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Persistence of Unemployment in the Canadian Provinces

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Abstract

We analyze the degree of persistence of the unemployment rates of the 10 Canadian provinces using quarterly data for the period 1976:1-2005:4. We apply a two-break minimum Lagrange Multiplier unit root statistic, which, unlike standard unit root statistics (without or with breaks), makes it possible to find the stationarity of the different unemployment rates, giving support to the theory of the natural rate. We use the methodology of Bai and Perron (1998, 2003) to estimate a linear model with multiple structural changes to estimate the different degrees of persistence over the different regimes. The results suggest that the degree of persistence decreases when multiple breaks are allowed. Issues regarding the Canadian labor market, the insurance benefits program, inter-provincial transfers, and inter-provincial mobility are discussed as potential explanations for the results.

Keywords: Persistence, Unemployment, Unit Root, Break Dates, Canadian Provinces.

JEL Classification: C22, C52, R11

Resumen

En este documento analizamos el grado de persistencia en las tasas de desempleo de las 10 provincias Canadienses utilizando datos trimestrales para el periodo 1976:1-2005:4. Como parte de la metodología, aplicamos un estadístico de raíz unitaria con dos quiebres calculado como el mínimo de los estadísticos LM estimados para cada observación. A diferencia de estadísticos de raíz unitaria tradicionales (con o sin quiebres), dicho estadístico permite encontrar estacionariedad en las diferentes tasas de desempleo dando respaldo a la teoria de la tasa natural de desempleo. Luego, usamos la metodología propuesta por Bai y Perron (1998, 2003) para estimar un modelo lineal con múltiples cambios estructurales que permite estimar los diferentes grados de persistencia en los diferentes regímenes seleccionados. Los resultados sugieren que el grado de persistencia disminuye cuando múltiples quiebres estructurales son permitidos y modelados. Aspectos relacionados con la estructura del mercado laboral Canadiense, el sistema y programa de beneficios sociales, las transferencias inter-provinciales, y la mobilidad inter-provincial son discutidos como explicaciones potenciales de los resultados obtenidos.

Palabras Claves: Persistencia, Desempleo, Raíz Unitaria, Quiebres Estructurales, Provincias Canadienses

Classificación JEL: C22, C52, R11

Persistence of Unemployment in the Canadian Provinces[1]

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1 Introduction

A significant volume of research is devoted to analyzing the behavior of the unemployment rate. One of the main theories in macroeconomics is the theory of the natural rate of unemployment proposed by Phelps (1967) and Friedman (1968). According to this theory, the unemployment rate is stationary around a natural rate and shocks have transitory effects on unemployment.

The validity of this theory, however, has been questioned as a consequence of the behavior of the unemployment rates in the 1970s and the 1980s which were characterized by high degrees of persistence for a long period of time. When the effects of shocks are permanent, we have a phenomenon called hysteresis (see Blanchard and Summers 1987). This effect can be explained by the fact that individuals who have been unemployed for long periods, are likely to remain unemployed, since their skills eventually decline. This means that the past unemployment affects current and future unemployment rates.

In the empirical literature, Blanchard and Summers (1986) estimated an AR(1) model for the unemployment rates in the United States and the United Kingdom. Based on the estimated value of the autoregressive coefficient, they concluded that the unemployment rates in these two countries are highly persistent with a weak tendency to return to their mean. From a more analytical perspective, Milbourne, Purvis and Scoones (1991) built a labor market model to investigate the effect of the Canadian unemploy-

[1] This paper is drawn from the third chapter of the PhD Dissertation of Firouz Fallahi when Gabriel Rodríguez was Associate Professor at the Department of Economics of the University of Ottawa. We thank Gamal Atallah, Lynda Khalaf, Serge Coulombe, Marcel Voia, and Greg Tkacz for constructive comments. We also acknowledge the interesting comments of the Editor and two anonymous referees. Useful conversations with Jean-Fançois Tremblay from the University of Ottawa are also acknowledged.

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ment insurance system on unemployment. Using this model and monthly unemployment data for Canada for the period 1970-1988, they showed that features of the Canadian unemployment insurance system make the natural rate of unemployment dependent on its past values, which means there is persistence in the data.

Jaeger and Parkinson (1994) used data on unemployment rates for Canada, Germany, the United Kingdom, and the United States to test for the presence of hysteresis. Using standard unit root statistics and an unobserved component model, they concluded that hysteresis was present in the data. An opposite result, however, was found by Jones (1995). Using an expanded data set, Røed (1996) examined the unemployment rates in 16 OECD countries including Canada and the results, obtained using unit root statistics, confirmed the existence of hysteresis in the unemployment rates of Canada, Australia, Japan, and several European countries over the period 1970-1994. Using a similar sample of countries, Bianchi and Zoega (1998) applied Markov-Switching models to study the persistence and found no evidence of persistence in unemployment rates. In the case of Canada, they determined the existence of three different regimes for the period 1970-1996.

From another perspective, Keil and Pantuosco (1998) used a panel model to address the persistence of the unemployment rates in Canadian and US regions. With the coefficient of the lagged dependent variable in the panel data model as an indicator of persistence, they found some evidence of persistence in Canada. Their results show that this coefficient varies between 0.73 and 0.92, depending on the specification of the model.

Using the unemployment rates of 19 OECD countries for the period 1956-2001, Camarero and Tamarit (2004) applied multivariate SURE unit root statistics, which indicated there was no evidence of hysteresis in most of these countries, including Canada. In a more recent paper, Camarero, Carrión-i-Silvestre, and Tamarit (2006) used a stationary panel test with breaks to test for hysteresis in unemployment rates in OECD countries. Based on data for the period 1956-2001, exploiting the cross-section variations of the series, and additionally, allowing for a different number of endogenous breakpoints in the unemployment series, their findings stress the importance of accounting for exogenous shocks in the series and give support to the natural-rate hypothesis of unemployment for the majority of the countries analyzed.

