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Convergence in the Canadian Provinces: Evidence using Unemployment Rates

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Abstract

Quarterly time series data from Canada and the Canadian provinces for the period 1976:1-2005:3 are examined to determine if the unemployment rates in the Canadian provinces are converging to the national rate of unemployment. Firstly, we check for existence of stochastic convergence using recent unit root statistics, see Perron and Rodríguez (2003) and Rodríguez (2007). Secondly, we verify for existence of $\beta-$convergence using recently proposed methods by Vogelsang (1998), Perron and Yabu (2009a, 2009b), and Bai and Perron (1998, 2003). Results from different unit root tests, without and with structural breaks, confirm that stochastic convergence exists in all provinces except British Columbia. The other results show strong evidence that deterministic convergence exists and the unemployment rates of the Canadian provinces are converging to the unemployment rate of Canada. This conclusion is stronger when multiple breaks are allowed in the trend function using the approach of Bai and Perron (1998, 2003).

Keywords: Stochastic convergence, β-convergence, unit roots, structural breaks, unemployment.

JEL classification: C22, O40, R00.

Resumen

Datos trimestrales para Canadá y sus provincias para el periodo 1976:1-2005:3 son analizados para determinar si las tasas de desempleo en las provincias Canadienses están convergiendo a la tasa nacional de desempleo. En primer lugar se verifica la existencia de convergencia estocástica usando recientes estadísticos de raíz unitaria, ver Perron y Rodríguez (2003) y Rodríguez (2007). En segundo lugar se contrasta la hipótesis de β convergencia usando métodos recientemente propuestos en la literatura por Vogelsang (1998), Perron y Yabu (2009a, 2009b), y Bai y Perron (1998, 2003). Los resultados de los diferentes estadísticos de raíz unitaria, con y son quiebre estructural confirman que convergencia estocástica existe excepto para British Columbia. Los otros resultados sugieren fuerte evidencia de convergencia determinística y que las tasas de desempleo de las provincias Canadienses están convergiendo a la tasa de desempleo nacional. Esta conclusión es más fuerte cuando múltiples quiebres son permitidos en la función de tendencia usando el método de Bai y Perron (1998, 2003).

Palabras Claves: Convergencia Estocástica, β convergencia, Raíces Unitarias, Quiebres Estructurales, Desempleo

Classificación JEL: C22, O40, R00.

Convergence in the Canadian Provinces: Evidence using Unemployment Rates[1]

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1 Introduction

Regional disparities have been in the centre of the policy makers concern for many years and has been studied by many researchers all over the world. Regional disparities are capable of creating imbalances in the development process of any country and in some countries, this issue has already created political problems. In Indonesia, for example, the rich regions are trying to separate from the central government because the people believe that their wealth is used to subsidize the poor regions.

Researchers use different measures of regional disparities, such as the relative per-capita incomes, per-capita earnings, wages, unemployment rates, and so on to study the existence and process of convergence. The main objective of this paper is to study the disparities in the unemployment rates in the Canadian provinces. Therefore, we use the difference in the unemployment rates of these provinces from the unemployment rate of Canada as our disparity indicator. We also want to investigate the trend of these disparities to see whether or not it has changed over last three decades.

The disparity problem is important in any country; however, it is more significant in diversified countries, like Canada. As Gunderson (1996, p.1) stated, "[b]eing a geographically large country with substantial heterogene-

[1] This paper is drawn from the second chapter of the PhD Dissertation of Firouz Fallahi at the Department of Economics of the University of Ottawa. We acknowledge comments from Gamal Atallah and Serge Coulombe.

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ity in resource endowments and accesses to markets, regional issues have always been important in Canada". Canada spent a massive amount of money during the 1970s and 1980s on a policy called economic development policy, to improve the economic conditions of the less developed regions. In 1968, the federal government created the Department of Regional Economic Expansion (DREE) which was active until 1981.[3]

The aim of the DREE was to promote economic expansion to adjust for the disparities in poor provinces. By doing so, Ottawa tried to minimize the disparity gaps in the growth rate and economic conditions among the Canadian provinces. In other words, the aim of the federal government of Canada was to create convergence among the Canadian provinces and one of the factors in this policy was the focus on unemployment rates.

Convergence is a simple notion which comes from neoclassical growth models, see Solow (1965) for example, for a closed economy. In these models, if the regions only differ from each other in the initial level of per-capita income and capital, they will reach the same steady-state level, since the productivity of capital in the poor regions is higher than the productivity of capital in the rich regions. Therefore, the poor regions will experience a higher growth rate than the rich regions. If the poor regions grow faster than the rich, they will eventually catch-up, the disparity gap will shrink and eventually close over time. On the other hand, if the growth rates of the rich regions are higher than that of poor regions, the disparity gap will increase over time. In this case, we say that these regions are diverging from each other.

Barro and Sala-i-Martin (1991) divided the concept of the convergence into three categories: σ-convergence, absolute β-convergence and conditional β-convergence and they measure two different things. The σ-convergence occurs when the dispersion of, say, per-capita income or unemployment

[3] In 1982, DREE was replaced by the Department of Regional Industrial Expansion (DRIE).

rate shrinks over the span of the study. Whereas, the β-convergence examines whether or not the poor regions are catching-up with the rich regions. In absolute β-convergence, the regions converge to the same level of, say, income or unemployment rate. However, the conditional β-convergence takes into consideration the differences in the regions, like different endowment, weather, infrastructure, and so on. Therefore, in the conditional β-convergence, the regions could have different steady sate levels and they do not need to converge to the same steady state level.

Another type of convergence that exists in literature is called stochastic convergence, developed by Bernard and Durlauf (1995). This deals with the effect of shocks on the variable under study, mainly using unit root tests and cointegration techniques. To have stochastic convergence in any variable, the variable must be stationary, i.e., the shocks to the variable must die-out exponentially.

