illustrated in Figures 2 to 7) using the same methodology as for the unemployment rate. Stationarity tests performed on the first differences of the structural factors are reported in the upper panel of Table 1. For five of the six structural factors {*union*, *paytax*, *minw*, *uirr* and *uidx*} the ADF, PP and KPSS tests indicate that the first difference of each series is mean-stationary (at conventional levels). The exception is the demographic variable *dem* for which the PP test rejects the unit-root hypothesis, while the KPSS test rejects the mean-stationarity hypothesis. This contradiction may be due to the fact that this series displays strong negative serial correlation, in which case the PP and KPSS tests are known to suffer from severe size distortion problems (see Stock 1992). The ADF test, which is known to be more reliable in this situation, cannot reject the unit root in the first difference of *dem*, suggesting that *dem* is I(2). Stationarity tests performed on the levels of the structural factors are reported in the lower panel of Table 1. The ADF and PP tests cannot reject a unit root in the levels of the six series, while the KPSS test rejects trend stationarity only for one series — *dem*. Taken together, these tests cannot determine whether the structural factors — *union, paytax*, *minw*, *uirr* and *uidx* are I(1) with drift or whether they are I(0) around a time trend.[19] We proceed under the assumption that the levels of all six structural factors are I(1) with drift.

In addition to the unemployment rate and the six structural variables, our analysis will also consider two cyclical variables that are intended to capture cyclical movements in the unemployment rate. One cyclical variable is the level of excess supply / demand in the product market as proxied by the log deviation of real GDP from its potential level: (y_t - yp_t).

19. Since stationarity tests are generally known to lack power, particularly in small samples, it is often difficult to characterize the order of integration of the series with a high degree of certainty (see Delong et al. 1992a and 1992b, and Stock 1992).

Potential output is measured by applying the Hodrick-Prescott (HP) filter to the log of real GDP. The other cyclical variable corresponds to movements in the real wage over the business cycle arising from cyclical variations in the demand for labour. The cyclical component of the real wage is proxied using the log deviation of the real wage from its equilibrium level. The latter is derived within the neoclassical framework so that in equilibrium, labour is paid its marginal product as given by:

$$w_t - pfc_t = mpl_t \tag{3.5}$$

where w_t is the log of the nominal wage rate, pfc_t is the log of the producer price at factor cost and mpl_t represents the log of the marginal product of labour derived using a conventional Cobb-Douglas production function.

Our analysis assumes that the “output gap” (y_t - yp_t) and the “real wage gap” (w_t - pfc_t - mpl_t) are cyclical phenomena. In statistical terms, these “gaps” are assumed to be I(0) and hence cannot account for the stochastic trend in the unemployment rate.[20] Because potential output is measured using the HP filter in this paper, the output gap is mean-stationary by construction. We can, however, test whether the real wage gap is I(0).

20. The nominal wage rate (w) which is defined as total labour income per employed persons (see Appendix 1) can be decomposed into two components:

$w = ws*w + paytax*w$,

where *paytax* represents the average payroll tax rate and *ws* represents wages and salaries as a proportion of total labour income. The nominal wage rate (w) represents the marginal cost of labour as paid by the firm whereas $ws*w$ represents the marginal payment to labour received by the worker. Payroll taxes drive a wedge between the marginal cost of labour and the marginal compensation received by workers. In equilibrium, the firm uses labour inputs to the point where the marginal product of labour is equal to its marginal cost which includes wages and salaries as well as payroll taxes. In the long-run, the nominal wage rate (inclusive of payroll taxes) is determined by the value of the marginal product of labour. In this sense, workers ultimately bear the burden of payroll taxes since they are paid the value of their marginal product less payroll taxes (i.e. workers are paid: $ws*w = (1-paytax)*w$). In this case, payroll taxes should not have any long-run effect on the unemployment rate. However, payroll taxes can affect the unemployment rate as long as it drives a wedge between the marginal cost of labour paid by the firm (i.e. real wages) and its marginal product.

The real producer wage (w_t - pfc_t) and the marginal product of labour (mpl_t) are illustrated in the upper panel of Figure 9.[21] The deviation of the real wage from the marginal product of labour (w_t - pfc_t - mpl_t) is illustrated in the lower panel of Figure 9. Based on casual observation, the real producer wage and the marginal product of labour seem to be cointegrated. This conclusion is supported by two stationarity tests out of three reported in Table 2, which is reassuring, since standard economic theory predicts that workers will be paid their marginal product. Based on the balance of the evidence, the real wage gap is therefore considered an I(0) process, implying that it has only a temporary influence on the unemployment rate.

3.2 Estimating the Long-Run Parameters

We estimate the long-run parameters defining the trend unemployment rate using five alternative estimation procedures, the first of which is the two-step estimation procedure proposed by Engle and Granger (1987). The first step in the Engle-Granger (EG) procedure involves estimating the hypothesized cointegration vector using a static regression. For example, one can estimate the long-run relationship between u_t and ΦS_t by applying OLS to the following static regression:

$$u_t = \Phi S_t + \eta_t \tag{3.6}$$

where the "residual" η_t is I(0) under the cointegration hypothesis. We will refer to OLS estimates of Φ obtained from the static regression (3.6) as Φ^{EG}.

The EG estimation procedure has been perhaps the most widely-used procedure for estimating cointegrating relationships in applied work. The main advantage of the EG approach is that one does not have to specify the dynamics of the model in order to estimate the long-run parameters. This avoids the numerous specification issues encountered in determining the dynamics of the model. Unfortunately, simulation studies indicate that

21. It should be noted that the wage and marginal product measures are not defined in terms of hours but only in terms of the number of total employed persons, excluding the armed forces. See the appendix for further details.

the estimates of the long-run parameters obtained from the static regression (Φ^{EG}) have poor finite sample properties (see Inder 1993). Although Φ^{EG} is super-consistent, it is not asymptotically efficient, because it does not take into account information on the underlying dynamics of the model.

One way to take the model dynamics into account is through the fully modified (FM) procedure, developed by Phillips and Hansen (1990). The FM procedure basically involves estimating the static regression (3.6) using a semiparametric method to correct for possible serial correlation in the "residuals" η_t.[22] The FM procedure has the same advantage as the EG procedure in that one does not have to specify the dynamics of the model in order to estimate the long-run parameters. However, unlike the EG procedure, the FM procedure results in asymptotically efficient estimates, enabling us to make valid inferences regarding Φ. We will refer to the long-run estimates of Φ obtained from the FM procedure as Φ^{FM}.

One can also obtain asymptotically efficient estimates of the long-run parameters by estimating the underlying dynamics using a parametric approach. Most of the existing empirical work on measuring the trend unemployment rate involves estimating the autoregressive distributed lag specification and then calculating the long-run parameters Φ (e.g., Burns 1990, Coe 1990). As mentioned earlier, this is equivalent to estimating Φ within the error-correction model (2.4), where Z_t are stationary cyclical variables. We will refer to the estimates of Φ obtained along these lines as Φ^{ECM}.

$$C(L)\Delta u_t = D(L)\Delta S_t + E(L)Z_t - \gamma[u_{t-1} - \Phi S_{t-1}] + \nu_t \qquad \text{(eqn. 2.4)}$$

Phillips and Loretan (1991) show that estimating long-run parameters (such as Φ) within the error-correction model does not generally result in asymptotically efficient estimates. They have developed an estimation

22. We obtain a semiparametric estimate of the long-run covariance matrix corresponding with η_t using the VAR prewhitened quadratic spectral kernel estimator proposed by Andrews and Monahan (1992). This was implemented using a GAUSS procedure provided by Bruce Hansen.

procedure that does.[23] The Phillips and Loretan (PL) procedure can be implemented using the following dynamic specification:

$$F(L)(u_t - \Phi S_t) = G(L)\Delta S_t + \xi_t \tag{3.7}$$

where F(L) represents a polynomial lag operator, while G(L) represents a vector of polynomial lead/lag operators; ξ_t is a random error term.[24] The PL procedure normalizes the dependent variable and extends the dynamics of the error-correction model (2.4) to include leads as well as lags of the explanatory variables. We will refer to estimates of Φ obtained from estimating (3.7) as Φ^{PL}. Finally, the estimates of the long-run parameters using the Stock and Watson (1993) leads-and-lags procedure will also be presented and referred to as Φ^{SW}.[25]

The estimation procedure proposed by Phillips and Loretan does not involve cyclical variables. Phillips and Loretan show that their procedure results in asymptotically efficient estimates of the cointegrating parameters without modelling the complete dynamics underlying the system. There is a special case where the estimates of Φ from the error-correction model are asymptotically efficient. Phillips and Loretan (1991) show that this arises when the explanatory cyclical variables, Z_t, from the error-correction model, are strongly exogenous (as defined by Engle, Hendry, and Richard 1983). In our analysis, it is unlikely that the output gap and the real wage gap are strongly exogenous with respect to the unemployment rate. Moreover, Van Rijckeghem (1993) presents some statistical evidence using Granger causality tests, which suggest that some of the structural factors as well are not strongly exogenous with respect to the

23. This estimation procedure has been considered in numerous papers, including Phillips (1991), Phillips and Loretan (1991), Saikkonen (1991) and has been extended to higher order integrated systems by Stock and Watson (1993).

24. For example, a polynomial lead/lag operator of order (2n+1) is defined as: $G(L)x_t = \sum_{i=-n}^{n} G_i x_{t+i}$.

25. The Stock and Watson procedure is implemented using a RATS program in which the long-run covariance matrix is computed by means of the Newey-West (1987) estimator.

unemployment rate.[26] In the absence of strong exogeneity, the FM, PL and SW estimates of the cointegrating parameters Φ are asymptotically efficient, whereas the EG and ECM estimates are not. In addition, simulation studies by Phillips and Loretan (1991) and Stock and Watson (1993) indicate that the FM, PL and SW estimates have more desirable finite sample properties than the EG and ECM estimates.

Instead of examining all combinations of possible cointegrating vectors involving the six structural factors listed above, we pursued a "general-to-specific" testing procedure in an attempt to isolate a unique cointegration vector. We therefore present our main results beginning with the general specification given by:

$$S_t = \Phi_1 uirr \text{ (or } uidx\text{)} + \Phi_2 minw + \Phi_3 union + \Phi_4 paytax + \Phi_5 dem \qquad (3.8)$$

Since the data series, *uirr, uidx* and *minw* are available only beginning in 1961, we initially confine our analysis to the 1963-94 period. Also, as discussed earlier, we examine alternatively the UI replacement rate adjusted for coverage (*uirr*) and Sargent's index of UI generosity (*uidx*), along with the other structural factors.

Estimates of the long-run parameters corresponding to (3.8) obtained using the five estimation procedures over the 1963Q1-94Q4 period are reported in the first panel of Table 3.[27] Four of the five estimation procedures result in a negative estimate of Φ_5, implying that *dem* has a negative effect on the unemployment rate in the long run, which is contrary to our priors. Excluding *dem* from the vector results in:

$$S_t = \Phi_1 uirr \text{ (or } uidx\text{)} + \Phi_2 minw + \Phi_3 union + \Phi_4 paytax \qquad (3.9)$$

Estimates of the long-run parameters corresponding to (3.9) are reported in the second panel of Table 3. All five estimation procedures result in nega-

26. Van Rijckeghem finds that the unemployment rate *Granger causes* three structural factors used in our analysis: *minw, union* and *paytax*.

27. The estimation results with the UI replacement rate adjusted for coverage (*uirr*) or Sargent's index of UI generosity (*uidx*) are qualitatively the same. Thus, only the results based on the UI replacement rate adjusted for coverage are reported in Table 3.

tive estimates of Φ_1 and Φ_2, implying that *uirr* (or *uidx*) and *minw* have negative effects on the unemployment rate in the long run, which is contrary to our priors. We choose to exclude *minw* from the vector, since its estimated parameter, Φ_2, is larger compared with that of Φ_1 across the five estimation procedures. Excluding *minw* from the vector results in:

$$S_t = \Phi_1 uirr \text{ (or } uidx) + \Phi_3 union + \Phi_4 paytax \qquad (3.10)$$

Estimates of the long-run parameters corresponding to (3.10) are reported in the third panel of Table 3. All five estimation procedures result in a negative estimate of Φ_1, implying that *uirr* (as well as *uidx*) has a negative effect on the unemployment rate in the long run, which again is contrary to our priors.