On the other hand, Mikhail, Eberwein and Handa (2005) investigated the persistence in aggregate and sectorial Canadian unemployment for the period 1976-1999. Their results provided strong evidence of high persistence and they concluded that the fluctuations in aggregate and sectorial Canadian unemployment are persistent. Recently, Mikhail, Eberwein and Handa (2006) reinvestigated persistence in the aggregate Canadian unemployment using Bayesian ARFIMA models. Using data for the period 1976-1998, they found that persistence in the unemployment rate of Canada holds in the short and medium run. In the long run, however, persistence is uncertain.

In a recent study, Gustavsson and Osterholm (2006) used a new unit root statistic suggested by Kapetanios, Shin, and Snell (2003) to investigate the presence of hysteresis in monthly unemployment data for Australia, Canada, Finland, Sweden, and the United States. They concluded that by allowing for nonlinearity in the statistics, much less evidence of hysteresis may be found in the data in comparison with standard unit root statistics. They rejected the presence of hysteresis in all countries, except Australia.

From our observation of the previous literature, we find that persistence of Canadian unemployment at the aggregate level has been previously studied. To the best of our knowledge, however, this paper is the first attempt to study the degree of persistence in the unemployment rates of the Canadian provinces using novel procedures. In this paper, we investigate the degree of persistence in the unemployment rates of the 10 Canadian provinces. We also include the national aggregate. One aspect to be investigated is the stationarity or the nonstationarity of the unemployment rates. To carry out this investigation, we use recent unit root statistics that are more powerful than standard unit root statistics. We also apply unit root statistics that allow for one or two breaks. According to the results, another issue deserving study is the calculation of the degree of persistence of unemployment rates in the Canadian provinces. In economic terms, the central theme of the paper is to show how labor markets in Canada adjust following an economic shock.

Our conjecture is that unemployment rates are stationary after allowing for structural breaks. This means that unemployment rates are mean reverting where the mean moves between regimes. Furthermore, we believe that allowing for structural breaks makes it possible for us to obtain smaller levels of persistence in comparison with the linear case. Therefore, the speed of adjustment is faster when we allow for structural breaks.

The application of recent unit root statistics (without breaks) indicates that unemployment rates are mostly nonstationary with a few exceptions. The application of augmented Dickey-Fuller (ADF)-type unit root statistics that allow for one and two structural breaks again produces mixed results. One drawback of ADF-type tests, however, is that their size properties are affected by breaks under the null. To overcome this weakness, we use a two-break minimum Lagrange Multiplier (LM) statistic. In this case the null hypothesis is rejected in favor of stationarity with two breaks for all provinces, showing support to the hypothesis of natural rate of unemployment. Furthermore, we complement the results using a panel unit root statistic with two breaks. The results give strong confirmation of stationarity of the unemployment rates.

When we use the approach of Bai and Perron (1998, 2003) which entails estimating a linear model with multiple breaks, the estimated break dates (and their intervals of confidence) are very precise. For most provinces the first break date is around 1981-1982 and the second break dates are around 1989-1991. Ontario and Nova Scotia are the only provinces that have a third break date and these happen in 2000:4 and 1993:3, respectively. Seven breaks occur during 1981-1982 (the recession period), five breaks between 1988 and 1990, two breaks in 1993, and two breaks in 2000, which show some sort of clustering.

Two points of interest can be deduced from the results. The first point is the fact that the speed of adjustment in provincial markets differs noticeably across provinces. This suggests that something affects labor markets differentially across the country. This issue is discussed when we analyze the results in Section 3. The second point is the fact that the speed of adjustment is quite slow when simple linear autoregressive models are estimated. In these cases the speed of adjustment implies that a return to the mean takes many years, which is in contrast with macroeconomic models that often assume markets adjust fairly quickly. However, when we estimate the persistence using the approach of Bai and Perron (1998, 2003), that is, working with stationarity around multiple breaks, values of persistence decrease between regimes, in particular comparing first and last regime. These small values of persistence imply a high speed of adjustment which is in agreement with what most macroeconomic models suggest.

The remainder of this paper is organized as follows. The next section presents and applies different unit root statistics to determine whether the unemployment rates are stationary or nonstationary. Section 3 estimates the levels of persistence for the different provinces using a multiple regression model with the structural changes proposed by Bai and Perron (1998, 2003). Section 4 concludes.

2 Identifying Stationarity or Nonstationarity of Unemployment Rates

We use quarterly unemployment rates for the 10 Canadian provinces for the period 1976:1-2005:4 obtained from Statistics Canada. The 10 provinces are Alberta, British Columbia (BC), Manitoba, New Brunswick (NB), New-Foundland (NF), Nova Scotia (NS), Ontario, Prince Edward Island (PEI), Quebec, and Saskatchewan (SAS). To complete the analysis, we also include the national aggregate. Figure 1 shows the evolution of the 11 time series.

A natural way to examine if a time series is stationary is to use unit root statistics. If the null hypothesis of a unit root cannot be rejected, we say that the time series is highly persistent and any shock to this variable has permanent effects. This result is compatible with the hypothesis of hysteresis. In the opposite case, if a rejection of the unit root hypothesis is obtained, the time series is stationary and shocks have transitory effects. In this case the autoregressive coefficient measuring persistence is less than unity and the theory says that unemployment returns to its mean. It is consistent with the existence of a natural rate of unemployment.

In formal terms, let y_t represent the unemployment rate for $t = 1, 2, ..., T$. We consider that the data-generating process is given by

$$y_t = d_t + u_t, \tag{1}$$

$$u_t = \alpha u_{t-1} + v_t, \tag{2}$$

where $d_t = \psi' z_t$, and z_t includes the deterministic components. The most classic examples consider $z_t = \{1\}$ or $z_t = \{1, t\}$, that is, a model with only an intercept, and a model with an intercept and a linear trend, respectively.