Researchers adopted the definition of convergence of Sala-i-Martin (1991) and used cross sectional data and methods to test for the existence of convergence in different regions. Barro and Sala-i-Martin (1992) used personal income since 1840 and gross state product for the U.S. states since 1963 and found that the states had converged during the span of study. Barro (1991) applied cross section models to the GDP data of 98 countries from 1960 to 1985 and concluded that the trend of growth for the countries that had lower GDP in 1960 were steeper than that of the countries with high GDP in 1960. In other words, these countries were converging to each other. The results of these two papers show that the speed of convergence is about 2% both in the U.S. states and in 98 countries included in the study. However, they found that this speed is about 1% for 20 original members of OECD, which has been considered an homogenous group of countries in the study. This finding is one of the main critiques of Quah (1996) which showed that the speed of convergence is very sensitive to the cross sectional and time dimensions of the sample. Using Monte-Carlo simulations he showed that he could get 2% convergence rate for two independent random walk variables by changing the number of cross sectional units and the time span of the sample. He concluded that the results from the convergence studies for heterogenous group of regions might be misleading. He developed a theoretical model in which the poor regions create one polar and rich regions another, and over time rich regions become more prosperous, the poor become more underprivileged and the middle class disappears. More critics come from Friedman (1992), Quah (1993, 1996), and Bernard and Durlauf (1995). They criticized the definition and the use of the cross section models for testing the convergence.

Carlino and Mills (1993) developed the necessary conditions for convergence in per-capita income in U.S. regions using time series techniques. They stated that to have a convergence consistent with neoclassical growth models, two conditions are required. The series must be stationary to satisfy the stochastic convergence condition and they must also satisfy β-convergence. Myatt (1992), using annual data from 1966 to 1990 and Zellner's system of seemingly unrelated regression (SUR), studied the Canadian provincial unemployment rate disparities. He studied the trend, causes, and remedies of the provincial unemployment rates in Canada. The results showed that the unemployment rate disparity had an upward trend, which was worsened by federal to provincial transfer payments. He concluded that there was divergence in the unemployment rates in the Canadian provinces.

Coulombe and Lee (1995) used different measurements of per-capita income and output to investigate convergence in the Canadian provinces. Their study used a pooled cross section time series data of the Canadian provinces from 1961 to 1991. They concluded that the results are in favor of the existence of convergence among the Canadian provinces. Furthermore, Coulombe and Day (1999) studied the evolution of the relative regional income disparities in Canada and compared it with the evolution of twelve US states bordered with Canada. They confirmed that convergence is occurring in the Canadian regions.

DeJuan and Tomljanovich (2005) using a method proposed by Vogelsang (1998), investigated the existence of convergence in the Canadian provinces using personal income and earning from 1926 to 1996. Their results support the existence of both the stochastic and β-convergences for most of the provinces, after allowing for one break in the data. In addition, using the data of personal earnings they found that the governmental redistribution program had an accelerating effect on the convergence of the provinces.

Recently in another study, Rodríguez (2006) used two sets of data for per-capita income of the Canadian provinces from 1926 to 1999. One set includes the interprovincial transfers and the other without these transfers, to examine the convergence and assess the effects of interprovincial transfers on the convergence process. Using a class of tests proposed by Vogelsang (1998) and allowing for a break in the data, he studied the convergence in two cases of known and unknown break points. The results showed that the relative per-capita income of the Canadian provinces were converging and, in addition, the interprovincial transfers did not have any role in accelerating the convergence in these provinces.

Rodríguez-Oreggia (2005), studied the convergence process in the Mexican regions using per-capita GDP. He found that after liberalization in Mexico, the northern regions, bordering the U.S., have moved away from a "falling-behind" situation to being a situation that the author called a "winner". On the other hand, the southern regions fell behind. In addition, he concluded that human capital played an important role in these disparities. Costanini and Lupi (2006) using panel unit root tests, investigated properties of the Italian regional unemployment rates. They found some evidence of stochastic convergence in these regions, however, there was not enough evidence to conclude that the stochastic convergence existed among these regions.

In this paper, we use some very recent methods to study convergence in the unemployment rates of the ten Canadian provinces using quarterly unemployment data for the period 1976:1-2005:3.[4] After defining the data, we check for stochastic convergence using unit root tests without and with breaks. We use three unit root tests that allow for one and two break points in the data, see Perron and Rodríguez (2003), Rodríguez (2007), and Lumsdaine and Papell (1997). Next, we use three different methods to investigate the existence of β-convergence in the data. These methods are introduced by Vogelsang (1998), Perron and Yabu (2009a, 2009b), and Bai and Perron (1998, 2003). The first two methods are robust to the presence of unit roots, however, the last method is not. In addition, except for the latter method, which allows for multiple breaks, the other methods allow for only one break point in the data. We find strong evidence of convergence in most of the provinces by introducing structural breaks in the data.

The Canadian provinces are different in terms of their characteristics, industry structures, and industry performances, consequently, their labor market would be different as well. For example, most of manufacturing industries are centered in Ontario, while, most of oil related industries are located in Alberta. Therefore, oil shocks could be a potential cause of breaks in Alberta and any shock to the manufacturing industry would have a significant effect in Ontario. On another hand, different unemployment insurance system and also government policies toward the labor market could be considered as the cause of difference that we see in the unemployment rate behavior of the provinces. In addition, in theory, we would expect to have labor mobility between provinces, that is, from a province with a high unemployment rate to the provinces with a lower unemployment rate which would lead to a convergence. In the reality, however, factors like language barriers, social and financial uncertainties which arise when relocating could cause rigidity in labor mobility (Tarzwell, 1999). Moreover, provinces with

[4] Using similar data set, Fallahi and Rodríguez (2011) analyzed the topic of persistence in unemployment data series.

a diverse industrial mix and higher labor productivity will have lower unemployment rates.

The remainder of this paper is organized as follows. The next section, explains the notion of stochastic convergence and show the results obtained applying different unit root statistics without and with breaks. The third section presents the approaches of Vogelsang (1998), and Perron and Yabu (2009a, 2009b). The results obtained using these approaches are analyzed. Furthermore, we analyze the results obtained using the approach of Bai and Perron (1998, 2003). Section 4 concludes.

2 Stochastic Convergence

We use quarterly unemployment rates for the ten Canadian provinces for the period 1976:1-2005:4 obtained from Statistic Canada. The ten provinces are Alberta, British Columbia (BC), Manitoba, New Brunswick (NB), Newfoundland (NF), Nova Scotia (NS), Ontario, Prince Edward Island (PEI), Quebec, and Saskatchewan (SAS). Because we are interested in examining the trend of the regional disparities of the unemployment rates in the Canadian provinces, we define disparities as the difference between the unemployment rates of each province and the unemployment rate of Canada. Let y_t^i denotes| the difference between the unemployment rate of the province i at period t (u_t^i) and the unemployment rate of Canada at period t (u_t^{can}). The unemployment rate of Canada, u_t^{can}, is the reference variable, and consequently, convergence means that the unemployment rates of the provinces move toward the unemployment rate of Canada. Figures 1 and 2 show y_t^i for each province during 1976:1-2005:4. Figure 1, shows y_t^i for provinces where the unemployment rates are higher than the national rate of unemployment (Atlantic Canada[5] and Quebec). Figure 2, represents the y_t^i for a group of provinces that their unemployment rates are close and usually lower than

[5]We name Atlantic Canada as the territory formed by NB, NS, NF and PEI, respectively.

the unemployment rate in Canada.