The absence of significant positive effects of changes in the generosity of the UI program is somewhat surprising, given the evidence presented by others (see Rose 1988; Coe 1990; Milbourne, Purvis and Scoones 1991; Card and Riddell 1993; van Rijckeghem 1993; Corak 1994; and Fougère 1995) that UI is an important determinant of unemployment.[28] Our finding that there is no significant effect of *uirr* or *uidx* on trend unemployment may be indicative of two facts: 1) the *uirr* variable is more of a dummy variable that shifts discretely in 1971-72 and is relatively stationary thereafter (see Figure 2); 2) although Sargent's index shifts like a dummy as well in 1971-72, its decline over the 1971-94 period seems unable to account for the trend increase in unemployment observed over the same period (see Figure 3). Thus, it may be that the generosity of the UI program (as proxied by *uirr* or *uidx*), while an important determinant of the natural rate, is not a structural factor for trend unemployment. In this sense, *uirr* and *uidx* might better belong among the other structural factors (S^O in equation 2.3) instead of in S^T.

Excluding *uirr* (or *uidx*) from the vector results in:

---

28. Evidence from microeconomic data suggests that the UI program does have important effects on labour market behaviour. This is particularly the case for seasonal workers and those marginally attached to the labour force. See Corak (1994) for an overview of this evidence.

$$S_t = \Phi_3 union + \Phi_4 paytax \tag{3.11}$$

Estimates of the long-run parameters corresponding to (3.11) are reported in the fourth panel of Table 3. Note that the estimates of the long-run parameters obtained from the five estimation procedures are qualitatively the same. These estimates indicate that both *union* and *paytax* have a positive effect on the unemployment rate in the long run. Moreover, the t-statistics obtained from the ECM, PL, FM and SW procedures indicate that the estimated parameters of both of these variables are statistically significant (at the 0.10 level).[29]

As mentioned previously, the ECM estimates are asymptotically efficient only for the special case where the right-hand side variables, in this case the structural factors, are strongly exogenous. Given the similarity between the ECM estimates and the asymptotically efficient PL, FM and SW estimates, our results suggest that the strong exogeneity condition is of minor importance for the purpose of estimating Φ.[30]

In order to simplify our presentation, we will focus on the estimates of Φ obtained using the PL procedure. These results are, however, qualitatively the same as those obtained using the EG, ECM, FM and SW procedures. The first line in Table 4 refers to the vector: $u - \Phi_3 union - \Phi_4 paytax$. If we exclude *paytax* from this vector to obtain: $u - \Phi_3 union$, we find that the estimate of Φ_3 increases from 0.417 to 0.847, while its t-statistic increases from 1.76 to 4.33 (see line 2 in Table 4). Alternatively, if we exclude the *union* variable to obtain: $u - \Phi_4 paytax$, we find that the estimate of Φ_4 increases from 0.484 to 0.842 while its t-statistic increases from 1.84 to 5.43 (see line 3 in Table 4). These estimates suggest that the *union* and *paytax* variables are highly collinear, which makes it difficult to identify their separate influences.[31]

29. The t-statistics derived from the EG estimation procedure are severely biased and, hence, are not reported.

30. Van Rijckeghem (1993) finds that simultaneity bias plays an important role in the analysis of the natural rate of unemployment.

31. The correlation between the variable *union* and *paytax* is of the order of 0.82 over the 1963Q1-94Q4 period and 0.81 over 1955Q1-94Q4.

Recall that we obtained counterintuitive (negative) long-run effects associated with the structural factors *uirr* (or *uidx*), *minw* and *dem*. If we exclude these variables from our analysis, we can extend the sample period back to 1955. Over the 1955-94 period, the results are qualitatively the same as those obtained over the shorter sample. The variables *union* and *paytax* can account for the stochastic trend in the unemployment rate over the 1955-94 period, although we cannot identify their individual influence.

Before we move on to further tests, an important caveat must be acknowledged: the counterintuitive long-run effects associated with certain of our structural factors may be due to measurement error in some variables, to multicollinearity amongst the explanatory variables, or to the possibility that an important long-run determinant of unemployment is omitted from the analysis. Accordingly, it is quite possible that some additional data, that is, more information in the data set, would alter the outcome of the process of elimination followed here.

3.3 Testing for Cointegration

Up to this point, we have examined estimates of long-run parameters relating the unemployment rate to various structural factors. We now examine whether one of the long-run relationships forms the basis of a cointegrating relationship.

We test for cointegration using “residual-based” versions of the ADF, PP and KPSS tests outlined in the previous section. The ADF test has been extended to the cointegration framework by Engle and Granger (1987). Engle and Granger show that one can test for cointegration by applying the ADF unit-root test to OLS residuals obtained from a static regression. For example, testing for cointegration between u_t and ΦS_t entails applying the ADF test to the estimated “residuals” η_t obtained from the static regression (3.6). The PP parameter bias test can also be used to test for cointegration in a similar manner (see Phillips and Perron 1988). As mentioned above, estimates of long-run parameters obtained from the static regression (Φ^{EG}) have poor finite-sample properties. This is likely to affect the cointegration tests. For this reason, we also apply the ADF and PP tests using estimates of Φ obtained from the ECM, PL, FM and SW procedures.

We "reverse the burden" of the cointegration tests by testing the cointegration null hypothesis against the non-cointegration alternative hypothesis using the KPSS procedure, which has been extended to the cointegration framework by Shin (1992). Shin shows that this extension is valid, provided that the estimates of the cointegration vector are asymptotically efficient. As mentioned above, only the estimates obtained using the FM, SW and PL procedures satisfy this criterion. Hence, we apply the KPSS test only to estimates of Φ obtained from the FM, SW and PL procedures.

Results of cointegration tests are reported in Table 4 for the cases where the estimated long-run parameters have the a priori expected (positive) sign. The reported results are based on the PL estimates; however, the same inferences are made using estimates obtained from the ECM, FM and SW procedures.

The upper panel in Table 4 refers to estimates obtained over the 1963-94 period. The first line in Table 4 corresponds to the vector: $(u - \Phi_3 union - \Phi_4 paytax)$. The ADF, PP and KPSS tests provide strong evidence of cointegration between the variables — one can reject the *non-cointegration* hypothesis on the basis of the ADF test (at the 0.10 level) and the PP test (at the 0.05 level), whereas one cannot reject the *cointegration* hypothesis on the basis of the KPSS test (even at the 0.10 level). When the variable *paytax* is excluded from the vector, we find strong evidence of cointegration between *u* and *union* (see line 2 in Table 4). When the variable *union* is excluded from the vector, we find mixed evidence of cointegration between *u* and *paytax*. One can reject the non-cointegration hypothesis on the basis of the PP test (at the 0.05 level) but not on the basis of the ADF test (even at the 0.10 level), whereas one cannot reject the cointegration hypothesis on the basis of the KPSS test. Overall, the cointegration tests suggest that there is a cointegrating relationship between the unemployment rate and the structural factors *union* and *paytax*, but we are not very confident about the relative size of their respective estimated parameters.

The lower panel in Table 4 refers to estimates obtained over the longer sample period. The ADF and PP statistics corresponding to the three vectors generally increase when the sample period is extended back to 1955, providing stronger evidence of cointegration. Moreover, we no longer find mixed evidence of cointegration between *u* and the vector involving the *paytax* variable alone. This result would be expected if these were actually cointegrating relationships, since increasing the number of observations increases the power of the ADF and PP tests.

4 SENSITIVITY ANALYSIS

As shown by Rose (1988) and Setterfield, Gordon and Osberg (1992), estimates of structural models of the unemployment rate tend to be quite fragile, in the sense that seemingly innocuous changes in model specification can have important implications for the estimated measure of trend unemployment. To address this concern, we examine the robustness of our results by performing an extensive sensitivity analysis. We illustrate our main findings with reference to the vector: $(u - \Phi_3 union - \Phi_4 paytax)$, summarized by the estimates reported in Tables 5 through 9.

4.1 Extending the Estimation Sample Period

We first examine the estimates of the long-run parameters derived from the alternative estimation procedures when the estimation sample period is extended back to 1955. When we compare the estimates obtained over the longer sample period reported in Table 5 with those of the shorter sample period reported in the fourth panel of Table 3, we find that they are very similar. This indicates that the estimates of the long-run parameters are robust with respect to extending the estimation sample period. Inferences relating to the cointegration hypothesis are also robust with respect to extension of the estimation sample period.[32]

32. Over the shorter sample period, inferences relating to the cointegration hypothesis are robust with respect to the alternative estimation procedures. Only the results based on the PL estimates are reported (see Table 4).

4.2 Stability Tests of the Estimated Long-Run Parameters

We now examine the stability of the estimated long-run parameters using the SupF, MeanF and L_c tests proposed by Hansen (1992). Each of these tests examines the null hypothesis of a stable cointegrating relationship among I(1) variables against different alternative hypotheses. The SupF test is designed to detect a discrete break in the parameters at an unknown breakpoint, while the MeanF and L_c tests are designed to detect gradual time variation in the parameters. We apply these tests using estimates obtained from the FM procedure over the extended sample period. The estimates and results of the stability tests are reported in Table 6.[33] Overall, stability tests when applied over the 1955-94 period suggest that the estimated long-run relationship between the unemployment rate and the *union* and *paytax* variables is stable.

4.3 Alternative Dynamic Specification

The estimation results reported in Table 5, which will be used as benchmark estimates, are derived with two lags and two leads on all variables. This second-order dynamic specification was selected on the basis of both Akaike's (1969) Information Criterion and Schwarz's (1978) Bayesian Information Criterion. When we apply an alternative dynamic specification to the ECM, PL and SW procedures, we generally obtain similar estimates of the long-run parameters. For example, estimates of the long-run parameters Φ_3 and Φ_4 obtained using a fourth-order dynamic specification (see Table 7) are qualitatively the same as those of the benchmark estimates (Table 5). This is, however, less the case for the ECM estimates. This indicates that the ECM estimates of the long-run parameters are more sensitive with respect to how the dynamics are modelled. The cointegration tests results reported in Table 7 are, however, robust to the dynamic specification issues of this nature.

33. The SupF, MeanF and L_c tests were implemented using a GAUSS procedure provided by Bruce Hansen.

4.4 Adding Cyclical Variables to Complement Long-Run Determinants

In our analysis up to this point, we have focussed on examining whether the unemployment rate is cointegrated with the various structural factors within a single-equation framework, using an unemployment equation without modelling the complete dynamics. We now extend the cointegration analysis by adding cyclical variables to complement the long-run determinants. The addition of the cyclical variables is performed using two different approaches, which are outlined in greater detail in the following subsections. The first approach, which is essentially the second step of the Engle-Granger procedure, is performed by adding the vector Z of cyclical variables to the error-correction model (see equation 2.4). The second approach uses the Phillips-Loretan procedure to jointly estimate the long-run factors along with the cyclical variables.

4.4.1 Two-Step Engle-Granger Procedure

First, we estimate the long-run parameters defining the trend unemployment rate using the two-step estimation procedure proposed by Engle and Granger (1987). The first step involves estimating the hypothesized cointegration vector using a static regression (see static regression 3.6). The hypothesized cointegration vector is reported in line 1 of Table 8. In the second step, we estimate an error-correction model (see equation 2.4) in which the change in unemployment is regressed on the residuals of the static regression (3.6) from the first step, along with the inclusion of the vector Z of cyclical variables to model unemployment dynamics.[34] An estimate of the error-correction term (γ) that is negative and statistically significant ensures that the unemployment rate converges towards its structural determinants in the long run and provides further evidence of cointegration.

Alternative error-correction models including different cyclical variables are estimated. Line 2 of Table 8 is the benchmark specification in

34. The Granger Representation Theorem states that if two variables (or a variable versus a vector of variables) are cointegrated, then there exists an error-correction model that can capture the dynamics underlying the cointegrating relationship between the variables (see Engle and Granger 1987).

which we include the first difference of the long-run determinants only. The estimated parameter associated with the error-correction term is negative, as expected, and statistically significant (at the 0.01 level). In addition to this benchmark specification, we examine a specification that includes the output gap series, defined using a measure of potential output generated by the HP filter with a "smoothness parameter" of 1600.[35,36] The estimated parameter associated with the error-correction term, which is reported in line 3 of Table 8, is again negative and statistically significant (at the 0.05 level). The inclusion of the real wage equilibrium condition in the error-correction model does not alter very much the results of the benchmark specification (line 4).

We also consider the implications of including other cyclical factors in the analysis. Following Coe (1990), we include the relative price of energy along with a terms of trade variable (see line 5 of Table 8).[37] Again, the estimated parameter associated with the error-correction term is of the expected sign and statistically significant (at the 0.01 level).