As of late, the most powerful unit root statistic has been the so named ADF^{GLS} statistic proposed by Elliott, Rothenberg, and Stock (1996). The regression to be estimated is

$$\Delta \tilde{y}_t = \alpha_0 \tilde{y}_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta \tilde{y}_{t-i} + \epsilon_t, \tag{3}$$

where $\tilde{y}_t = y_t - \widehat{\psi}' z_t$ and the coefficients $\widehat{\psi}$ have been estimated by generalized least squares (GLS). In other words, $\tilde{y}_t$ is GLS detrended data at some particular point of the alternative hypothesis.[3] The null hypothesis is $\alpha_0 =$

[3] The alternative hypothesis is $H_1 : \alpha = \overline{\alpha} = 1 + \overline{c}/T$. Elliott, Rothenberg and Stock (1996) recommended using $\overline{c} = -7.0$ when $z_t = \{1\}$ and $\overline{c} = -13.5$ when $z_t = \{1, t\}$.

$\alpha - 1 = 0$, Δ is the first difference operator, and k is the number of lags that must be determined in such a way that ϵ_t is serially uncorrelated.

Application of the ADF^{GLS} statistic results in rejections for NB (10%), Ontario (5%), and aggregate Canada (10%).[4] A direct interpretation of these results suggests permanent effects of shocks in eight provinces. Only three cases appear to suggest stationarity in the unemployment rates.

Another way to examine the persistence of a time series is by constructing confidence intervals for the respective autoregressive coefficient. Constructing a confidence interval may be more informative than the unit root statistics. If the confidence interval contains the unity, then the variable is nonstationary and is highly persistent. When this interval does not include unity, the variable is stationary and is less persistent.

There are several ways of building confidence intervals of the autoregressive coefficient. Asymptotic confidence interval is the conventional way. In this approach the lower and the upper bounds of the 90% confidence interval can be calculated as the point estimate of the root ± 1.645 times the standard error. However, this way of calculating the confidence interval is not appropriate when the value of the root is large and near to unity, which might be the case in this study. In addition, this method is not useful in constructing the confidence interval when the time series has a unit root, since the traditional asymptotic theory is discontinuous in this case. This represents a serious issue because most macroeconomic variables have a root close or equal to unity.

New methods have been developed that are robust to the presence of a root close to or on the unit circle. One approach is described and applied in Stock (1991). Recently, Hansen (1999) proposed a grid bootstrapping procedure to build the confidence interval of α with correct first-order asymptotic coverage. This procedure gives a confidence interval that is asymptotically valid and can control for size, even in cases where the root is close to one. Using Monte Carlo simulations, Hansen (1999) found that the grid point bootstrapping way of constructing a confidence interval is asymptotically valid and provides good coverage in finite samples.

The approach of Hansen (1999) is applied to our 11 time series. The point estimates[5] show that the autoregressive coefficients are very close to unity, reflecting very high persistence. The lowest coefficient is obtained for NF (0.885). The confidence interval, however, is so broad that it includes

[4] A complete set of results is available upon request.

[5] *A* complete set of results is available upon request. An alternative procedure is described in Elliott and Stock (2001), but similar results are obtained.

the unity. Alberta shows the highest autoregressive coefficient (0.969) and again the confidence interval includes unity. All these results go in the same sense as those obtained with the application of the unit root statistics. This means that most of the unemployment rates are nonstationary, indicating that shocks have permanent effects.

However, it is difficult to think of shocks that have permanent effects on unemployment even if the labor markets are very rigid. Furthermore, even when provincial economies have different structures of production, it is difficult to think that some provincial unemployment rates have permanent effects and others do not. In general, we expect that the unemployment dynamics would be characterized as a mean-reverting process with different degrees of persistence. The mean or "natural rate" may move over time due to changes in real provincial (macroeconomic) variables or in the institutional environment. Therefore, there could be structural breaks of the steady-state path of a stationary process. In this context, high persistence levels imply slow speeds of adjustment towards the long-run equilibrium level. It is a special case of the hypothesis of a natural rate of unemployment. In fact, the presence of structural breaks could be the explanation to the previous results. In this sense what we expect are different degrees of persistence between regimes.

In fact, a non-rejection of the null hypothesis of a unit root may be due to a misspecification of the set of deterministic components. In a seminal contribution, Perron (1989) showed that the ADF-type test fails to reject the null hypothesis of a unit root when the series is stationary with a broken trend function. The approach suggested by Perron (1989) was extended to an unknown break point by Zivot and Andrews (1992) and Perron (1997). Furthermore, Perron and Rodríguez (2003) extended the GLS detrending approach of Elliott, Rothenberg, and Stock (1996) and Ng and Perron (2001) to the context of one unknown structural change.

The results of the application of the ADF^{GLS} statistic proposed by Perron and Rodríguez (2003) reveal that only PEI rejects the null hypothesis of unit root. Again, as the previous results, most of the series appear to be nonstationary and coherent with the fact that shocks have permanent effects.[6]

Following in the framework of structural change, another plausible alternative is the existence of two breaks under the alternative hypothesis of stationarity. One of the most used statistics is the ADF statistic proposed

[6] Similar results are also found using unit root statistics with GLS detrended data with break only in the intercept as proposed by Rodríguez (2007).

by Lumsdaine and Papell (1997). The break points are selected based on the minimum value of the t statistic for $\widehat{\alpha}$ as proposed by Zivot and Andrews (1992). The application of this unit root statistic allows us to obtain 6 rejections out of 11 series. The rejections are obtained for Alberta, BC, Manitoba, NF, SAS, and Canada. The dates of the breaks are clustered in 1980-1981 and 1990-1991, which are in close concordance with the recession periods of the country.

A drawback of ADF-type tests is that their size properties are affected by the presence of breaks under the null. Furthermore, these statistics suffer divergence under the null hypothesis. To overcome this issue, we use a two-break minimum LM statistic proposed by Lee and Strazicich (2003), which is related to the one-break LM unit root test developed by Amsler and Lee (1995). Unlike conventional unit root statistics, the distribution of the LM statistic is invariant to the break point nuisance parameters. In addition, it does not suffer from bias and spurious rejections in the presence of breaks under null.