Following Carlino and Mills (1993), convergence exists if stochastic convergence and β-convergence are verified. A natural way to examine the existence of stochastic convergence is to verify for stationarity of the respective time series using unit root statistics. If the null hypothesis of a unit root cannot be rejected, we say that the time series is not stationary and consequently there is not stochastic convergence. On the opposite case, if a rejection of the unit root hypothesis is obtained, the time series is stationary, shocks have transitory effects and stochastic convergence exists.

In formal terms, at this stage, 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.

In recent times, the most powerful unit root statistic developed in this literature is 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 GLS. In other words, $\tilde{y}_t$ is GLS detrended data at some particular point of the alternative hypothesis.[6] The null hypothesis is $\alpha_0 = \alpha - 1 = 0$, Δ is the first difference operator and k is the number of lags which must be determined in such a way that ϵ_t become serially uncorrelated.

[6] The alternative hypothesis is $H_1 : \alpha = \overline{\alpha} = 1 + \overline{c}/T$. Elliott, Rothenberg and Stock (1996) recommended to use $\overline{c} = -7.0$ when $z_t = \{1\}$ and $\overline{c} = -13.5$ when $z_t = \{1, t\}$. Data is transformed as $y_t^{\overline{\alpha}} = [y_1, (1 - \overline{\alpha}L)y_t]$, $z_t^{\overline{\alpha}} = [z_1, (1 - \overline{\alpha}L)z_t]$, for$t = 2, 3, ..., T$, and where $\overline{\alpha} = 1 + \frac{\overline{c}}{T}$.

Table 1 presents results obtained from the application of the ADF^{GLS} unit root statistics where the lag length is selected using BIC. The results show that the null hypothesis of unit root can be rejected for Manitoba, NB, NF, NS, PEI, Quebec and Saskatchewan. The null hypothesis cannot be rejected for Alberta, BC and Ontario. Therefore, according to these results, stochastic convergence is not found for these last three provinces.

A non rejection of the null hypothesis of a unit root may be due to a misspecification of the set of deterministic components (z_t). In fact, 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. The approach suggested by Perron (1989) was extended to an unknown break point by Zivot and Andrews (1992) an 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. Another interesting approach is suggested by Lumsdaine and Papell (1997) who proposed application of the minimum ADF unit root tests allowing for two breaks under the alternative hypothesis of stationarity. Furthermore, we consider the approach of Rodríguez (2007) who uses GLS detrended data for the so named crash model.

In fact, the Figures 1 and 2 seem to suggest possible presence of more than one break. Table 2 presents results using different unit root statistics. We present the MSB^{GLS} and the P_T^{GLS} statistics proposed by Perron and Rodríguez (2003) which allows for one break. Using these statistics, stochastic convergence is confirmed only for Atlantic provinces. Furthermore, we include the minimum ADF statistic of Lumsdaine and Papell (1997), which is denoted by LP in the Table 2.[7] Based on the results from LP, with two break points, the null hypothesis of unit root can be rejected for Alberta and Ontario. Therefore, British Columbia is the only province for which stochastic convergence is not found[8].

The dates of the break points are such diverse that is very difficult to find their relationship with some particular events of the Canadian economy. For some provinces, breaks appear to be concentrated around 1989-1991 (Alberta, Manitoba, NS, Ontario, PEI and Saskatchewan). In other cases, like Quebec, 2000-2001 appears to be a break.

The Canadian provinces are different in terms of their characteristics, industry structures, and industry performances, consequently, their labor market would be different as well. For example, most of manufacturing industries are centered in Ontario, while, most of oil related industries are located in Alberta. Therefore, oil shocks could be a potential cause of breaks in Alberta and any shock to the manufacturing industry would have a significant effect in Ontario.

3 β-Convergence

β-convergence implies that the provinces that initially had an unemployment rate lower than the national rate of unemployment will have a positive slope in the trend function. On the other hand, the provinces with a higher initial

[7] Lumsdaine and Papell (1997) proposed a modified version of the ADF test with two endogenous breaks:

$$\Delta y_t = \mu + \beta t + \theta_1 DU_{1t} + \delta_1 DT_{1t} + \theta_2 DU_{2t} + \delta_2 DT_{2t} + \alpha y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + \epsilon_t,$$

where $DU_{it} = 1$ if $t > TB_i$ $(i = 1, 2)$ and zero otherwise; $DT_{it} = t - TB_i$ if $t > TB_i$ $(i = 1, 2)$ and zero otherwise. The variables DU_{it} capture the structural breaks in the intercept at time TB_i while the other dummy variables, DT_{it} $(i = 1, 2)$ capture the structural changes in the slope at times TB_i $(i = 1, 2)$. The break points are selected based on the minimum value of the t statistic for $\widehat{\alpha}$.

[8] The results obtained applying Rodríguez (2007) do not differ from the results obtained using Perron and Rodríguez (2003). Results are available upon request.

unemployment rate than the national rate will have a negative slope in the trend function. These requirements can be directly linked to the parameters of the deterministic trend function of y_t.

Suppose that y_t is modeled as[9]

$$y_t = \mu + \beta t + u_t, \tag{4}$$

where μ represents the initial level of y_t, β represents the average change at y_t, and u_t is a serially correlated mean zero random process. The concept of β-convergence requires to have a negative relationship between μ and β, i.e. if $\mu > 0$ then $\beta < 0$ and if $\mu < 0$ then $\beta > 0$. Therefore, the evidence on β-convergence can be obtained from the estimates of the trend function in (4).

If u_t in (4) is serially uncorrelated, then we can use the OLS estimates of μ and β and these estimates will be unbiased and efficient. However, in practice u_t in (4) might be autocorrelated and may even have a unit root; and consequently, the interpretation of the trend parameters in the autoregressive representation of y_t depends on the nature of u_t. When u_t is an $I(0)$ process, then the inference about β can be obtained from the estimate of the slope. But if u_t is an $I(1)$ process this coefficient is zero and the inference must be made from an estimate of the intercept in the autoregressive representation of y_t. Therefore, the possibility of a unit root in u_t can make the interpretation of the coefficients of the trend function very difficult.