Finally, we also examined a specification that includes a measure of the term structure of interest rates defined as the differential between short-term and long-term interest rates: (*rs* - *rl*) as a cyclical variable in place of the output gap (see line 6 of Table 8). The estimated parameter associated with the error-correction term remains negative and statistically significant (at the 0.05 level).

---

35. The "smoothness parameter" in the HP filtering framework basically determines the variation in the trend component of the series relative to that in the cyclical component. The smoothness parameter is typically set to a value of 1600 in empirical applications.

36. The estimation results with alternative measures of the output gap indicate that the estimated parameter associated with the error-correction term is robust with respect to the measurement of the output gap. Thus, only the results based on the measure of potential output generated by the HP filter with a "smoothness parameter" of 1600 is reported in Table 8.

37. The relative price of energy is measured using the variable *rpeng*, while the terms of trade are measured using the relative price of (non-energy) commodities, *rpcne*. Given that these variables are both I(1), they are specified in difference form so that they affect unemployment dynamics but have no long-run effect on the unemployment rate. As we emphasize in the introduction, our analysis does not encompass the entire set of structural factors that have been proposed in the literature. Hence, issues such as sectoral shifts arising from movements in the terms of trade are beyond the scope of our analysis.

To sum up, the estimation results of the various error-correction models reported in Table 8 provide further evidence of cointegration between the unemployment rate and the structural factors *union* and *paytax* over the 1955-94 period.

4.4.2 Joint Estimation of the Long-Run Determinants and the Cyclical Variables

After extending the cointegration analysis by adding cyclical variables to complement the long-run factors using a two-step procedure, we now consider their joint estimation using the Phillips-Loretan procedure. For the PL procedure, this entails estimating equation (3.7) with the vector Z of stationary cyclical variables. We perform a sensitivity analysis in order to examine the robustness of our results when long-run factors and cyclical variables are jointly estimated. We illustrate our main findings with reference to the vector: $(u - \Phi_3 union - \Phi_4 paytax)$. More specifically, we examine whether the benchmark estimates reported in line 1 of Table 9 (reported previously in Table 4 and in Table 5) are sensitive to the inclusion of cyclical influences such as the output gap, the real wage gap, commodity prices and the term structure of interest rates.

We first examine a specification that includes the output gap series defined using a measure of potential output generated by the HP filter with a "smoothness parameter" of 1600. This measure of potential output is admittedly arbitrary, and hence we consider alterative measures as well. We apply the HP filter with the smoothness parameter set alternatively to 500 and 10 000. This essentially enables us to vary the amplitude and duration of the measured output gap. The resulting estimates of the long-run parameters Φ_3 and Φ_4 are reported in lines 2, 3 and 4 of Table 9. The inclusion of the output gap results in large changes in the estimated long-run parameters relative to the benchmark specification. The estimated parameter associated with the *union* variable (Φ_3) increased from 0.4 to 0.7, while that of the *paytax* variable (Φ_4) decreased from 0.3 to about 0.05. Hence, the estimates of the long-run parameters Φ are sensitive to the inclusion of the output gap when they are jointly estimated. They seem to be robust, however, with respect to its measurement.

When the real wage equilibrium condition specified in level form is added to the benchmark specification, the long-run parameters Φ_3 and Φ_4 change somewhat relative to the benchmark specification but not as much as is the case for the output gap (see line 5 in Table 9). We also consider the implications of including the relative price of energy along with a terms of trade variable. In this case, the estimated long-run parameters are very similar to those of the benchmark specification (see line 6 of Table 9). We also examine a specification that includes a measure of the term structure of interest rates: (*rs* - *rl*). As shown, the results are qualitatively similar to those of the benchmark specification (line 7 of Table 9).

Finally, we consider the impact of simultaneously including all the above cyclical factors. More specifically, we begin with a general specification that includes all the cyclical variables along with the long-run factors *union* and *paytax*. We then reduce this general specification by eliminating short-run variables that yield estimated parameters that are not statistically significant.[38] The long-run parameters estimates Φ_3 and Φ_4 of this reduced specification are reported in line 8 of Table 9 (see also trend_eq8 in Figure 11).[39] As can be seen, including cyclical variables that are statistically significant to the long-run factors results in estimated long-run parameters Φ_3 and Φ_4 that change somewhat relative to the benchmark specification. In particular, the estimated parameter associated with the *union* variable (Φ_3) rises from 0.4 to 0.6, while that of the *paytax* variable (Φ_4) falls from 0.3 to 0.15. Note also that richer dynamics improve the precision of the estimated long-run parameters relative to the other specifications. As a result, although the estimated parameter on *paytax* is about half as large as in the baseline specification, it is nonetheless statistically significant at the 0.10 level.

38. We tested the joint hypothesis that the parameters on the excluded cyclical variables are zero and found that we cannot reject this exclusion restriction at the 0.10 level using the likelihood ratio test (LR test).

39. This reduced specification includes the following cyclical variables: second lead, contemporaneous value and second lag of the output gap, and the second lag of the first difference of the relative price of non-energy commodities.

In an effort to improve the precision of the estimates further, we also attempted to add further information to our regression model. In particular, following Ford and Rose (1989), we extended our cointegration analysis by considering a bivariate system that consists of the unemployment equation examined previously, along with a wage Phillips curve. By using more than one source of information, the systems approach has the potential to be helpful in the joint estimation of the cyclical factors and the long-run determinants. Unfortunately, our results to this point on this front have been disappointing. This is not unique to structural models of unemployment and is also pervasive in the study of business cycles (see Quah 1992).

4.5 Payroll Taxes and the Recent Cycle

To provide some insight into the empirical importance of the fragility problem, we compare alternative measures of the trend unemployment rate based on the estimates reported in Table 9. Figure 12 compares two measures of the trend unemployment rate, each corresponding to the vector $\Phi_3 union + \Phi_4 paytax$.[40] One measure is based on a high estimate of Φ_3 (represented by trend_eq3); the other is based on a higher estimate of Φ_4 (represented by trend_eq6). These two measures of the trend unemployment rate move together in a broad manner throughout most of the sample period. The difference between the measures has a standard deviation of 0.60 of a percentage point over the 1961-94 period (0.67 of a percentage point over the 1955-94 period). The divergence between the two measures increased to 1.5 percentage points in 1994, however. These calculations show that the fragile nature of the estimates can give rise to empirically important differences in measures of the trend unemployment rate during certain episodes. They also underscore the potential benefits of incorporating a variety of elements when constructing estimates of potential output, as opposed to relying on a single empirical relationship.

40. The constant term or intercept of trend unemployment rate is calculated as the mean, over the 1955Q1-94Q4 period, of the difference between the observed unemployment rate and the cointegrating vector.

Figure 12 also illustrates that the measure of the trend unemployment rate based on the larger estimated parameter on the *paytax* variable rises substantially over the 1990-94 period. This coincides with an increase in payroll taxes (see Figure 6), that is due primarily to an increase in UI premiums. The increase in UI premiums may have in turn been related, at least in part, to an increase in the unemployment rate.[41] Hence, there may be simultaneity between payroll taxes and the unemployment rate, which would tend to bias our estimates. In other words, the actual rise in trend unemployment is likely to have been less than suggested by the figure. Such interpretation leads to the use of judgement when implementing the multivariate filter to estimate potential output.

Further examination of this issue reveals that our inferences do not hinge on the 1990-94 period, however.[42] When we restrict our estimation analysis to the 1955-89 period, the estimated parameter on the payroll tax variable remains virtually unchanged, although the confidence intervals associated with our point estimates are wider. Nevertheless, the rise in payroll taxes over the 1990Q1-93Q1 period (and to a lesser extent over the 1993Q2-94Q4 period) does have important empirical implications for our measures of the trend unemployment rate. Our estimates suggest that payroll taxes may have increased the trend unemployment rate by over 1 percentage point between 1989 and 1994.

5 Conclusions

The Bank's staff have adopted a methodology called the multivariate filter for the purpose of estimating potential output. The objective of the present paper is to identify the trend unemployment rate, an empirical concept, which we define in terms of cointegration, for use within the multivariate filter. As such, the paper does not deal explicitly with either the

41. The increase in UI premiums was legislated through the 1990 reform act (see footnote 13), which made the UI account self-financing. Subsequent increases in UI premiums can be attributed to the sharp increase in the unemployment rate beginning in late 1990, which was unanticipated at the time.

42. The estimation analysis is performed here, over the 1955-89 period, with long-run factors only. The results are then compared with those of the benchmark specification (long-run factors only) reported in line 1 of Table 9.

natural rate of unemployment, or the NAIRU. Trend unemployment can be thought, however, as being part of a more general specification of the NAIRU, while being in all likelihood more variable than the underlying natural rate. Accordingly, we think of the trend unemployment rate as conceptually between the natural rate and the NAIRU, and we see the research undertaken here as pointing the way for further work on the NAIRU.

In this paper, we examine which structural factors best account for the stochastic trend in unemployment using a cointegration approach that explicitly takes into account econometric complications that arise from the nonstationary nature of the data. This basically involves determining whether the unemployment rate is cointegrated with various structural factors that have been proposed in the literature. Our analysis is limited to the structural factors proposed by Rose (1988), Coe (1990) and Sargent (1995).

The main conclusion that we draw from our analysis is that the stochastic trend in the unemployment rate can best be explained by two structural factors: the degree of unionization in the labour force and payroll taxes. There is little evidence that the other four structural factors considered — the demographic composition of the labour force, the UI replacement rate, Sargent's index of UI generosity and the minimum wage rate — have played an important role, although the UI variables may have been important episodically and thus have played an important role in determining the natural rate in certain periods. This conclusion is supported by formal statistical tests for cointegration. However, given the high degree of collinearity between the unionization and payroll tax variables, we are unable, based on the cointegration tests, to determine whether this cointegrating relationship is due to one or both of these factors. We also have difficulty in obtaining precise estimates of their individual contributions.

Nevertheless, in the single-equation framework, our inferences appear to be robust with respect to many things. This is contrary to much of the existing empirical work in this area. For example, our results are qualitatively robust with respect to different estimation procedures, to different estimation sample periods as well as to several specification issues, including stability. This is also the case with respect to inferences relating to the cointegration hypothesis. Moreover, the results are robust to the inclusion of cyclical influences when we use either a two-step or joint estimation procedure. The joint estimation results are, however, quantitatively sensitive to the inclusion of some cyclical variables. Nonetheless, on the basis of the work done to date, we would favour the trend unemployment rate measure implied by equation 8 of Table 9 for use within the multivariate filter. The inclusion of cyclical influences on unemployment in this equation allows us to estimate the long-run parameters with more precision.

It should also be emphasized that our analysis does not encompass the entire set of structural factors that have been proposed in the literature to explain unemployment. In particular, we have not considered the implications of sectoral shifts. There is little doubt that sectoral shifts arising from large commodity price movements in the 1970s and 1980s have had important effects on the Canadian labour market. Although we view the role of sectoral shifts to be of great importance for understanding the Canadian unemployment experience over the last few decades, the relevant modelling issues are non-trivial. In particular, it is not clear how one can best separate the effects of sectoral shifts from some of the other cyclical influences that we have considered in the present paper. The same is true for other potentially important macroeconomic variables not considered in our analysis, such as real interest rates (considered by Fortin, Keil and Symons 1995) and the real exchange rate, for example. Issues of this nature are deferred to future research.

We would also like to draw attention to the nature of the statistical evidence relating to the unemployment rate and the various structural factors. Four points are relevant in this respect. First, it is important to keep in mind that our analysis has focussed primarily on the long run. Consequently, we have little to say about how the various structural factors affect unemployment rates in the short run. Second, the confidence intervals associated with our point estimates are quite wide, making it meaningless to attempt to measure the trend unemployment rate with precision. Third, we believe that misspecification arising from measurement error plays an important role in determining which structural factors seem to be relevant for trend unemployment. We have used proxies for structural factors constructed in a manner that is consistent with previous research. In our view, these empirical proxies at best provide only rough measures of the structural factors that are of economic interest. Fourth, multicollinearity may play an important role in determining which structural factors seem to matter. Although the trend unemployment rate is best captured by the proportion of the labour force that is unionized, and payroll taxes, there is a theoretical case for the other variables to be important in identifying trend unemployment as well. It might easily be the case that with the accumulation of more data the trend would be best captured by another combination of cointegrated variables.