We apply the two-break minimum LM statistic for two models. In the first model breaks are present in the intercept (Model 1). In the second model breaks are present in the intercept and the slope of the trend function (Model 2). The results (see Table 1) for Model 2 indicate rejection of the null hypothesis for all provinces and for the aggregate, with the exception of Ontario. However, we may reject the null hypothesis for this province using Model 1.[7]

We use the two-break LM statistic to construct a panel LM statistic as proposed by Im et al. (2002).[8] As is well known, panel unit root statistics are considered more powerful because they add the cross-section dimension and increase the amount of information for each time period.[9] Several features used in the panel unit root statistic are worth noting. First, each province is allowed to have unique fixed effects, different time trend coefficients, and varying persistence parameters. Second, the number of structural breaks varies by province. Third, heterogeneous break points are allowed and de-

[7] The results are robust to the change of lag specification. All results are available upon request.

[8] The panel test statistic uses the optimal LM univariate t-statistic and then standar*d*izes that to compute the t-statistic for the panel.

[9] The task of allowing for structural breaks in the existing panel unit root tests such as proposed by Im et al. (2003) and Levin et al. (2002) would be quite difficult to implement, *because* the fact that the distribution of these panels with structural breaks will critically depend on nuisance parameters indicating their location as noted by Im et al. (2002). Therefore, it would be impossible to control for the numerous possible combinations of heterogeneous structural breaks that might occur when using these panel unit root tests.

termined endogenously for each province. Fourth, time-specific fixed effects are allowed to capture any common structural breaks. Fifth, the number of lags is heterogeneously determined jointly with the breaks.

The result of the panel unit root statistic for Model 2 allowing for time-fixed effects is -17.319, which is significant at 1%. If time-fixed effects are not allowed, the value of the statistic is -21.048, which is again significant at the 1%.[10]

Table 1 also shows the break dates identified by the two-break minimum LM statistic. These dates are clustered in the periods 1981-1982, 1991-1992, and 1988-1987. The two first clusters are the recession times frequently identified by different business cycle studies. These two recession times are selected for the aggregate and most of the provinces. The cluster 1986-1987 was a recession period experienced by some provinces; it was not a national recession.

Another possibility is to use a unit root statistic allowing for more than two breaks in the same spirit of Lumsdaine and Papell (1997). However*, this is* not available in the literature.[11] As we established before, we expect and think that the unemployment rates are characterized as a mean-reverting process with different degrees of persistence, where the mean or natural rate may move over time due to different provincial aspects. Furthermore,

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[10] An important issue to consider in the application of the panel unit root statistics is the frequent assumption of cross-section independence. As an anonymous referee well argues, these tests have some flaws in terms of size distortion and lack of power when correlation *exists* among contemporaneous cross-sectional error terms. One potential solution is firstly to use the statistic proposed by Ng (2006) to assess whether individuals in the panel are cross-section independent or not. If we find cross-sectional dependence, the second step is the application of the statistic proposed by Maddala and Wu (1999), which is a bootstrapped statistic to accommodate general forms of cross-section dependence. Another possibility is to use directly a statistic controlling for the presence of different kinds of cross-section dependence. An example is Camarero et al. (2006). Even when we recognize the importance of all *the* above mentioned issues (cross-section independence is hardly found in practice), we believe that they are beyond the methodological scope of the paper*, which* is based on a time series approach. Furthermore, we want to keep the econometric stuff at the minimum level as possible. The panel unit root statistic used in this paper is used only as an approximation to confirm the results obtained by the two-break minimum LM statistic of Lee and Strazicich (2003). We thank an anonymous referee for advice concerning panel unit root tests allowing for different kinds of cross-section dependence.

[11] What is available in the literature is a panel stationarity test allowing for more than two breaks; see Harris, Leybourne and McCabe (2005). As we mentioned before, *however,* this paper is based on a time-series approach, and we want to keep econometric aspects at a minimum needed level. We prefer do not merge different approaches. However, we recognize that the approach based on *a* panel with more than two breaks is interesting to improve the paper in future research. We thank an anonymous referee for this remark.

we think that more than two breaks are possible. Therefore, we use the approach of Bai and Perron (1998, 2003) in the next section.

3 Measuring Persistence in Unemployment Rates

Bai and Perron (1998, 2003) have proposed a method to estimate the coefficients in a multiple linear model with endogenous structural changes when the time series is stationary. The method allows for testing for the number of breaks using different statistics, if the null hypothesis of no structural change is rejected, the method allows for the estimation of the parameters for each regime. Therefore, in the context of an autoregressive model, the approach may allow the degree of persistence for each regime to be estimated.

Following Bai and Perron (1998, 2003) and using a slightly modified notation, we consider the following multiple linear regression with m breaks:

$$\left\{ \begin{array}{cc} y_t = \mu_1 + \beta_1 t + \sum_{i=1}^{k} \rho_{1i} y_{t-i} + \epsilon_t, & t = 1, 2, ..., T_1 \\ y_t = \mu_2 + \beta_2 t + \sum_{i=1}^{k} \rho_{2i} y_{t-i} + \epsilon_t, & t = T_1 + 1, ..., T_2 \\ . & \\ . & \\ . & \\ y_t = \mu_m + \beta_m t + \sum_{i=1}^{k} \rho_{mi} y_{t-i} + \epsilon_t, & t = T_m + 1, ..., T \end{array} \right\} \quad (4)$$

where y_t is the dependent variable (unemployment rate), μ_i is the intercept, β_i is the slope, and ρ_i are the autoregressive coefficients. We allow all coefficients to change according to the number of regimes. The indices (T_1, $T_2, ..., T_m$) are the break points and they are treated as unknown.

To estimate the number of breaks, Bai and Perron (1998, 2003) propose different statistics. The first statistic is the $\sup F$-type test for no break ($m = 0$), against an alternative hypothesis of h breaks ($m = h$). The critical values can be found in Bai and Perron (2003). The maximum number of breaks depends on a trimming parameter. For example for a 15% trimming, the maximum number of allowed breaks is $five$.[12]

Bai and Perron (1998, 2003) propose two other statistics, called double maximum statistics, which can be used to test the null hypothesis of no break against an unknown number of breaks, given some upper limit for the number of breaks, M. The statistics are the $UD_{\max}F_T(M, q)$ and the $WD_{\max}F_T(M, q)$. The critical values can be found in Bai and Perron (2003).