In order to overcome this problem, in this paper, we use two methods that are robust to the nature of u_t. These methods are proposed by Vogelsang (1998) and Perron and Yabu (2009a, 2009b) and are robust to the presence of a unit root and serial correlations in the series. The robustness of these methods to the presence of unit roots guarantees that the results

[9] In order to save notation, we eliminate the superscript i (indicating province i) in the variable y_t.

will not be spurious. Vogelsang (1998) proposed a class of statistics that consists of estimating two OLS regressions. The first regression, named the y_t regression, is given by

$$y_t = \mu_1 DU_{1t} + \beta_1 DT_{1t} + \mu_2 DU_{2t} + \beta_2 DT_{2t} + u_t \quad (5)$$

where $DU_{it} = 1$ if $t \leq T_B$ or zero otherwise $(i = 1, 2)$, $DT_{1t} = 1$ if $t \leq T_B$ or 0 otherwise, and $DT_{2t} = t - T_B$ if $t > T_B$ and zero otherwise, T_B is the date of a shift in the parameters of the trend function of y_t and is considered either known or unknown. The parameters μ_1 and μ_2 show whether the unemployment rate of the province is higher or lower than the unemployment rate of Canada in periods 1 and T_B, respectively.

The second regression, named the z_t regression, is given by:

$$z_t = \mu_1 DT_{1t} + \beta_1 SDT_{1t} + \mu_2 DT_{2t} + \beta_2 SDT_{2t} + S_t \quad (6)$$

where $z_t = \sum_{j=1}^{t} y_j$, $SDT_{it} = \sum_{j=1}^{t} DT_{ij}$, $i = 1, 2$, $S_t = \sum_{j=1}^{t} u_j$ and DT_{it} as defined before.

Let t_y and t_z denote the t-statistics for testing the null hypothesis that the individual parameters in the y_t and z_t regressions are zero. The modified t-statistics for the y_t regression is denoted by $T^{-1/2} y_t$, where T is the sample size. On the other hand, the modified t-statistics for z_t regression are defined as $t - PS_T = T^{-1/2} t_z \exp(-bJ_T)$, where b is a constant (to be calculated) and J_T is T^{-1} multiplied by the Wald statistic for testing $c_2 = c_3 = ... = c_9 = 0$ in the following regression[10]

$$y_t = \mu_1 DU_{1t} + \beta_1 DT_{1t} + \mu_2 DU_{2t} + \beta_2 DT_{2t} + \sum_{i=2}^{9} c_i t^i + u_t. \quad (7)$$

The J_T term can be calculated as $(RSS_Y - RSS_J)/RSS_J$, where RSS_Y is the sum of the squared residuals from regression (5), and RSS_J is the

[10] In fact, the J_T-statistic is a unit root statistic, which was proposed by Park and Choi (1988) and Park (1990).

sum of the squared residuals from regression (7). For a given significance level, the constant value of b can be computed in such a way that the critical values of t-statistics are the same whether u_t is $I(0)$ or $I(1)$.[11] Consequently, the modified t-tests by J_T from the z_t regression are robust to $I(1)$ errors. Note that if $b = 0$, the J_T modification will not have any effect on t-tests. Therefore, the distribution of $t - PS_T$ is different when u_t is $I(1)$ compared to the case when u_t is $I(0)$. Hence, it is recommended to only use $b = 0$ if we are sure that the errors are $I(0)$. On the other hand, there is no need for any modification in the y_t regression because $T^{-1/2}t_y$ statistics have well defined asymptotic distribution when u_t is $I(1)$ and when u_t is $I(0)$ these statistics converge to zero. Therefore, $T^{-1/2}t_y$ statistics are conservative tests when the errors are stationary.

The asymptotic distributions of $T^{-1/2}t_y$ and $t - PS_T$ are not standard and they depend on the break date used in the regressions which must be estimated from the data. The estimated break date is the break date that has the largest normalized Wald statistic. The critical values of the $T^{-1/2}t_y$ and $t - PS_T$ statistics are taken from Vogelsang (1998).

Tables 3-5 present the results obtained using the $t - PS_T$ without J_T correction, the $t - PS_T$ with J_T correction, and the $T^{-1/2}t_y$, respectively. These statistics are calculated considering an unknown break date in the regressions. The last column in these tables shows the estimated break dates.[12]

Table 3 reports the estimated μ, β and the break dates using regression z_t. The $t - PS_T$ statistics without J_T correction are given in parenthesis below each coefficient and the asymptotic critical values are given at the bottom two rows. The estimates of μ_1 for provinces are statistically different

[11] In addition, these critical values are robust to the serial correlation in the series.

[12] We also estimate the trend function using Vogelsang (1998) without any break points. The results show the convergence only for NB, QUE, and SAS. For other provinces no convergence found. To save the space we do not report these results, however, they are available upon request.

from zero for most of the provinces. This indicates that the disparities between the unemployment rates of the provinces and the national level of the unemployment rate existed in 1976. In addition, the results reveal that there is more evidence for deterministic convergence after the break than before. However, we should be cautious in using these results since these statistics are obtained assuming that the residuals are $I(0)$ process. This is because, even clearly stationary residuals could spuriously increase the $t - PS_T$ statistics in a small sample; see Tomljanovich and Vogelsang (2002).

The first method to overcome this problem is to use the $t-PS_T$ statistics using the z_t regression but corrected for the possibility that a unit root is present in the errors of process using J_T correction. The results are shown in Table 4. The estimated coefficients are the same as Table 3, but the statistics are smaller now. In this table, the $t - PS_T$ statistics are reported for 5% and 10% in parenthesis below each coefficient. These results show that before break, deterministic convergence is found in Manitoba, NF, NS, and Saskatchewan and there is a divergence in NB. On the other hand, after the estimated break, only PEI and NS demonstrate deterministic convergence and NF shows deterministic divergence.

The second way to overcome the problem of the behavior of u_t is to use the y_t regression to calculate the $T^{-1/2}t_y$ statistics. Table 5 reports the results. In the periods before the break, there was convergence in NF, NS, and Saskatchewan. At the same time, Alberta, Manitoba, NB, and Quebec have experienced divergence. After the break, there was convergence in PEI and divergence in NF.

In terms of the break dates, they appear to be clustered around a few dates. The first cluster of break dates are around 1990-1991 which is a known period of strong recession and affected particularly British Columbia, NB, Ontario, PEI, and Saskatchewan. Other dates are concentrated around 1982-1984 which are related to the oil crisis.