We would also like to stress that the scope of our analysis has been limited to being empirical in nature and does not address a number of underlying theoretical issues. In particular, we have made no attempt to provide a theoretical framework that can explain the linkages between unemployment rates, wages and structural factors such as payroll taxes and unionization. The conventional view of the labour market assumes that wages adjust to equate labour supply and labour demand in the long run. According to this view, increases in payroll taxes will eventually be offset by reductions in real wages received by employees to establish an equilibrium whereby labour is paid its marginal product. If wages are sufficiently flexible, increases in payroll taxes should have no effect on unemployment in the long run.

Our empirical results suggest that payroll taxes affect unemployment above and beyond the real wage channel. It is unclear whether this is due to theoretical shortcomings or to our inability to adequately measure theoretical concepts such as the real wage and the marginal product of labour. Nevertheless, if real wage inflexibility exists, the wedge between the marginal cost of labour paid by the firm and its marginal product may influence the unemployment rate in the long run. Also, if payroll tax increases affect labour costs for a very long period of time, say longer than a business cycle, it is not surprising that empirically we find the payroll tax variable to have a permanent effect on unemployment.

With regard to the more broadly defined research agenda, it would be of interest to examine the evidence for other countries as well as for Canada to determine whether the results uncovered in this paper can be corroborated. We are, however, somewhat sceptical that more can be learned about the underlying linkages between the unemployment rate and structural factors using aggregate time-series data, particularly with the serious measurement problems involved. We believe that longitudinal data sets, such as Statistics Canada's Labour Market Activity Survey, provide a rich source of information for many of the issues at hand, particularly relating to labour supply behaviour. It is hoped that research along these lines can improve our understanding of the various factors that determine unemployment rates and thereby complement the time series evidence.

We would like to end by briefly commenting on the policy implications of our main empirical findings. Our analysis suggests that two structural factors — the degree of unionization in the labour market and payroll taxes — can best account for the upward "drift" in the Canadian unemployment rate observed since the mid-1950s. This is not to say that other factors play no role in determining unemployment over the course of the business cycle. Our analysis has abstracted from issues relating to the cyclical component of the unemployment rate by focussing exclusively on its underlying trend. Hence, we have little to say about the cyclical nature of the unemployment problem. Our analysis offers instead a long-run perspective. According to our results, the upward "drift" in the Canadian unemployment rate must be addressed with reference to the underlying structural factors.

APPENDIX 1: Description of the Data[43]

AIB = dummy variable representing the Anti-Inflation Board:
AIB = 1 over 1976Q1-78Q2 period; 0 otherwise.

dem = the percentage of adult women and youths (aged 15-24) in the labour force.

minw = the relative minimum wage rate defined as the weighted (by labour force shares) average of provincial minimum wages relative to the average hourly wage in the manufacturing sector. The provincial minimum wages are from Human Resources Development Canada.

mpl = the marginal product of labour defined as: $\log(\alpha) + (y - n)$,
where α is labour's average share of output (0.67)
and $(y - n)$ is the log of the output/labour ratio defined as real GDP / the number of total employed persons, excluding the armed forces.

p = the (log of the) GDP price deflator at market prices (D20011/ D20463).

paytax = the payroll tax rate defined as:
supplementary labour income / wages and salaries.

pc = the (log of the) CPI excluding food and energy (B820555).

pfc = the (log of the) GDP price deflator at factor costs defined as:
nominal GDP - total indirect tax revenue less subsidies/real GDP, i.e., (D20011-D20008)/D20463.

---

43. The mnemonics given in parentheses refer to Cansim data series; the other variables have been constructed using various data sources. All data series include information available up to the end of June 1995 (including the June 1995 National Accounts). They are available upon request.

q = the (log of the) real effective exchange rate defined using a trade-weighted index of GDP/GNP price deflators for the G-6 countries.[44]

rpcne = the relative price of non-energy commodities defined as: the Bank of Canada non-energy commodity price index in $US / U.S. producer price index for finished goods.

rpeng = the relative price of energy defined as: the Bank of Canada energy commodity price index in $US / U.S. GDP deflator.

rs - rl = differential between short-term and long-term interest rates where: rs is a (nominal) yield on 90-day treasury bills (B14007) and rl is a (nominal) yield on long-term G. of C. Bond (B14013).

u = the percentage of the labour force that is unemployed defined as: the number of unemployed/total labour force, D767611 or (D767609/D767606). The Statistics Canada labour force series are available beginning in 1976. The series prior to 1976 have been linked at the Bank of Canada by the Research Department.

uidx = an index of the generosity of the Unemployment Insurance Program developed by Timothy Sargent (1995).

uirr = the maximum UI replacement rate adjusted for coverage defined as: (the maximum weekly UI benefit/the average weekly earnings in the commercial sector)*(the proportion of the labour force covered by unemployment insurance).

uirra = the average UI replacement rate adjusted for coverage defined as: (the average weekly UI benefit/the average weekly earnings in the commercial sector)*(the proportion of the labour force covered by Unemployment Insurance).

44. Trade weights for the G-6 countries are given by:
U.S.: 0.8303, Japan: 0.0598, U.K.: 0.0502, Germany: 0.0275, France: 0.0176, Italy: 0.0146.

union = the percentage of the labour force that is unionized from: "Union Membership as a Percentage of Non-Agricultural Paid Workers" in *The 1994-95 Directory of Labour Organization in Canada* (Table 1), Human Resources Development Canada.

w = the (log of the) nominal wage rate defined as: the (log of) total QPM labour income/the number of total employed persons, excluding the armed forces.

y = the (log of) domestic real GDP (D20463).

yf = the (log of) foreign real GDP constructed using a trade-weighted index of the G6 countries (see footnote 44).

Table 1

Stationarity Tests on the Unemployment Rate and Structural Factors

	Tests in the Absence of Drift [a]		
	Unit-Root Tests [b]		*Stationarity Tests* [c]
Variables	***ADF:*** $\hat{\tau}_\mu$	***PP:*** $Z(\hat{\alpha})$	***KPSS:*** $\hat{\eta}_\mu$
	Sample: 1955Q1-94Q4 (160 Observations)		
Δu	-3.35 [.025]	-67.35 [<.01]	0.025 [>.10]
Δdem	-1.65 [>.10]	-140.79 [<.01]	2.043 [<.01]
Δunion	-2.76 [.10]	-111.59 [<.01]	0.056 [>.10]
Δpaytax	-3.19 [.025]	-150.70 [<.01]	0.380 [.10]
	Sample: 1963Q1-94Q4 (128 Observations)		
Δminw	-2.48 [>.10]	-137.56 [<.01]	0.306 [>.10]
Δuirr	-4.12 [<.01]	-129.44 [<.01]	0.110 [>.10]
Δuidx	-10.56 [<.01]	-119.12 [<.01]	0.172 [>.10]
	Tests in the Presence of Drift		
Variables	***ADF:*** $\hat{\tau}_\tau$	***PP:*** $Z(\tilde{\alpha})$	***KPSS:*** $\hat{\eta}_\tau$
	Sample: 1955Q1-94Q4 (160 Observations)		
u	-2.63 [>.10]	-11.27 [>.10]	0.010 [>.10]
dem	0.70 [>.10]	1.93 [>.10]	0.254 [<.01]
union	-1.84 [>.10]	-7.99 [>.10]	0.033 [>.10]
paytax	-3.26 [.10]	-4.89 [>.10]	0.029 [>.10]
	Sample: 1963Q1-94Q4 (128 Observations)		
minw	-1.79 [>.10]	-3.47 [>.10]	0.072 [>.10]
uirr	-1.85 [>.10]	-5.96 [>.10]	0.032 [>.10]
uidx	-1.49 [>.10]	-5.87 [>.10]	0.042 [>.10]

a. In the *absence* of drift the ADF, PP and KPSS tests include a constant term but do not include a linear time trend, whereas in the *presence* of drift they include a constant term as well as a linear time trend.

b. The ADF and PP normalized bias statistics test the null hypothesis of nonstationarity (i.e. H_0: y is I(1)) against the alternative hypothesis of stationarity (i.e. H_1: y is I(0)). P-values for the ADF t-statistics and the PP normalized bias statistics (reported in square brackets) are obtained from the critical values reported by Davidson and MacKinnon (1993, Table 20.1).

c. The KPSS test statistics test the null hypothesis of stationarity (i.e. H_0: y is I(0)) against the alternative hypothesis of nonstationarity (i.e. H_1: y is I(1)). P-values (reported in square brackets) for the KPSS statistics are obtained from the critical values reported by Kwiatkowski et al. (1992, Table 1).

Table 2

Stationarity Tests on Wages, Prices and Labour Productivity

Sample: 1955Q1-94Q4 (160 Observations)

Tests in the Absence of Drift			
	Unit-Root Tests		*Stationarity Tests*
Variables	***ADF:*** $\hat{\tau}_\mu$	***PP:*** $Z(\hat{\alpha})$	***KPSS:*** $\hat{\eta}_\mu$
Δw	-2.10 [>.10]	-83.65 [<.01]	0.763 [<.01]
Δpfc	-2.10 [>.10]	-47.26 [<.01]	0.766 [<.01]
Δw – Δpfc	-4.37 [<.01]	-165.30 [<.01]	0.642 [.025]
Δmpl	-4.42 [<.01]	-163.28 [<.01]	0.578 [.025]
Δpc - Δpfc	-3.30 [.025]	-144.42 [<.01]	0.627 [.025]
w - pfc - mpl	-2.78 [.10]	-9.52 [>.10]	0.061 [>.10]
Tests in the Presence of Drift			
	Unit-Root Tests		*Stationarity Tests*
Variables	***ADF:*** $\hat{\tau}_\tau$	***PP:*** $Z(\tilde{\alpha})$	***KPSS:*** $\hat{\eta}_\tau$
w - pfc	-1.32 [>.10]	-3.51 [>.10]	0.205 [.025]
mpl	-1.30 [>.10]	-5.41 [>.10]	0.151 [.05]
pc - pfc	-1.07 [>.10]	0.179 [>.10]	0.161 [.05]
w - pfc - mpl	-2.80 [>.10]	-9.53 [>.10]	0.058 [>.10]

- See Notes to Table 1.

Table 3

Single-Equation Estimates of the Long-Run Parameters

Long-Run Factors Only

(Sample: 1963Q1-94Q4, 128 Observations)

Estimation Procedures	***Estimates of the Long-Run Parameters***
	$\Phi S_t = \Phi_1 uirr + \Phi_2 minw + \Phi_3 union + \Phi_4 paytax + \Phi_5 dem$
EG	.004uirr - .14minw + .708union + .398paytax - .288dem
ECM	-.059uirr + .045minw - .433union + .263paytax + .761dem (0.87)[a] (0.25) (0.45) (0.62) (0.95)
PL	-.033uirr + .139minw + .301union + .115paytax - .048dem (0.53) (0.67) (0.37) (0.25) (0.08)
FM	.043uirr - .209minw + 1.114union + .266paytax - .684dem (1.60) (3.34) (4.28) (1.45) (3.16)
SW	.01uirr - .166minw + 1.275union + .316paytax - .718dem (0.51) (4.17) (6.54) (3.63) (4.27)
	$\Phi S_t = \Phi_1 uirr + \Phi_2 minw + \Phi_3 union + \Phi_4 paytax$
EG	-.013uirr - .107minw + .404union + .426paytax
ECM	-.014uirr - .130minw + .474union + .276paytax (0.38) (1.44) (1.95) (0.89)
PL	-.029uirr - .09minw + .495union + .413paytax (0.67) (0.87) (1.63) (1.11)
FM	-.001uirr - .127minw + .377union + .421paytax (0.05) (2.07) (2.67) (2.15)
SW	-.032uirr - .115minw + .517union + .446paytax (1.48) (2.71) (3.22) (4.17)
	$\Phi S_t = \Phi_1 uirr + \Phi_3 union + \Phi_4 paytax$
EG	-.025uirr + .379union + .651paytax
ECM	-.043uirr + .580union + .524paytax (1.13) (2.44) (2.16)
PL	-.054uirr + .553union + .631paytax (1.36) (2.32) (2.56)
FM	-.027uirr + .434union + .683paytax (0.94) (2.62) (3.94)
SW	-.042uirr + .413union + .752paytax (1.63) (2.58) (8.87)

a. Absolute t-statistics reported in parentheses in Tables 3 to 9.