[12] These values are used in the estimations.

Another statistic to test for a given number of breaks l, against $l+1$ breaks is the sup $F_T(l+1|l)$. A model with $l+1$ breaks will be selected if the overall minimal value of the square sum of residuals is sufficiently smaller than the square sum of residuals from a model with l breaks. We can repeat this test for $l=1$ until $l=M$.

Additionally, the number of breaks can be determined using information criteria such as the Bayesian Information Criterion (BIC) and the criterion proposed by Liu, Wu, and Zidek (1997), denoted by LWZ. The number of breaks that minimizes these information criteria is the selected number of breaks for that time series. Finally, repartition and sequential procedures can also be used to determine the exact number of breaks. See Bai and Perron (1998, 2003) for further details.

Table 2 shows the results allowing up to five breaks. The $UD_{\max}$ and the $WD_{\max}$ suggest the existence of at least one break for all provinces and the national aggregate. A similar conclusion is obtained using the sup $F_T(h)$ for $h=1,2,3,4,5$. The rejection of the null hypothesis of no break is obtained at high levels of significance, which supports strong rejection. The sup $F_T(l+1|l)$ also indicates structural change for most of the unemployment rates.

As for the information criteria, the LWZ does not find any break in the unemployment rates. The BIC suggests structural change for Alberta, BC, NS, PEI, SAS, and Canada. In most cases one break is identified. Three breaks are selected for NS and two breaks for Canada.[13]

The last row of Table 2 shows the number of breaks selected by the sequential procedure at different levels of significance. In the case of Manitoba, for example, no breaks are selected using critical values at 1% of significance. However, one break is identified at 2.5% and 5% of significance. Finally, four breaks are selected using critical value at 10% of significance. Therefore, one or more breaks are selected for all unemployment rates except for Quebec. The case of Quebec is interesting because the $UD_{\max}$, the $WD_{\max}$, the sup $F_T(k)$, and the sup $F_T(l+1|l)$ suggest that this province has experienced at least one structural break. Furthermore, sup $F_T(l+1|l)$ selects two breaks for this province. Therefore, we decide to estimate a model with two breaks for this province.

Therefore, to determine the number of breaks, we use the sequential procedure at 5%. According to the results, there are three break points in the unemployment rates of NS and Ontario; one break point in the unemploy-

[13] Notice that many studies have found three regimes for Canada; see Bianchi and Zoega (1998) for example.

ment rates of Alberta, BC, Manitoba, NF, PEI, and SAS; and two breaks in Canada and NB. As mentioned before, we estimate a model with two breaks for Quebec.

Table 3 presents estimates of the intercept, slope, and persistence. Several comments may be made from the estimates of the degrees of persistence. In general, the degrees of persistence are not stable, suggesting that the process is mean reverting but this mean or "natural rate" is moving over time (due to changes in real macroeconomic variables or in institutional environment).[14] In most provinces the value of the persistence decreases when the first and the last regimes are compared. In the case of BC the persistence decreases from an extremely high value of 1.157 to 0.895. The level of persistence in Manitoba decreases as well, from 0.886 to 0.764. In NB, persistence moves from 0.698 to 0.799 and then to almost no persistence in last regime. The persistence in NF decreases from 0.856 to 0.207. The estimated coefficients for Ontario show that the persistence goes up first and then declines to 0.442 at regime 4. The results for PEI show a level of persistence of around 0.50 for all regimes. On the opposite side, the degree of persistence in Alberta, NS, Quebec, SAS, and Canada is larger at the last regime. SAS also shows an increase in the level of persistence. The province of Quebec shows a degree of persistence of 0.799 in the first regime. The persistence at the second and the third regimes are 0.352 and 0.920, respectively. For aggregate Canada, the level of persistence passes from 0.708 in the first regimen to 0.841 in the third regime, which is not a particularly high degree of persistence.

Overall, the degree of persistence in the last regime is less than 0.9 except for Quebec where it is 0.920. Values of persistence for the Atlantic provinces are the smallest. The message is not surprising. If we allow for structural breaks in the trend function, the degree of persistence is reduced, which is in contrast with the results obtained from linear AR(k) models and the results obtained in Section 2 using different unit roots and confidence intervals for unit roots. Our results are in agreement with results found by Keil and Pantuosco (1998), and Gustavsson and Osterholm (2006) where introduction of nonlinearities implies reductions of levels of persistence in the unemployment rates. In addition, the speed of adjustment in provincial labor markets differs noticeably across provinces, which means the response of these markets to the common shocks would be quite different.[15]

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[14] In other words, it means that we have structural breaks in the steady-state path of a stochastic variance stationary time series (which are the unemployment rates).

[15] Panigo et al. (2004) find similar results for the case of Argentina. They state that when shocks to the unemployment rates are persistent and the trend of the unemployment rates

Table 4 presents the estimated break dates (and their intervals of confidence). The dates are very precise and for most provinces, the first break date is around 1981-1982, the second break date around 1989-1991. Ontario and NS are the only provinces that have a third break date, which happen in 2000:4 and 1993:3, respectively. Seven breaks occur during 1981-1982 (the recession period), five breaks between 1988 and 1990, two breaks in 1993, and two breaks in 2000, revealing some sort of clustering.

Many factors may influence the values of the persistence. The Canadian labor market has been characterized by high rigidity. Inter-provincial mobility is probably an important adjustment mechanism, a factor that would limit emigration from provinces with high unemployment rates and would probably tend to modify the persistence of unemployment. Factors that would tend to increase the average duration of unemployment spells in regions with high unemployment also have a similar effect. We may think of two important factors. The first is the inter-provincial redistribution, which has been implemented through the equalization programs as well as through federal transfers to provinces (which are financed from federal tax revenues collected in all provinces). Net redistribution towards any given province will tend to be higher as the tax capacity of the province declines. Thus, net redistribution will tend to increase as GDP goes down and unemployment goes up. That increases the net fiscal benefits in provinces with high unemployment compared with provinces that have low unemployment. There is much less inter-regional redistribution of that sort in the United States, and the inter-regional mobility is significantly greater, which acts as an important adjustment mechanism in response to shocks that hit different regions asymmetrically.