Using the same notation as Rodríguez (2006) Table 7 summarizes the results of Tables 3, 4, and 5. In this table, a (capital) C denotes point estimates consistent with β-convergence ($\mu > 0$ and $\beta < 0$, or $\mu < 0$ and $\beta > 0$) and both estimates are statistically significant at least at 10 percent level. A (small) c denotes point estimates consistent with β-convergence but only with one coefficient statistically significant at least at 10 percent level. Divergence is shown by D and d, where D indicates that both coefficients are statistically significant and d signifies that only one coefficient is statistically significant. An E denotes point estimates that are not statistically different from zero and are small in magnitude which implies that β-convergence has already occurred. Lastly, a u means that no conclusion is possible using all information from Tables 3 to 5. This situation is characterized when the coefficients are not significant but they are not small enough in magnitude to be considered as equilibrium (E).

According to Table 7 there is no evidence of β-convergence at all for NB and, in fact, the results are in favor of divergence in this province. This divergence is observed for this province in both periods before and after the break date. On the other hand, β-convergence has or is occurring for NS and Saskatchewan all the time. At the same time in NF we observe that β-convergence was occurring before break and after break there is strong evidence of divergence, that is the disparities between the unemployment rate of this province and national unemployment rate of Canada is increasing. Divergence is observed for PEI before the estimated break date; however, it changes to β-convergence after the break. We see β-convergence in Alberta only after the estimated break. This is also the case for Quebec. The results for Ontario are inconclusive and, finally, for BC the results support β-convergence and or equilibrium.[13]

[13] These results are consistent with the results Rodríguez (2006) who uses the same methodology to study the convergence in the per-capita personal income of the Canadian provinces and to investigate the effect of interprovincial transfers on the divergence or convergence process. He found that the most of provinces are converging or have converged

In a recent paper, Perron and Yabu (2009b) have proposed a novel procedure which has more power than the approach of Vogelsang (1998). The procedure is constructed based on Perron and Yabu (2009a). In this paper they analyzed the problem of hypothesis testing on the slope coefficient of a linear trend model when no information about the nature, $I(0)$ or $I(1)$, of the noise component is available. The method is based on Feasible Quasi Generalized Least Squares (FQGLS) that uses a superefficient estimate of the sum of the autoregressive parameters α when $\alpha = 1$. Perron and Yabu (2009b) extend the previous analysis to the case of testing for changes in level or slope of the trend function of an univariate time series. When the break date is known, the asymptotic results of Perron and Yabu (2009a) directly apply. If this is the case, Perron and Yabu (2009b) show that the procedure has good finite sample properties and a power function that is close to that attainable with the infeasible GLS procedure that uses the true value of α. When the break dates are unknown, the limit distributions of the Exp, $Mean$ and Sup functionals of the Wald tests across all permissible break dates (see Andrews and Ploberger, 1994) is no longer the same in the $I(0)$ and $I(1)$ cases. However, the limit distribution is nearly the same when considering the $Mean$ functional. Hence, it is possible to have tests with nearly the same size in both cases. To improve the finite sample properties of the test, Perron and Yabu (2009b) use a bias-corrected version of the OLS estimate of α as suggested by Roy and Fuller (2001). Perron and Yabu (2009b), based on simulations, show that their procedure has substantially more power than the procedure of Vogelsang (1998). Furthermore, their procedure has a power function that is close to that attainable if we knew the true value of α in many cases.

Let us assume that the univariate time series y_t, $t = 1, ..., T$, has been generated by (1) and (2). In addition, $-1 \leq \alpha \leq 1$ so that both stationary

into the unweighted provincial average. In addition, the results of this study showed that the interprovincial transfers did not have any significant role on the convergence of the provinces.

and integrated errors are allowed in the data generating process. Three models are considered. In the first model, there is a structural change in intercept, that is, $z_t = (1, DU_t, t)'$ and $\psi = (\mu_1, \mu_2, \beta_1)'$ where $DU_t = 1(t > T_B)$, T_B is the break date and $T_B = [\lambda_1 T]$. The hypothesis of interest in this case will be $\mu_2 = 0$. In the second model there is a structural change in slope, that is, $z_t = (1, t, DT_t)'$ and $\psi = (\mu_1, \beta_1, \beta_2)'$ where $DT_t = \mathbf{1}(t > T_B)(t - T_B)$. The null hypothesis in this case will be $\beta_2 = 0$. In the third model there is a structural change both in the intercept and the slope, that is, $z_t = (1, DU_t, t, DT_t)'$ and $\psi = (\mu_1, \mu_2, \beta_1, \beta_2)'$. The hypothesis of interest in this case is $\mu_2 = \beta_2 = 0$. In addition, $\mathbf{1}(.)$ is the indicator function, $\lambda_1 \in (0, 1)$, and $[.]$ denotes the largest integer. In all these models, we allow for only one break in intercept or slope or both of them.

When break date is unknown, Perron and Yabu (2009b) use the approach of Andrews and Ploberger (1994) and they suggest to use:

$$Mean - W_{FS} = T^{-1} \sum\nolimits_{\Lambda} W_{FS}(\lambda_1') \quad (8)$$

$$Exp - W_{FS} = \log[T^{-1} \sum\nolimits_{\Lambda} \exp(\frac{1}{2} W_{FS}(\lambda_1')] \quad (9)$$

$$\sup - W_{FS} = \sup_{\Lambda} W_{FS}(\lambda_1') \quad (10)$$

where W_{FS} is the Wald statistic using feasible generalized squares, $\Lambda = \{\lambda_1'; \epsilon \leq \lambda_1' \leq 1 - \epsilon\}$ for some $\epsilon > 0$. Notice that T_B' denotes a generic break date, which is used to construct the Wald statistic and T_B denotes the true break date. These statistics are optimal only when the errors are stationary. The asymptotic critical values are calculated by Perron and Yabu (2009b). They show that although the limiting distribution of these statistics are different, when considering the $Exp - W_{FS}$ functional, the relevant quantiles are very similar for the $I(0)$ and $I(1)$ cases. On the other hand, the results from the simulation show that the $Mean$ and Sup versions of the test are less useful than the Exp version.