(cont'd)

Table 3

Single-Equation Estimates of the Long-Run Parameters

Long-Run Factors Only

(Sample: 1963Q1-94Q4, 128 Observations)

Estimation Procedures	***Estimates of the Long-Run Parameters***
	$\Phi S_t = \Phi_3 union + \Phi_4 paytax$
EG	.322union + .587paytax
ECM	.480union + .393paytax (2.07)[a] (1.57)
PL	.417union + .484paytax (1.76) (1.84)
FM	.355union + .601paytax (1.93) (3.17)
SW	.292union + .652paytax (2.86) (7.97)

a. Absolute t-statistics reported in parentheses in Tables 3 to 9.

Table 4

Cointegration Tests in the Single-Equation Framework

	Estimates of the Long-Run Parameters[a]	***Cointegration Tests***[b]		
	(ΦS_t)	***ADF*** $\hat{\tau}_\mu$	***PP:*** $Z(\hat{\alpha})$	***KPSS:*** $\hat{\eta}_\mu$
Line	*(Sample: 1963Q1-94Q4, 128 Observations)*			
1	.417union + .484paytax	-3.60	-26.24	0.009
	(1.76) (1.84)	[.10]	[.05]	[>.10]
2	.847union	-3.33	-20.80	0.026
	(4.33)	[.10]	[.05]	[>.10]
3	.842paytax	-2.80	-19.52	0.012
	(5.43)	[>.10]	[.05]	[>.10]
Line	*(Sample: 1955Q1-94Q4, 160 Observations)*			
4	.449union + .346paytax	-3.75	-27.14	0.010
	(1.90) (1.51)	[.05]	[.05]	[>.10]
5	.804union	-3.58	-22.55	0.026
	(4.54)	[.05]	[.05]	[>.10]
6	.689paytax	-3.35	-24.69	0.012
	(5.07)	[.10]	[.025]	[>.10]

a. The estimates reported above are obtained using the PL procedure.

b. The ADF and PP statistics test the null hypothesis of *non-cointegration* (i.e. H_0: u - ΦS is I(1)) against the alternative hypothesis of *cointegration* (i.e. H_1: u - ΦS is I(0)). Probability values for the ADF t-statistics (reported in square brackets) are obtained from the critical values reported by MacKinnon (1991, Table 1), while those for the PP normalized bias statistics are obtained from the critical values reported by Haug (1992, Table 2). The KPSS test statistics test the null hypothesis of *cointegration* (i.e. H_0: u - ΦS is I(0)) against the alternative hypothesis of *non-cointegration* (i.e. H_1: u - ΦS is I(1)). Probability values (reported in square brackets) for the KPSS statistics are obtained from the critical values reported by Shin (1992, Table 1).

Table 5

Sensitivity Analysis of the Single-Equation Estimates

Extending the Sample Period

(Sample: 1955Q1-94Q4, 160 Observations)

	Estimates of Long-Run Parameters		Cointegration Tests[b]	
Line	$\Phi S_t = \Phi_3 union + \Phi_4 paytax$	$\hat{\tau}_\mu$	$Z(\hat{\alpha})$	$\hat{\eta}_\mu$
EG	.306union + .509paytax	-3.62 [.10]	-27.41 [.05]	0.007 [>.10]
ECM	.490union + .290paytax (2.14) (1.33)	-3.75 [.05]	-26.53 [.05]	0.014 [>.10]
PL	.449union + .346paytax (1.90) (1.51)	-3.75 [.05]	-27.14 [.05]	0.010 [>.10]
FM	.401union + .434paytax (2.08) (2.46)	-3.68 [.10]	-27.48 [.05]	0.007 [>.10]
SW	.287union + .539paytax (3.08) (6.71)	-3.57 [.10]	-27.17 [.05]	0.008 [>.10]

b. See notes to Table 4.

Table 6

Sensitivity Analysis of the Single-Equation Estimates

Stability Tests

(Sample: 1955Q1-94Q4, 160 Observations)

Estimates of Long-Run Parameters[a]			*Stability Tests*[b]		
Line	***(ΦS_t)***		***L_c***	***MeanF***	***SupF***
1	.401union	+ .434paytax	0.15	3.05	6.03
	(2.08)	(2.46)	[>.20]	[>.20]	[>.20]
2	.88union		0.31	4.12	8.19
	(5.94)		[>.20]	[.07]	[>.20]
3	.758paytax		0.10	1.46	2.89
	(6.26)		[>.20]	[>.20]	[>.20]

a. The estimates of the long-run parameters reported above are obtained using the FM procedure.

b. The L_c, MeanF and SupF tests examine the null hypothesis of a stable long-run relationship among I(1) variables against different alternative hypotheses (see text). A rejection of the null hypothesis provides evidence of parameter instability. Probability values are reported in square brackets.

Table 7

Sensitivity Analysis of the Single-Equation Estimates

4th-Order Dynamic Specification

(Sample: 1955Q1-94Q4, 160 Observations)

Estimates of Long-Run Parameters			*Cointegration Tests*[b]		
Line	***$\Phi S_t = \Phi_3 union + \Phi_4 paytax$***		$\hat{\tau}_\mu$	***$Z(\hat{\alpha})$***	$\hat{\eta}_\mu$
ECM	.676union	+ .110paytax	-3.67	-24.11	0.022
	(2.47)	(0.41)	[.10]	[.10]	[>.10]
PL	.584union	+ .215paytax	-3.73	-25.76	0.015
	(2.12)	(0.78)	[.10]	[.05]	[>.10]
SW	.283union	+ .535paytax	-3.59	-27.24	0.007
	(2.52)	(5.01)	[.10]	[.05]	[>.10]

b. See notes to Table 4.

Table 8

Sensitivity Analysis of the Single-Equation Estimates

Adding Cyclical Variables

Two-Step Engle-Granger Procedure

(Sample: 1955Q1-94Q4, 160 Observations)

	1st Step - Estimates of the Long-Run Parameters with the Engle-Granger Procedure	***Cointegration Tests[b] of the Hypothesized Cointegration Vector***		
Line	$\Phi S_t = \Phi_3 union + \Phi_4 paytax$	$\hat{\tau}_\mu$	$Z(\hat{\alpha})$	$\hat{\eta}_\mu$
1	.306union + .509paytax	-3.62 [.10]	-27.41 [.05]	0.007 [>.10]

	2nd Step - Estimation of Error-Correction Models $C(L)\Delta u_t = D(L)\Delta S_t + E(L)Z_t + \gamma[u_{t-1} - \Phi S_{t-1}]$	***Error-Correction Term (γ)***
2	*"Benchmark" Specification - Include ΔS_t Only*	-0.075 (3.61)
3	Z_t *= Output Gap (Potential Output: HP: θ = 1600)*	-0.064 (2.42)
4	Z_t *= Real Wage Equilibrium -- Level Condition*	-0.070 (3.21)
5	Z_t *= Commodity and Energy Prices*	-0.075 (3.32)
6	Z_t *= Term Structure of Interest Rates (rs - rl)*	-0.050 (2.15)

b. See notes to Table 4.

Table 9

Sensitivity Analysis of the Single-Equation Estimates

Adding Cyclical Variables

Joint Estimation of Long-Run Factors and Cyclical Variables

(Sample: 1955Q1-94Q4, 160 Observations)

	Estimates of Long-Run Parameters[a]	***Cointegration Tests***[b]			
Line	$\Phi S_t = \Phi_3 union + \Phi_4 paytax$	$\hat{\tau}_\mu$	$Z(\hat{\alpha})$	$\hat{\eta}_\mu$	$RBAR^2$
	"Benchmark" Specification - Long-Run Factors Only				
1	.449union + .346paytax	-3.75	-27.14	0.010	0.9866
	(1.90) (1.51)	[.05]	[.05]	[>.10]	
	Include Output Gap (Potential Output: HP: θ = 1600)				
2	.710union + .048paytax	-3.60	-22.72	0.032	0.9906
	(3.64) (0.24)	[.10]	[.10]	[>.10]	
	Include Output Gap (Potential Output: HP: θ = 500)				
3	.720union + .037paytax	-3.59	-22.56	0.032	0.9903
	(3.29) (0.17)	[.10]	[.10]	[>.10]	
	Include Output Gap (Potential Output: HP: θ = 10 000)				
4	.729union + .046paytax	-3.61	-22.95	0.028	0.9905
	(3.66) (0.23)	[.10]	[.10]	[>.10]	
	Include Real Wage Equilibrium -- Level Condition				
5	.655union + .206paytax	-3.71	-25.78	0.011	0.9867
	(1.73) (0.63)	[.10]	[.05]	[>.10]	
	Include Commodity and Energy Prices				
6	.382union + .398paytax	-3.73	-27.35	0.010	0.9871
	(1.78) (1.91)	[.10]	[.05]	[>.10]	
	Include Term Structure of Interest Rates (rs - rl)				
7	.581union + .212paytax	-3.73	-25.68	0.016	0.9880
	(2.07) (0.77)	[.10]	[.05]	[>.10]	
	Include Statistically Significant Cyclical Variables				
8	.622union + .152paytax	-3.69	-24.61	0.022	0.9909
	(5.56) (1.67)	[.10]	[.10]	[>.10]	

a. The estimates reported above are obtained using the PL procedure.
b. See notes to Table 4.

Figure 1
The Unemployment Rate
(The Number of Unemployed as a Percentage of the Labour Force)

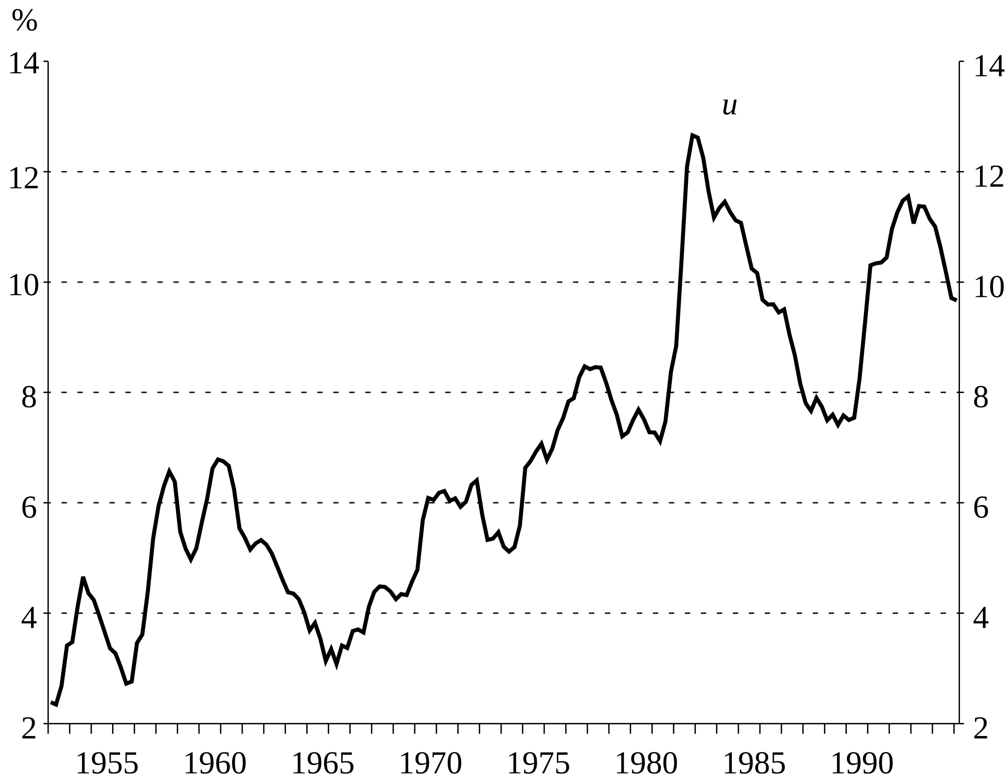


Figure 2
The UI Replacement Rate
(Adjusted for Coverage)