The other important element pertains to unemployment insurance benefits. Normally, such benefits are more generous (and the eligibility condition is less stringent) in regions with high unemployment. For example, an unemployed individual living in a region where the unemployment is high could get more weeks of benefits. Again this reduces significantly the incentives of unemployed workers to move towards provinces with lower unemployment rates. It also tends to increase the average duration of unemployment spells in regions with high incidence of unemployment. Both of these factors tend to increase the overall persistence of unemployment. We think they partly explain the high levels of persistence found in some provinces like Quebec.

changes, policy makers should follow policies that are capable of changing the structure of the labor market. However, if these shocks are persistent but they do not cause changes in the trend of the unemployment rates, then *a* counter-cyclical policy toward the labor market would be enough to help this market adjust to its trend.

In addition, it is worth noting that an important reform of the unemployment insurance system in 1996 made the system much less generous and the eligibility condition much more stringent, and reduced the inter-regional variations in conditions and benefits.

Furthermore, different industry structures and industry performances could create diverse adjustment speeds in the labor markets of the provinces. For example, shocks to the manufacturing industry would affect the unemployment rate in Ontario much more than in Alberta, simply because the manufacturing industry is centered in Ontario. On the contrary, oil shocks would have a significant effect in Alberta. Government policies could also cause the different degrees of persistence and adjustment speeds, because the provincial governments might adopt different policies toward the labor markets. Take for example, the attitude towards the role of migration. Alberta has historically attracted labor when it does well and shed labor when it has done poorly, more than have other provinces.[16]

4 Conclusions

We analyzed the persistence of unemployment rates in the 10 Canadian provinces and the national aggregate. The application of standard unit root statistics suggests no rejection of the null hypothesis of a unit root, which means the unemployment rate does not revert to its mean. Application of the ADF statistic allowing for one or two breaks indicates a few rejections of the null hypothesis. Unlike these statistics, we applied a two-break minimum LM statistic that is invariant to the location of the break points. The application of this statistic suggests strong rejection of the null hypothesis of a unit root, implying that unemployment rates are mean reverting. The application of a panel unit root statistic strongly confirms these results.

Given stationarity, we estimated the degree of persistence using the approach of Bai and Perron (1998, 2003). This approach allows an estimation of the parameters of a model with multiple structural changes. Therefore, we may estimate the different degrees of persistence along different regimes. Overall, the estimates indicate that allowing for one or more breaks reduces the estimated degrees of persistence. Most of provinces present a reduction of the persistence between the first and the last regimes. The highest values of degree of persistence are around 0.85, indicating a low speed of adjustment but still smaller compared with standard (high) measures of persistence.

[16] We thank an anonymous referee for pointing this out.

The rigidity of the Canadian labor market may explain the degree of persistence found in the estimations. Other factors that contribute to these levels of persistence are the unemployment insurance benefits and the program of inter-provincial redistribution. In 1996, however, the government applied a reform of the unemployment insurance system. This reform made the system much less generous and the eligibility condition much more stringent, and reduced the inter-regional variations in conditions and benefits. This implies a reduction in the values of the persistence of unemployment rates.

In addition, different industry structures and industry performances could create diverse adjustment speeds in the labor markets of the provinces. For example, the unemployment rate in Ontario would be more affected than the rate in Alberta by shocks to the manufacturing industry simply because the manufacturing industry is centered in Ontario. On the other hand, Alberta would be significantly affected by oil shocks. Government policies could also cause different degrees of persistence and adjustment speeds, because provincial governments might adopt different policies toward the labor markets. This may be seen, for example, in the attitude towards migration. Alberta, more than any other province, has historically attracted labor when it does well and shed labor when it has done poorly.

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Table 1. Two-Break LM Unit Root Statistics

	Model 1				Model 2			
Province	TB_1	TB_2	t-statistics	Lag	TB_1	TB_2	t-statistics	Lag
Alberta	1982:2	1994:3	-2.650	9	1981:2	1987:2	-6.202^a	9
BC	1982:2	1986:3	-3.315	10	1981:2	1987:4	-6.606^a	10
Manitoba	1984:2	1991:3	-2.668	3	1981:3	1987:1	-5.593^b	10
NB	1980:1	1981:3	-3.567	3	1984:4	1991:2	-5.715^b	7
NF	1981:2	1997:2	-3.740^c	7	1985:1	1992:1	-5.465^b	10
NS	1995:3	1997:2	-2.202	3	1986:4	1992:2	-5.858^a	10
Ontario	1994:2	1995:1	-3.616^c	9	1983:2	1991:2	-4.673	10
PEI	1987:2	1997:2	-2.181	7	1986:4	1992:1	-6.211^a	0
Quebec	1985:1	1996:4	-2.680	7	1982:3	1991:3	-5.031^c	7
SAS	1982:1	1997:4	-2.145	0	1982:1	1994:3	-5.411^b	3
Canada	1981:3	1982:1	-3.194	10	1983:1	1991:2	-5.497^b	10

^{a,b,c} denote statistical significance at 1%, 5% and 10%, respectively.