It is known that the OLS estimate of the autoregressive coefficient is biased downward especially when α is near unity. Therefore, to improve finite sample properties, they use a modified version of the weighted symmetric least square estimates described by Roy and Fuller (2001). The procedure consists of the following steps:

1. For any given break date, use OLS to detrend the data and obtain the residuals $\widehat{u}_t$;

2. Estimate an $AR(1)$ for $\widehat{u}_t$ in order to find the estimate of $\hat{\alpha}$ and the $t-$ratio, $\hat{\tau}$;

3. Use $\hat{\alpha}$ and $\hat{\tau}$ to get the Roy and Fuller (2001) bias corrected estimate $\hat{\alpha}_M$;

4. Apply the following truncation

$$\hat{\alpha}_{MS} = \begin{cases} \hat{\alpha}_M, & \text{if } |\hat{\alpha}_M - 1| > T^{-1/2} \\ 1, & \text{if } |\hat{\alpha}_M - 1| \leq T^{-1/2} \end{cases};$$

5. Use $\hat{\alpha}_{MS}$ and GLS method to estimate the coefficients of the trend function and obtain the estimate of the variance of the residuals. Then create the standard Wald statistic;

6. For a case with unknown break date, repeat the previous steps for all possible break dates and construct the statistics given in (8)-(10).

The results from Yabu and Perron's (2009b) approach are presented in Table 6.[14] They reject the null hypothesis that the trend function is stable in favor of a trend function with a break in all provinces with the exception of Manitoba, NB, PEI, and Saskatchewan. This table also shows the estimated break date obtained by minimizing the sum of squared residuals and they are the same as the break dates estimated using the method of Vogelsang (1998). The estimated μ and β for the periods before and after the estimated breaks are also shown in Table 6. According to these results, Ontario and

[14] The subscript RQF stands for Robust Quasi Feasible GLS.

Quebec experience divergence from the national rate of unemployment in Canada during the sample period. On the other hand, the results support the deterministic convergence for Alberta, BC, and NS. Finally, the unemployment rate in NF was approaching the national rate before the break in 1983:3 so there is evidence of β-convergence in this province before break date. However, in the period after the break, we see divergence. For Manitoba, NB, PEI, and Saskatchewan, this test cannot reject the null hypothesis of stable trend function; therefore, we cannot use these results to discuss the divergence or convergence in these provinces.

An important issue is worth to be discussed at this point. The results may be sensitive to the fact that only one break is allowed in the above statistics. Perron and Yabu (2009a, 2009b) give some indications about how to extend their procedure to the case where multiple breaks are present in the time series. However, it is beyond the scope of the paper and we prefer to follow another approach.

When the time series is stationary, Bai and Perron (1998, 2003) have proposed a method to estimate the coefficients in a multiple linear model with endogenous structural changes. Firstly, the method allows to test for the number of breaks using different statistics. Second, if the null hypothesis of no structural change is rejected, the method allows to estimate the parameters for each regime.

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

$$\left\{ \begin{array}{lr} y_t = \mu_1 + \beta_1 t + u_t, & t = 1, 2, ..., T_1 \\ y_t = \mu_2 + \beta_2 t + u_t, & t = T_1 + 1, ..., T_2 \\ \vdots & \\ y_t = \mu_m + \beta_m t + u_t, & t = T_m + 1, ..., T \end{array} \right\} \quad (11)$$

where y_t is the dependent variable, μ_i is the intercept, and β_i is the slope (for $i = 1, 2, ..., m$). We allow all coefficients to be regime-dependent. 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 $(m = h)$[15]. 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 $WD_{\max}F_T(M, q)$ and the critical values can be found in Bai and Perron (2003).

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 minimizing information criteria such as the Bayesian Information Criterion (BIC) and the criterion proposed by Liu, Wu and Zidek (1997), denoted by LWZ. 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.

The results are reported in Table 8. This method allows for breaks in μ and β in the trend function. In the estimation process we use 15% trimming from both ends of the sample (17 observations) and we also impose the maximum number of breaks equal to 5. According to this table, the statistics UD_{MAX}, WD_{MAX}, and $SupF(0|l)$ suggest the existence of at least one break in each time series. The number of breaks can be determined using information criteria, such as BIC and LWZ, using $SupF(l + 1|l)$, and the sequential procedure.

The number of breaks for each series is reported in the bottom three rows of Table 8. Based on the sequential procedure at 5% level of significance,

[15] 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 5. We use this parameterization here.

there are 2 or 3 breaks in each series except for Manitoba with no break and Quebec with 4 breaks. However, at 10% level of significance, the results of this procedure show that there are 3 break points in the series for Manitoba, that is, there are 4 different regimes in the time series of disparities in Manitoba. For all series, we use the results from the sequential procedure at 5% level of significance, the exception being Manitoba, for which we use the result at 10% level.

The results from the approach of Bai and Perron, Table 9, show that the unemployment rate disparity was significant in 1976 and the unemployment rate in Alberta, Manitoba, Ontario, and SAS were below the national unemployment rate. On the other hand, all the remaining provinces had an unemployment rate higher than the national rate. It is also interesting to note that the disparity for Alberta was negative in all the regimes, i.e., it had experienced a lower unemployment rate than Canada at all the regimes. As before, we can use the estimates of μ and βs to study the (di)convergence in the provinces allowing for up to 5 break points in the data.

The results from the Table 9 are summarized in Table 10. In this Table we use the same notation as Table 6 to investigate the deterministic convergence or divergence in the series. We also call the period between two breaks, a regime. According to this Table, there is a more evidence of β-convergence in all provinces.

In all regimes, we see that BC[16], Manitoba, NB, and NS have experienced β-convergence. However, the other provinces had experienced convergence in some period and divergence in the others. Except for the second regime, Alberta shows divergence and NF supports the existence of β-convergence. In Quebec and PEI, we observe that the rate of unemployment of these provinces are converging to the Canadian rate of unemployment after the first break in 1982:3 and 1982:2, respectively. Before the second break in

[16] The results for BC should be approached cautiously, since based on the results from section 2. Apparently this time series has unit root and the approach of Bai and Perron (1998, 2003) is not robust to the presence of unit root.

1990:2, Ontario was diverging; however, there is clear convergence after this break date. Finally, the convergence in SAS changes to divergence after the second break in 1990:3.

The estimated break dates and 90% confidence interval of them along with the estimated coefficients and their t-statistics are reported in Table 9. The estimated break dates are very precise and at 90% confidence interval they have less than 1.5 years of standard errors. The point estimates of the first break dates are around 1982-1984 and the second and the third break dates are around 1989-1992 and 1996-1998, respectively. The last break date for Quebec is in 2001.

Different unemployment insurance system and also government policies toward the labor market could be considered as the cause of differences that we see in the unemployment rate behavior of the provinces. In addition, while, in theory, we would expect to have labor mobility between provinces, that is, from a province with a high unemployment rate to the provinces with a lower unemployment rate which would lead to a convergence. In the reality, however, factors like language barriers, social and financial uncertainties which arise when relocating, could cause rigidity in labor mobility (Tarzwell, 1999).