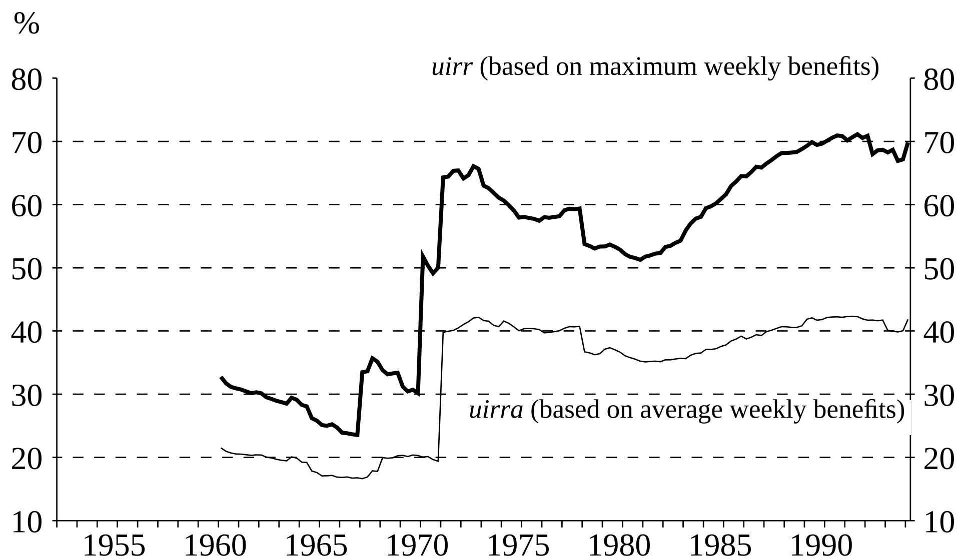


Figure 3
Sargent's Index of UI Generosity - Flexible Time Horizon
1970=100

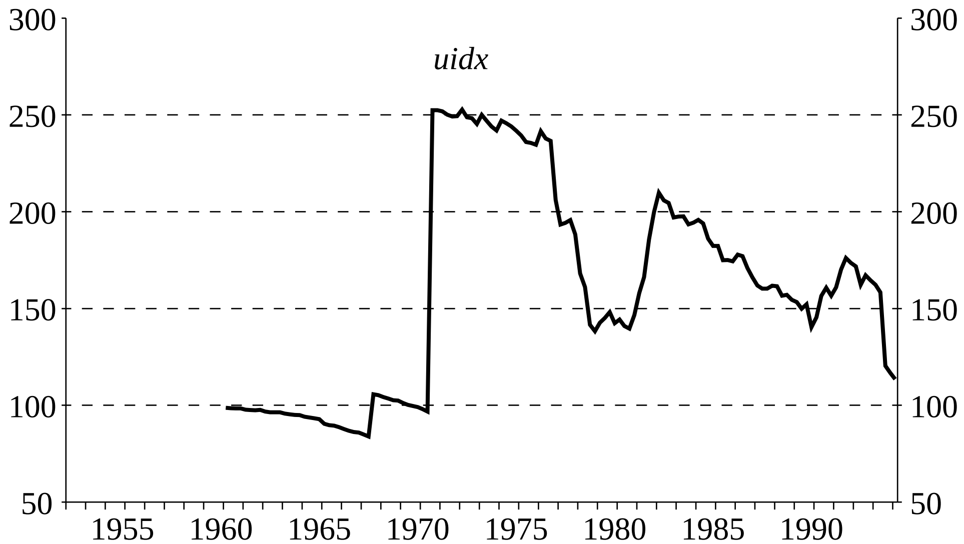


Figure 4
The Percentage of the Labour Force that is Unionized

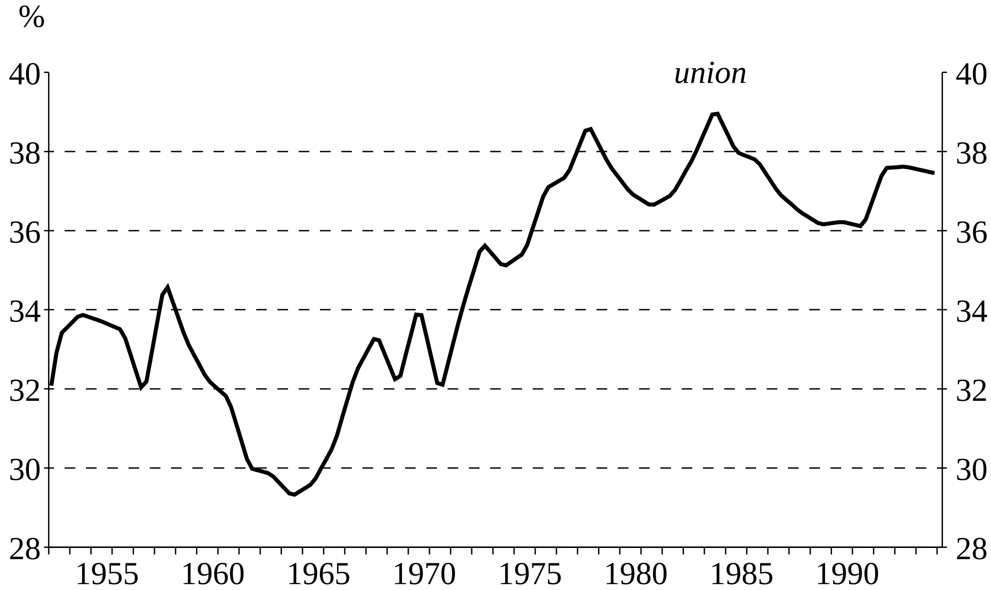


Figure 5
The Minimum Wage Rate Relative to Average Hourly Earnings

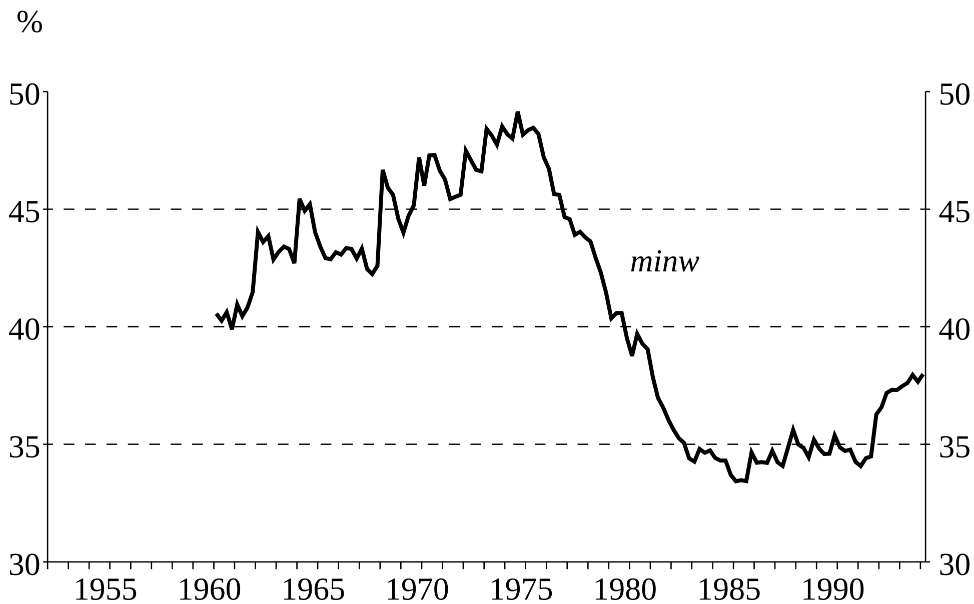


Figure 6
Payroll Taxes Relative to Wages and Salaries

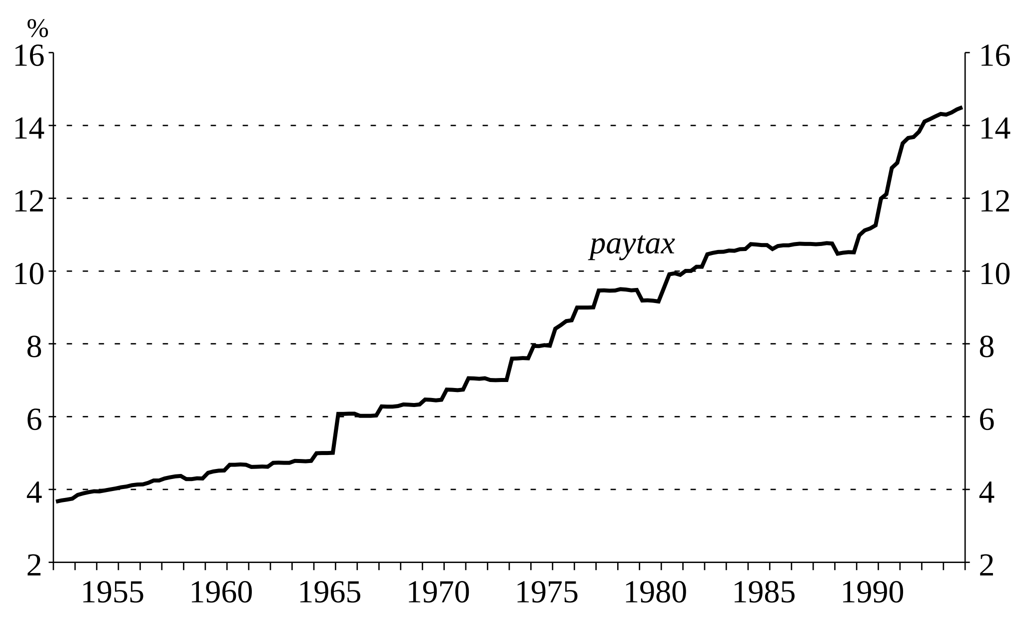


Figure 7
The Percentage of Adult Women and Youths in the Labour Force

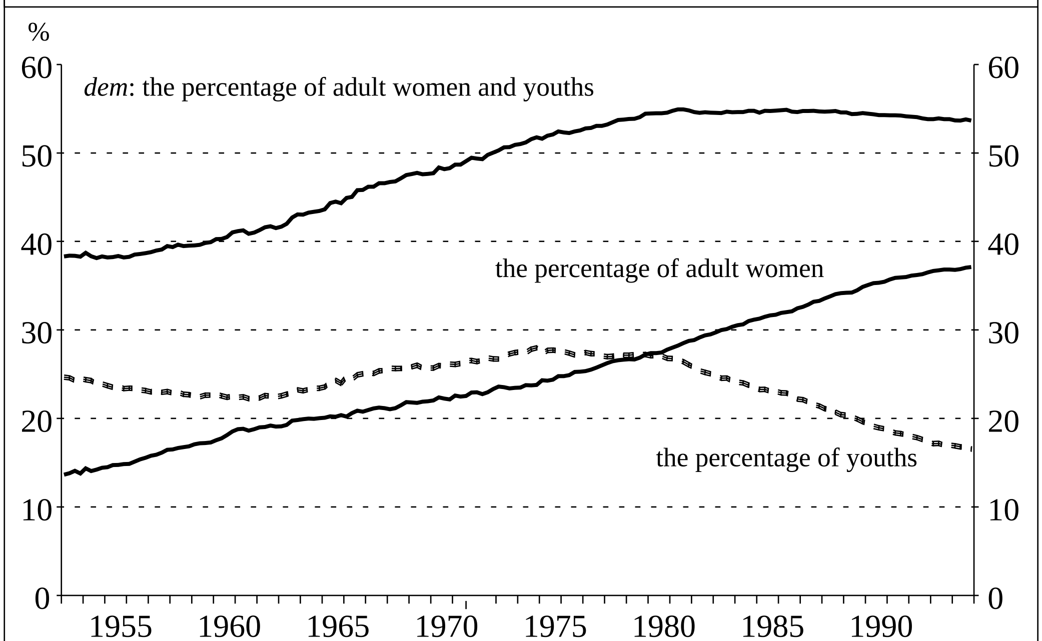


Figure 8
Nominal Wage Growth and Inflation
(Quarterly growth rates)