Table 2. Selection of Breaks using Bai and Perron's (1998, 2003) Procedure

Statistics	Alberta	BC	Manitoba	NB	NF	NS	Ontario	PEI	Quebec	Saskatchewan	Canada
$UD_{\max}$	30.840[a]	37.138[a]	18.399[a]	15.939[b]	32.441[a]	35.406[a]	25.801[a]	35.124[a]	34.491[a]	35.730[a]	38.750[a]
$WD_{\max}$	38.840[a]	37.138[a]	27.616[a]	26.357[a]	33.718[a]	45.733[a]	38.441[a]	35.124[a]	48.731[a]	45.249[a]	45.560[a]
$supF_T(1)$	38.840[a]	37.138[a]	16.233[b]	13.787[d]	32.441[a]	35.406[a]	19.557[b]	35.124[a]	8.056	35.730[a]	18.619[b]
$supF_T(2)$	21.175[a]	30.141[a]	16.005[a]	15.939[a]	25.225[a]	30.296[a]	25.801[a]	26.389[a]	34.491[a]	24.258[a]	38.750[a]
$supF_T(3)$	27.310[a]	27.683[a]	18.398[a]	15.304[a]	19.814[a]	33.989[a]	22.136[a]	23.005[a]	25.366[a]	28.081[a]	30.712[a]
$supF_T(4)$	22.400[a]	22.044[a]	17.877[a]	14.602[a]	23.341[a]	29.037[a]	24.554[a]	19.494[a]	21.451[a]	23.859[a]	26.591[a]
$supF_T(5)$	18.609[a]	19.538[a]	12.935[a]	14.065[a]	19.549[a]	22.191[a]	21.583[a]	15.212[a]	27.360[a]	24.146[a]	23.138[a]
$supF_T(2\|1)$	10.074	16.210[c]	15.408[c]	16.827[b]	16.954	18.096[b]	30.548[a]	12.139[d]	20.172[b]	10.167	55.667[a]
$supF_T(3\|2)$	10.169	18.699[c]	13.431	13.81	8.424	39.681[a]	25.527[a]	11.498	12.742	10.618	14.753
$supF_T(4\|3)$	10.791	9.514	15.408[d]	12.75	8.424	10.775	22.118[b]	6.409	12.056	8.187	11.871
$supF_T(5\|4)$	5.819	9.514		11.537	8.678		2.573				
BIC	1	1	0	0	0	3	0	1	0	1	2
LWZ	0	0	0	0	0	0	0	0	0	0	0
Sequential	1 [a,b,c,d]	1 [a,b,c] 3 [d]	0 [a] 1 [b,c] 4 [d]	0 [a,b,c] 2 [d]	1 [a,b,c,d]	1 [a] 3 [b,c,d]	0 [a] 3 [b,c,d]	1 [a,b,c,d]	0 [a,b,c,d]	1 [a,b,c,d]	0 [a] 2 [b,c,d]

[a,b,c,d] show significance at 1%, 2.5%, 5%, and 10%, respectively.

Table 3. Estimates of Persistence based on Bai and Perron (1998, 2003)

	Alberta	BC	Manitoba	NB	NF	NS
μ_1	1.474	-1.746	0.708	3.888	2.258	5.203
	(1.143)	(-1.347)	$(2.492)^a$	$(2.739)^a$	$(2.436)^a$	$(3.823)^a$
β_1	-0.022	0.049	0.006	-0.03	0.005	-0.028
	(-1.235)	$(2.383)^a$	$(1.642)^a$	$(-1.58)^c$	(0.689)	$(-1.773)^c$
ρ_{11}	0.519	1.671	0.886	0.698	0.697	0.516
	(1.446)	$(8.409)^a$	$(18.501)^a$	$(6.096)^a$	$(6.057)^a$	$(3.727)^a$
ρ_{21}	0.180	-0.514			0.267	
	(0.498)	$(-1.801)^c$			$(1.905)^b$	
ρ_{31}					0.192	
					(1.349)	
ρ_{41}					-0.300	
					$(-2.606)^a$	
μ_2	2.109	1.547	2.387	3.388	20.744	2.755
	$(5.667)^a$	$(2.082)^a$	$(1.672)^c$	$(3.595)^a$	$(4.691)^a$	$(3.567)^a$
β_2	-0.013	-0.008	-0.011	-0.014	-0.076	-0.041
	$(-5.292)^a$	$(-1.945)^b$	(-1.217)	$(-3.512)^a$	$(-4.541)^a$	$(-4.673)^a$
ρ_{12}	1.073	0.958	0.764	0.799	0.312	0.897
	$(11.572)^a$	$(10.008)^a$	$(7.747)^a$	$(13.528)^a$	$(2.136)^a$	$(17.063)^a$
ρ_{22}	-0.233	-0.063			0.063	
	$(-2.751)^a$	(-0.673)			(0.422)	
ρ_{32}					0.197	
					(1.311)	
ρ_{42}					-0.365	
					$(-2.673)^a$	
μ_3				20.359		-4.723
				$(3.495)^a$		(-1.461)
β_3				-0.092		0.236
				$(-2.904)^a$		$(1.590)^c$
ρ_{13}				-0.013		0.199
				(-0.048)		(0.391)
μ_4						5.002
						$(2.796)^a$
β_4						-0.025
						$(-2.405)^a$
ρ_{14}						0.741
						$(9.411)^a$
$\bar{R}^2$	0.965	0.952	0.926	0.891	0.815	0.943
F	403.420^a	285.613^a	243.32^a	107.569^a	42.728^a	161.275^a

^{a,b,c} denote significance at 1%, 5%, 5%, and 10%, respectively; t-statistics are shown in paranthesis.