4 Conclusions

This paper has examined the stochastic and deterministic convergence of the unemployment rates in the Canadian provinces for the period of 1976-2005, by using recently proposed methodologies in the literature. Using the ADF^{GLS} unit root test, we found evidence of the stochastic convergence for most, but not all of the provinces. Therefore, we used some new tests of unit root, which allow for the break in the data. By applying these tests and allowing for one and two structural breaks in the data, we could reject the null of unit root and conclude that the stochastic convergence exists for all the provinces except British Columbia. Next, we test for the deterministic convergence in the data, using tests proposed by Vogelsang (1998), Perron and Yabu (2009a, 2009b), and Bai and Perron (1998, 2003). The results of these tests show that the by introducing multiple structural breaks in the data, we get more evidence of the deterministic convergence in the Canadian provinces.

Overall, combining the results from the study of stochastic convergence in Section 2 with the results that we obtained using methodologies of Vogelsang, Perron and Yabu, and Bai-Perron, there is strong evidence that the unemployment rates of these provinces are converging to the national level of the unemployment rate. This conclusion is stronger when we allow for more than one break in the trend function.

The Canadian provinces are different in terms of their characteristics, industry structures, and industry performances, consequently, their labor market would be different as well. For example, most of manufacturing industries are centered in Ontario, while, most of oil related industries are located in Alberta. Therefore, oil shocks could be a potential cause of breaks in Alberta and any shock to the manufacturing industry would have a significant effect in Ontario. Different unemployment insurance system and also government policies toward the labor market could be considered as the cause of difference that we see in the unemployment rate behavior of the provinces. In addition, while, in theory, we would expect to have labor mobility between provinces, that is, from a province with a high unemployment rate to the provinces with a lower unemployment rate which would lead to a convergence. In the reality, however, factors like language barriers, unemployment insurance policies, social and financial uncertainties which arise when relocating, could cause rigidity in labor mobility (Tarzwell, 1999). Other factors such as participation rate and industrial mix play an important role, provinces with a diverse industrial mix will have lower unemployment rate. Labor productivity is considered as another possible cause of differences in the labor markets of provinces.

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Table 1. Results from unit root tests without break

Province	ADF^{GLS}	
	$z_t = \{1\}$	$z_t = \{1, t\}$
Alberta	-1.547	-1.601
British Columbia	-1.223	-1.585
Manitoba	-3.050^{a}	-3.061^{c}
New Brunswick	-2.828^{a}	-3.389^{c}
Newfoundland	-0.972	-3.038^{c}
Nova Scotia	-2.077^{c}	-2.937^{d}
Ontario	-1.232	-1.352
PEI	-2.117^{c}	-2.873^{d}
Quebec	-2.334^{c}	-2.763^{d}
Saskatchewan	-1.806^{d}	-2.254

^{a}, ^{b}, ^{c} and ^{d} show significance at 1%, 2.5%, 5%, and 10% level of significance, respectively.

Table 2. Unit Root Statistics with Structural Change

Province	Statistic	$t_{\hat{\alpha}}$	k^*	T_{B1}	T_{B2}
Alberta	MSB	0.174	1	1983:2	
	P_T	17.081	1	1983:2	
	LP	-6.357[c]	0	1982:4	1989:3
British Columbia	MSB	0.273	1	1989:4	
	P_T	41.069	1	1989:4	
	LP	-4.486	0	1981:3	1997:1
Manitoba	MSB	0.169	1	1987:1	
	P_T	15.115	1	1987:1	
	LP	-5.999[d]	0	1987:2	1990:2
New Brunswick	MSB	0.125 [b]	1	1991:4	
	P_T	7.988 [a]	1	1991:4	
	LP	-5.578	7	1991:4	1996:2
Newfoundland	MSB	0.110 [a]	3	1983:2	
	P_T	6.543 [a]	3	1983:2	
	LP	-5.959[d]	3	1979:2	1983:4
Nova Scotia	MSB	0.125 [b]	1	1997:4	
	P_T	8.852 [b]	1	1985:2	
	LP	-7.135[c]	0	1984:2	1992:4
Ontario	MSB	0.225	1	1989:3	
	P_T	27.312	1	1984:1	
	LP	-6.786[c]	7	1983:2	1990:2
PEI	MSB	0.138 [d]	1	1989:2	
	P_T	10.211 [c]	1	1987:2	
	LP	-5.219	0	1984:3	1988:4
Quebec	MSB	0.150	1	1977:1	
	P_T	11.419 [d]	1	1977:1	
	LP	-4.689	7	1983:4	1999:3
Saskatchewan	MSB	0.196	1	1993:1	
	P_T	19.724	1	1993:1	
	LP	-4.508	8	1985:1	1990:3

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

Table 3. Results using the Statistic $t - PS_T$ without J_T correction

Province	$\hat{\mu}_1$	$\hat{\beta}_1$	$\hat{\mu}_2$	$\hat{\beta}_2$	$\hat{T}_B$
	$(t - stat)$	$(t - stat)$	$(t - stat)$	$(t - stat)$	
Alberta	-3.319^d	-1.557	0.258	-4.054^c	1982:4
	(-1.956)	(-0.120)	(0.437)	(-3.176)	
British Columbia	-0.806	8.869^c	-1.495^d	3.891^d	1989:3
	(-0.967)	(2.523)	(-1.295)	(1.062)	
Manitoba	-2.208^c	0.267	-0.840^d	-2.837^c	1987:1
	(-3.427)	(0.083)	(-1.484)	(-1.852)	
New Brunswick	3.777^c	0.508	1.976^c	2.491	1990:3
	(6.902)	(0.234)	(2.186)	(0.810)	
Newfoundland	7.486^c	-6.405	8.526^c	0.929	1983:3
	(6.295)	(-0.771)	(17.064)	(0.831)	
Nova Scotia	2.439^c	-2.151	2.898^c	-0.824	1984:3
	(3.004)	(-0.422)	(6.751)	(-0.814)	
Ontario	-0.470^d	-4.450^c	-0.606^d	-0.486	1990:2
	(-1.575)	(-3.704)	(-1.284)	(-0.308)	
PEI	1.930^c	2.132	7.136^c	-4.595^c	1989:1
	(2.572)	(0.652)	(7.501)	(-1.575)	
Quebec	1.847^d	4.029	1.927^c	-0.074	1982:3
	(1.578)	(0.436)	(5.045)	(-0.091)	
Saskatchewan	-4.434^c	5.722^c	-3.193^c	2.763	1990:3
	(-5.705)	(1.856)	(-2.488)	(0.633)	
10% critical value	±1.570	±1.330	±1.140	±0.936	
5% critical value	±2.190	±1.760	±1.500	±1.270	

^{a}, ^{b}, ^{c} and ^{d} show significance at 1%, 2.5%, 5%, and 10% level of significance, respectively.