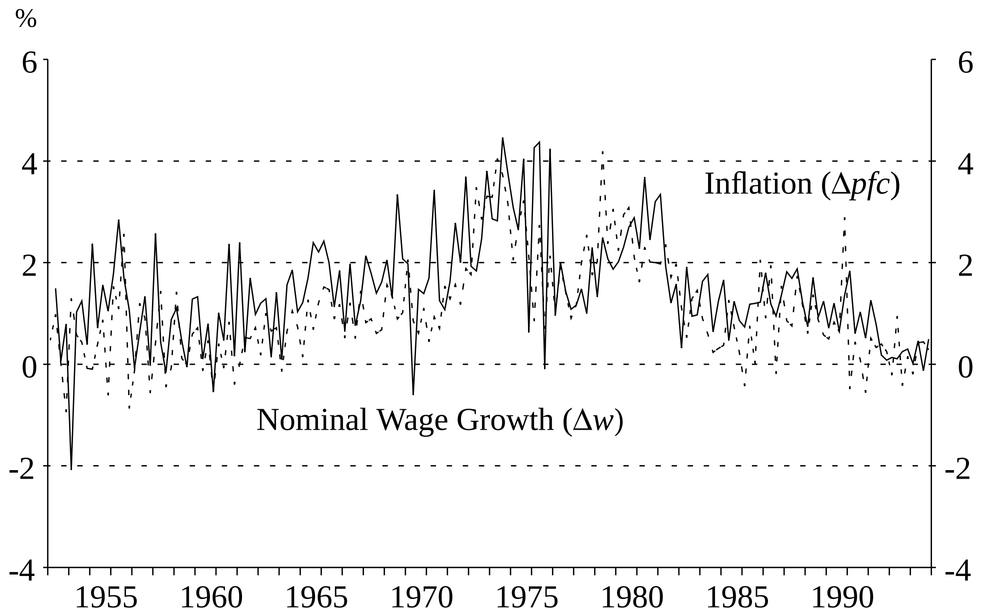
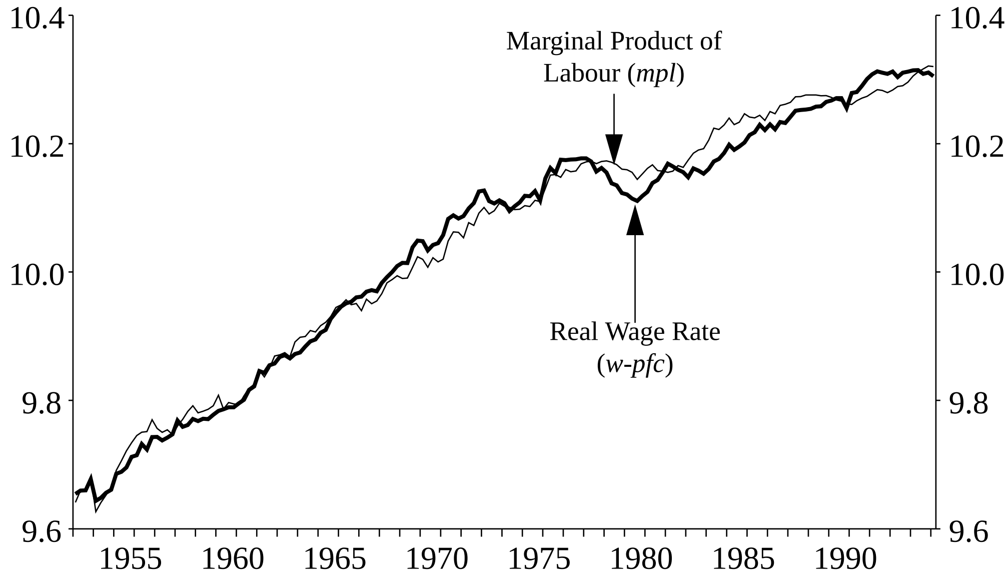


Figure 9
The Real Wage Rate and the Marginal Product of Labour
(Shown in log form)



The Deviation of the Real Producer Wage (w-pfc) from the Marginal Product of Labour (mpl)

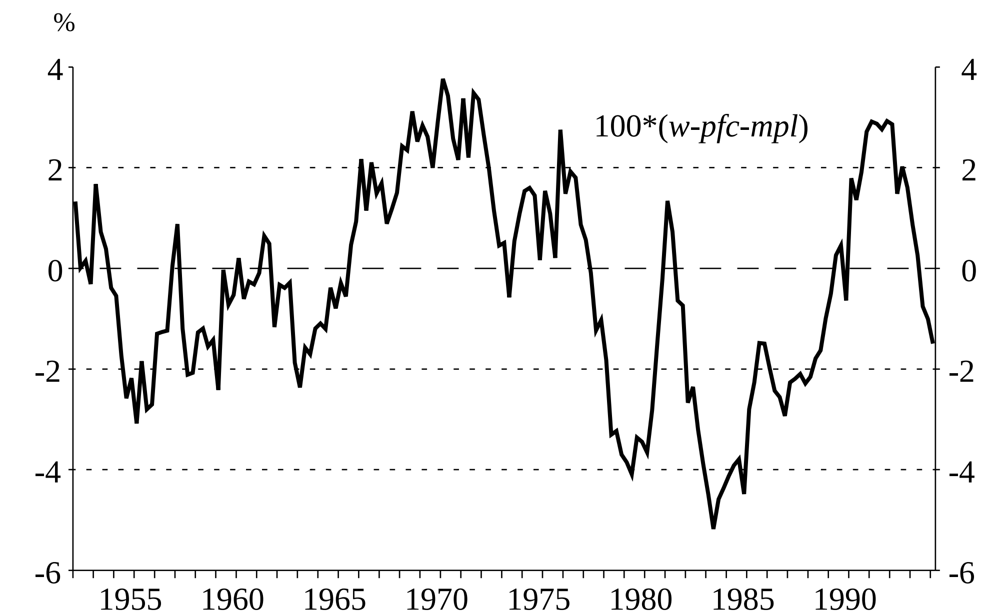


Figure 10
The Deviation of Real Output from its Trend
(Trend generated by the HP filter with the "smoothness parameter" set to 1600)

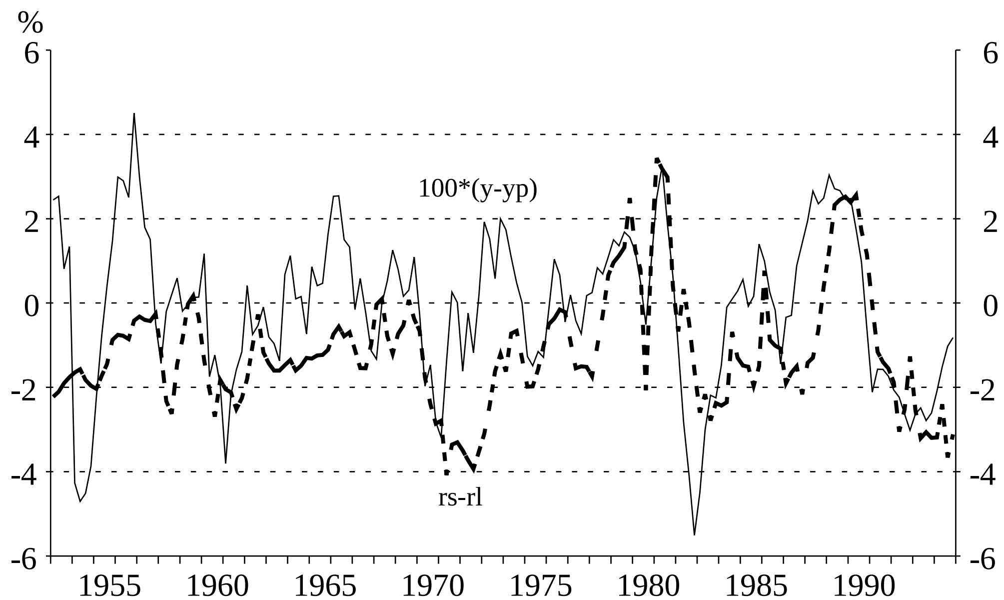


Figure 11
Trend Unemployment Rate
(Our best measure of trend unemployment based on estimates reported in line 8 of Table 9)

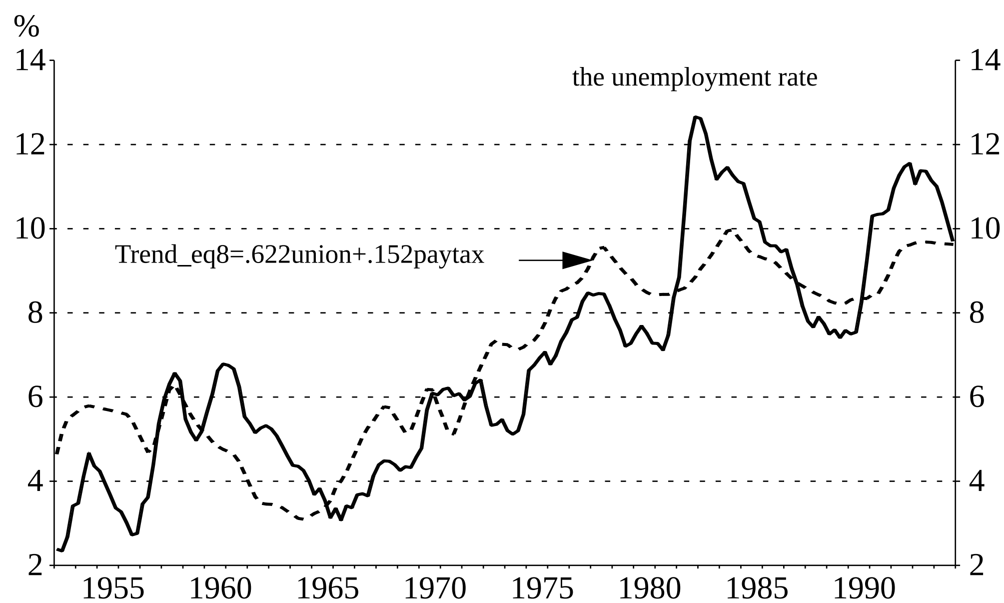
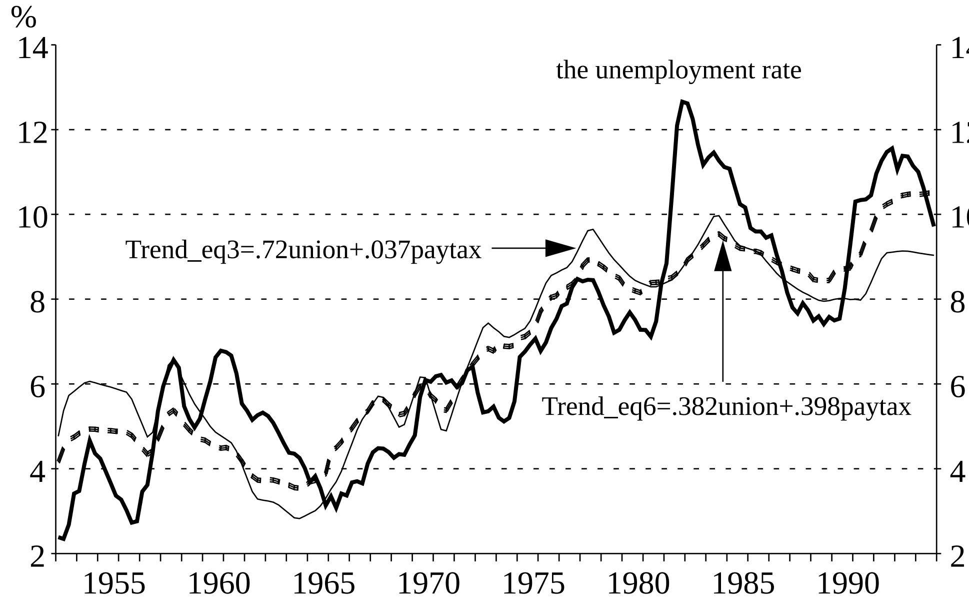


Figure 12
The Effect of Fragility

(Alternative measures of trend unemployment rate based on estimates reported in Table 9)



BIBLIOGRAPHY

Adams C. and D.T. Coe (1990), "A Systems Approach to Estimating the Natural Rate of Unemployment and Potential Output for the United States", *IMF Staff Papers* Vol. 37 No. 2 (June), 232-293.

Akaike H. (1969), "Statistical Predictor Identification", *Annals of the Institute of Statistical Mathematics* 21, 203-207.

Amano R.A. and S. van Norden (1992), "Unit Root Tests and the Burden of the Proof", Bank of Canada Working Paper 92-7, Ottawa.

Andrews D.W.K. and J.C. Monahan (1992), "An Improved Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimator", *Econometrica* Vol. 60, 953-966.

Armstrong J., R. Black, D. Laxton and D. Rose (1995), *A Robust Method for Simulating Forward-Looking Models*, The Bank of Canada's New Quarterly Projection Model, Part 2, Technical Report No. 73. Ottawa: Bank of Canada.

Ashenfelter O. and D. Card (1986), "Why Have Unemployment Rates in Canada and the United States Diverged?", *Economica* 53 Supplement (210), S171-196.

Banerjee A., J.J. Dolado and R. Mestre (1993), "On Some Simple Tests for Cointegration: the Costs of Simplicity", Bank of Spain Working Paper No. 9302.

Black R., D. Laxton, D. Rose and R. Tetlow (1994), *The Steady-State Model: SSQPM*, The Bank of Canada's New Quarterly Projection Model, Part 1, Technical Report No. 72. Ottawa: Bank of Canada.