Table 3 (continued). Estimates of Persistence based on Bai and Perron (1998, 2003)

	Ontario	PEI	Quebec	Saskatchewan	Canada
μ_1	3.28	3.91	2.064	1.598	2.286
	(2.785)[a]	(3.693)[a]	(2.025)[a]	(2.171)[a]	(2.613)[a]
β_1	0.006	0.032	0.008	0.003	0.001
	(0.566)	(3.018)[a]	(0.609)	(0.312)	(0.167)
ρ_{11}	1.146	0.597	0.799	0.628	1.342
	(4.958)[a]	(5.459)[a]	(7.512)[a]	(3.612)[a]	(7.621)[a]
ρ_{21}	-0.635				-0.634
	(-2.800)[a]				(-3.611)[a]
μ_2	5.584	12.176	14.309	2.775	6.528
	(5.473)[a]	(5.205)[a]	(6.335)[a]	(6.098)[a]	(6.858)[a]
β_2	-0.073	-0.063	-0.173	-0.011	-0.069
	(-5.182)[a]	(-5.032)[a]	(-6.654)[a]	(-5.598)[a]	(-6.468)[a]
ρ_{12}	1.241	0.528	0.352	0.710	1.059
	(10.638)[a]	(5.749)[a]	(3.189)[a]	(14.280)[a]	(8.479)[a]
ρ_{22}	-0.602				-0.446
	(-6.256)[a]				(-4.619)[a]
μ_3	4.014		1.470		2.637
	(6.040)[a]		(2.550)[a]		(4.620)[a]
β_3	-0.03		-0.007		-0.014
	(-5.548)[a]		(-2.529)[a]		(-4.465)[a]
ρ_{13}	1.030		0.920		1.186
	(8.510)[a]		(24.718)[a]		(10.331)[a]
ρ_{23}	-0.224				-0.345
	(-2.087)[a]				(-3.305)[a]
μ_4	5.801				
	(2.446)[a]				
β_4	-0.018				
	(-1.079)				
ρ_{14}	0.55				
	(1.385)				
ρ_{24}	-0.108				
	(-0.303)				
$\bar{R}^2$	0.967	0.897	0.941	0.937	0.976
F	213.197[a]	170.776[a]	209.519[a]	289.125[a]	390.785[a]

[a,b,c] denote significance at 1%, 5%, and 10%, respectively; t-statistics are shown in paranthesis.

Table 4. Estimated Break Points and their 90% Confidence Intervals

	Alberta	BC	Manitoba	NB	NF	NS	Ontario	PEI	Quebec	SAS	Canada
T_{B1}	1981:1	1982:1	1993:4	1981:1	1990:4	1981:1	1981:3	1990:2		1981:4	1981:3
$C.I.$	1980:3-1981:2	1981:3-1982:4	1993:2-1995:4	1980:4-1982:3	1990:3-1991:1	1980:3-1981:2	1981:1-1981:4	1989:3-1990:3		1981:1-1982:1	1981:1-1981:4
T_{B2}				2000:3		1988:4	1989:4				1989:4
$C.I.$				1998:4-2000:4		1988:2-1989:1	1989:2-1990:1				1989:2-1990:1
T_{B3}						1993:3	2000:4				
$C.I.$						1993:1-1993:4	2000:2-2001:2				

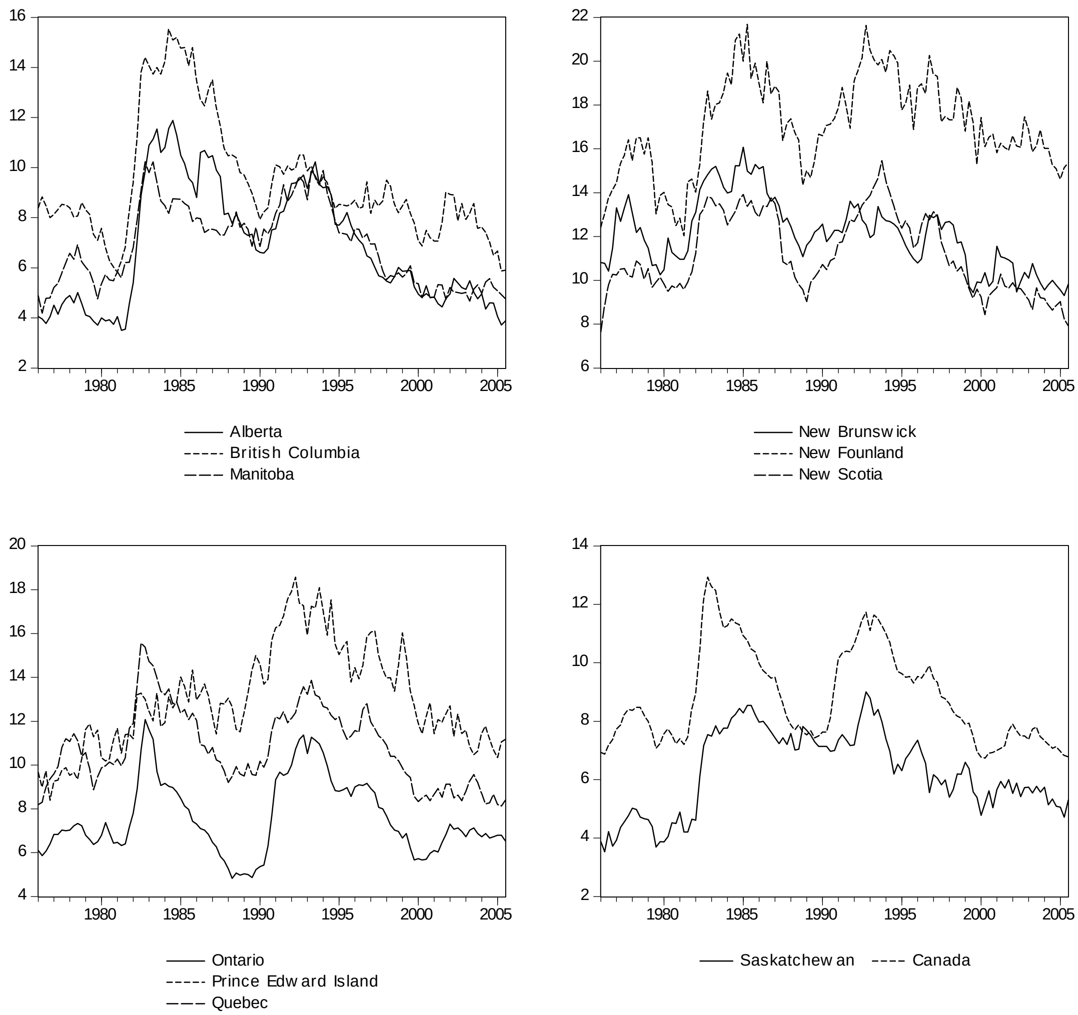

Figure 1. Unemployment Rates