Burns A. (1990), "Unemployment in Canada: Frictional, Structural and Cyclical Aspects", *Economic Council of Canada* Working Paper No. 1.

Butler L. (1996), *A Semi-Structural Method to Estimate Potential Output: Combining Economic Theory with a Time-Series Filter*, The Bank of Canada's New Quarterly Projection Model, Part 4, Forthcoming Bank of Canada Technical Report.

Card, D. and W. C. Riddell (1993), "A Comparative Analysis of Unemployment in Canada and the United States", in *Small Differences that Matter: Labor Markets and Income Maintenance in Canada and the United States*, edited by D. Card and R. Freeman, Chicago: University of Chicago Press.

Coe D.T. (1990), "Structural Determinants of the Natural Rate of Unemployment in Canada", *IMF Staff Papers* Vol. 37 No. 1 (March), 94-115.

Coletti D., B. Hunt, D. Rose and R. Tetlow (1996), *The Dynamic Model: QPM*, The Bank of Canada's New Quarterly Projection Model, Part 3, Forthcoming Bank of Canada Technical Report.

Corak M. (1994), "Unemployment Insurance, Work Disincentives, and the Canadian Labour Market: An Overview", in *Unemplyment Insurance and How to Make it Work*, C.D. Howe Institute, Toronto (Ontario, Canada), 86-159.

Côté D. and D. Hostland (1994), "Measuring Potential Output and the NAIRU as Unobserved Variables in a Systems Framework", in *Economic Behaviour and Policy Choice Under Price Stability*, Ottawa: Bank of Canada, 359-411.

Davidson R. and J. G. MacKinnon (1993), "Estimation and Inference in Econometrics", Oxford University Press (New York, Oxford).

Delong D.N., J.C. Nankervis, N.E. Savin and C.H. Whiteman (1992a), "Integration Versus Trend Stationarity in Time Series", *Econometrica* Vol. 60 No. 2 (March), 423-433.

———————. (1992b), “The Power Problems of Unit Root Tests in Time Series with Autoregressive Errors”, *Journal of Econometrics* Vol. 53 No. 1-3 (July-September), 323-344.

Dickey D.A. and W.A. Fuller (1979), “Distribution of the Estimators for Autoregressive Time Series with a Unit Root”, *Journal of the American Statistical Association* 74, 427-431.

Duguay P. and S. Poloz (1994), “The Role of Economic Projections in Canadian Monetary Policy Formulation”, *Canadian Public Policy,* 20 (2): 189-199.

Engle R.F. and C.W.J. Granger (1987), “Co-Integration and Error Correction: Representation, Estimation and Testing”, *Econometrica* 55, 251-276.

Engle R.F., D.F. Hendry, and J.-F. Richard (1983), “Exogeneity”, *Econometrica* 51 No. 2 (March), 277-304.

Ford R. and D. Rose (1989), “Estimates of the NAIRU Using an Extended Okun’s Law”, *Bank of Canada Working Paper* 89-3 (November).

Fortin P. (1984), “Unemployment Insurance Meets the Classical Labor Supply Model”, *Economic Letters* 14, 275-281.

———————. (1991), “The Phillips Curve, Macroeconomic Policy, and the Welfare of Canadians”, *Canadian Journal of Economics* 20, 774-803.

Fortin P., M. Keil, and J. Symons (1995), “The Sources of Unemployment in Canada, 1967-1991: Evidence from a Panel of Regions and Demographic Groups”, Working paper No. 36, Research Centre on Employment and Economic Fluctuations, Université du Québec à Montréal (March).

Fougère M. (1995), “Pourquoi le Taux de Chômage est-il plus élevé au Canada qu’aux États-Unis?”, Manuscript, Department of Finance, Government of Canada (June). Paper presented to the CEA Meetings, Montréal, June 1995.

Friedman M. (1968), "The Role of Monetary Policy", *American Economic Review* Vol. LVIII No. 1, 1-17.

Fuller W.A. (1976), *Introduction to Statistical Time Series*, (Wiley, New York).

Granger C. W. J. and P. Newbold (1974), "Spurious Regressions in Econometrics", *Journal of Econometrics* Vol. 2, 111-120.

Hansen B.E. (1992), "Tests for Parameter Instability in Regressions with I(1) Processes", *Journal of Business and Economic Statistics* Vol. 10 No. 3 (July), 321-335.

Haug A. A. (1992), "Critical Values for the $\hat{\mathbf{z}}_{\alpha}$ - Phillips-Ouliaris Test for Cointegration", *Oxford Bulletin of Economics and Statistics* Vol. 54 No. 3, 473-480.

Hodrick R. J. and E. C. Prescott (1980), "Post-War U.S. Business Cycles: An Empirical Investigation", Northwestern University Discussion Paper No. 451.

Inder B. (1993), "Estimating Long-run Relationships in Economics", *Journal of Econometrics* Vol. 57, 53-68.

Kremers J., N. Ericsson and J. Dolado (1992), "The Power of Cointegration Tests", *Oxford Bulletin of Economics and Statistics* Vol. 54 No. 3, 325-348.

Kwiatkowski D., P.C.B. Phillips, P. Schmidt and Y. Shin (1992), "Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root: How sure are we that economic time series have a unit root?", *Journal of Econometrics* Vol. 54, 159-178.

Laxton D. and R. Tetlow (1992), "A Simple Multivariate Filter for the Measurement of Potential Output", Technical Report No. 59. Ottawa: Bank of Canada.

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

MacKinnon J.G. (1991), "Critical Values for Cointegration Tests", Chapter 13 in *Long-Run Economic Relationships: Readings in Cointegration*, Engle R. and C. Granger (eds.), Oxford University Press (Oxford).

Milbourne, R.D., D.D. Purvis, and W.D. Scoones (1991), "Unemployment Insurance and Unemployment Dynamics", *Canadian Journal of Economics* XXIV No. 4 (November), 804-826.

Newey W.K. and K.D. West (1987), "A Simple, Positive Semi-Definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix", *Econometrica* Vol. 55 No. 3 (May), 703-708.

Ng S. and P. Perron (1995), "Unit Root Tests in ARMA Models with Data Dependent Methods for the Selection of the Truncation Lag", *Journal of the American Statistical Association* Vol. 90 No. 429 (March), 268-281.

Ogaki M. and J. Y. Park (1990), "A Cointegration Approach to Estimating Preference Parameters", Paper No. 236, Rochester Center for Economic Research.

Paquet A. (1994), "A Guide to Applied Modern Macroeconometrics", Department of Finance Working Paper No. 94-05, Government of Canada.

Phillips P.C.B. (1986), "Understanding Spurious Regressions in Econometrics", *Journal of Econometrics* 33 No. 3, 311-340.

———. (1987), "Time Series Regression with a Unit Root", *Econometrica* 55, 277-302.

———. (1991), "Optimal Inference in Cointegrated Systems", *Econometrica* Vol. 59 No. 2 (March), 283-306.

Phillips P.C.B. and B.E. Hansen (1990), "Statistical Inference in Instrumental Variables Regression with I(1) Processes", *The Review of Economic Studies* Vol. 57 No. 189 (January), 99-126.

Phillips P.C.B. and M. Loretan (1991), "Estimating Long-run Economic Equilibria", *The Review of Economic Studies* Vol. 58 No. 195 (May), 407-436.

Phillips P.C.B. and S. Ouliaris (1990), "Asymptotic Properties of Residual Based Tests for Cointegration", *Econometrica* Vol. 58 No. 1 (January), 165-193.

Phillips P.C.B. and P. Perron (1988), "Testing for a Unit Root in Time Series Regression", *Biometrica* 75, 335-346.

Poloz S. S. (1994), "The Causes of Unemployment in Canada: A Review of the Evidence", Bank of Canada Working Paper No. 94-11.

Poloz S. and G. Wilkinson (1992), "Is Hysteresis a Characteristic of the Canadian Labour Market? A Tale of Two Studies", Bank of Canada Working Paper No. 92-3.

Poloz S., D. Rose and R. Tetlow (1994), "The Bank of Canada's New Quarterly Projection Model (QPM): An Introduction", *Bank of Canada Review* (Autumn): 23-38.

Quah D. (1992), "The Relative Importance of Permanent and Transitory Components: Identification and Theorical Bounds", *Econometrica* Vol. 60, 107-118.

Rose D.E. (1988), "The NAIRU in Canada: Concepts, Determinants and Estimates", Technical Report No. 50. Ottawa: Bank of Canada.

Said S.E. and D.A. Dickey (1984), "Testing for Unit Roots in ARMA(p,q) Models with Unknown p and q", *Biometrica* 71, 599-607.

Saikkonen P. (1991), "Asymptotically Efficient Estimation of Cointegration Regressions", *Econometric Theory* 7, 1-21.

Sargent T. C. (1995), "An Index of the Generosity of Unemployment Insurance", Manuscript, Department of Finance, Government of Canada (May). Paper presented to the CEA Meetings, Montréal, June 1995.

Schmidt P. (1990), "Dickey-Fuller Tests with Drift", *Advances in Econometrics (Vol. 8): Co-Integration, Spurious Regressions and Unit Roots*, T.B. Fomby and G.F. Rhodes (eds.) JAI Press, 161-200.

Schwarz G. (1978), "Estimating the Dimension of a Model", *Annals of Statistics* Vol. 6, 461-464.

Setterfield M.A., D.V. Gordon and L. Osberg (1992), "Searching for a Will o' the Wisp: An Empirical Study of the NAIRU in Canada", *European Economic Review* Vol. 36 No. 1 (January), 119-136.

Shin Y. (1992), "A Residual-Based Test of the Null of Cointegration Against the Alternative of No Cointegration", Manuscript. Department of Economics, Michigan State University, East Lansing.

Solow R. M. (1986), "Unemployment: Getting the Questions Right", *Economica* Vol. 53 , S23-S34.

Stock J.H. (1992), "Unit Roots and Trend Breaks", mimeo (December).

Stock J.H. and M.W. Watson (1993), "A Simple Estimator of Cointegrating Vectors in Higher Order Integrated Systems", *Econometrica* Vol. 61 No. 4 (July), 783-820.

Topel R. (1993), "What Have We Learned from Empirical Studies of Unemployment and Turnover", *American Economic Review: AEA Papers and Proceedings* (May), 110-115.

Van Rijckeghem C. (1993), "Endogeneity in Structural Unemployment Equations: The Case of Canada", IMF Working Paper WP/93/94 (December).

Bank of Canada Working Papers

1996

96-1	Switching Between Chartists and Fundamentalists: A Markov Regime-Switching Approach	R. Vigfusson
96-2	Decomposing U.S. Nominal Interest Rates into Expected Inflation and Ex-Ante Real Interest Rates Using Structural VAR Methodology	P. St-Amant
96-3	Regime-Switching Models: A Guide to the Bank of Canada Gauss Procedures	S. van Norden and R. Vigfusson
96-4	Overnight Rate Innovations as a Measure of Monetary Policy Shocks in Vector Autoregressions	J. Armour, W. Engert and B. S. C. Fung
96-5	A Distant-Early-Warning Model of Inflation Based on M1 Disequilibria	J. Armour, J. Atta-Mensah, W. Engert and S. Hendry
96-6	Provincial Credit Ratings in Canada: An Ordered Probit Analysis	S. Cheung
96-7	An Econometric Examination of the Trend Unemployment Rate in Canada	D. Côté and D. Hostland

1995

Earlier papers not listed here are also available.

95-6	Inflation, Learning and Monetary Policy Regimes in the G-7 Economies	N. Ricketts and D. Rose
95-7	Analytical Derivatives for Markov Switching Models	J. Gable, S. van Norden and R. Vigfusson
95-8	Exchange Rates and Oil Prices	R. A. Amano and S. van Norden
95-9	Selection of the Truncation Lag in Structural VARs (or VECMS) with Long-Run Restrictions	A. DeSerres and A. Guay
95-10	The Canadian Experience with Weighted Monetary Aggregates	D. Longworth and J. Atta-Mensah
95-11	Long-Run Demand for M1	S. Hendry
95-12	The Empirical Performance of Alternative Monetary and Liquidity Aggregates	J. Atta-Mensah

Single copies of Bank of Canada papers may be obtained from
Publications Distribution, Bank of Canada, 234 Wellington Street Ottawa, Ontario K1A 0G9

E-mail: publications@bank-banque-canada.ca
WWW: http://www.bank-banque-canada.ca
FTP: ftp.bank-banque-canada.ca (login: anonymous, to subdirectory /pub/publications/working.papers)