Table 4. Results using the Statistic $t - PS_T$ with J_T correction

Province	$\hat{\mu}_1$	$\hat{\beta}_1$	$\hat{\mu}_2$	$\hat{\beta}_2$	$\hat{T}_B$
	(5% t-stat)	(5% t-stat)	(5% t-stat)	(5% t-stat)	
	(10% t-stat)	(10% t-stat)	(10% t-stat)	(10% t-stat)	
Alberta	-3.319	-1.557	0.258	-4.054^d	1982:4
	(-1.212)	(-0.028)	(0.064)	(-0.782)	
	(-1.302)	(-0.043)	(0.106)	(-1.158)	
British Columbia	-0.806	8.869	-1.495	3.891	1989:3
	(-0.073)	(0.001)	(0.000)	(0.001)	
	(-0.108)	(0.01)	(-0.001)	(0.005)	
Manitoba	-2.208^d	0.267	-0.84	-2.837	1987:1
	(-2.016)	(0.016)	(-0.177)	(-0.392)	
	(-2.182)	(0.026)	(-0.309)	(-0.606)	
New Brunswick	3.777^c	0.508	1.976	2.491	1990:3
	(4.848)	(0.080)	(0.531)	(0.288)	
	(5.111)	(0.109)	(0.769)	(0.385)	
Newfoundland	7.486^c	-6.405	8.526^c	0.929	1983:3
	(4.881)	(-0.354)	(6.162)	(0.395)	
	(5.070)	(-0.446)	(8.039)	(0.486)	
Nova Scotia	2.439^c	-2.151	2.898^c	-0.824	1984:3
	(2.392)	(-0.210)	(2.711)	(-0.418)	
	(2.475)	(-0.258)	(3.440)	(-0.504)	
10% critical value	±1.570	±1.330	±1.140	±0.936	
5% critical value	± 2.190	±1.760	±1.500	±1.270	

^{a}, ^{b}, ^{c} and ^{d} show significance at 1%, 2.5%, 5%, and 10% level of significance, respectively.

Table 4. (continued). Results using the Statistic $t - PS_T$ with J_T correction

Province	$\hat{\mu}_1$ (5% t-stat) (10% t-stat)	$\hat{\beta}_1$ (5% t-stat) (10% t-stat)	$\hat{\mu}_2$ (5% t-stat) (10% t-stat)	$\hat{\beta}_2$ (5% t-stat) (10% t-stat)	$\hat{T}_B$
Ontario	-0.470	-4.450	-0.606	-0.486	1990:2
	(-0.294)	(-0.022)	(-0.002)	(-0.002)	
	(-0.378)	(-0.100)	(-0.009)	(-0.009)	
PEI	1.930^d	2.132	7.136^d	-4.595	1989:1
	(1.499)	(0.125)	(0.864)	(-0.324)	
	(1.625)	(0.204)	(1.519)	(-0.505)	
Quebec	1.847	4.029	1.927	-0.074	1982:3
	(0.802)	(0.055)	(0.336)	(-0.013)	
	(0.888)	(0.101)	(0.681)	(-0.022)	
Saskatchewan	-4.434^d	5.722	-3.193	2.763	1990:3
	(-1.831)	(0.058)	(-0.026)	(0.023)	
	(-2.170)	(0.160)	(-0.086)	(0.058)	
10% critical value	$\pm$1.570	$\pm$1.330	$\pm$1.140	$\pm$0.936	
5% critical value	$\pm$ 2.190	$\pm$1.760	$\pm$1.500	$\pm$1.270	

^{a}, ^{b}, ^{c} and ^{d} show significance at 1%, 2.5%, 5%, and 10% level of significance, respectively.

Table 5. Results using the Statistic $T^{-1/2}t_y$

Province	$\hat{\mu}_1$	$\hat{\beta}_1$	$\hat{\mu}_2$	$\hat{\beta}_2$	$\hat{T}_B$
	(t-stat)	(t-stat)	(t-stat)	(t-stat)	
Alberta	-3.392^c	-0.722	0.044	-3.457	1982:4
	(-1.223)	(-0.043)	(0.029)	(-1.213)	
British Columbia	-0.389	6.882	-0.797	1.826	1989:3
	(-0.130)	(0.738)	(-0.287)	(0.246)	
Manitoba	-2.149^c	-0.278	-0.838	-2.495	1987:1
	(-1.189)	(-0.041)	(-0.598)	(-0.769)	
New Brunswick	3.755^c	0.700	2.040	1.883	1990:3
	(1.629)	(0.105)	(0.893)	(0.289)	
Newfoundland	7.471^c	-6.713	8.775^c	0.254	1983:3
	(2.093)	(-0.345)	(4.208)	(0.063)	
Nova Scotia	2.452^c	-2.342	3.066	-1.351	1984:3
	(1.075)	(-0.212)	(2.110)	(-0.455)	
Ontario	-0.570	-4.022	-0.828	0.350	1990:2
	(-0.514)	(-1.229)	(-0.766)	(0.115)	
PEI	1.666	3.516	6.917^d	-4.359	1989:1
	(0.515)	(0.337)	(2.392)	(-0.581)	
Quebec	1.725^c	5.249	1.995	-0.383	1982:3
	(0.985)	(0.480)	(2.145)	(-0.220)	
Saskatchewan	-4.510^c	6.191	-3.342	3.009	1990:3
	(-2.063)	(0.977)	(-1.542)	(0.487)	
10% critical value	±0.671	±1.470	±2.370	±1.480	
5% critical value	±0.875	±2.000	±3.000	±2.010	

^{a}, ^{b}, ^{c} and ^{d} show significance at 1%, 2.5%, 5%, and 10% level of significance, respectively.

Table 6. Results using Perron and Yabu (2009a, 2009b)

Province			Intercept		Slope	
	W_{RQF}	$\hat{T}_B$	Pre-break	Post-break	Pre-break	Post-break
Alberta	11.778^a	1982:4	-3.392^c	3.638^c	-0.007	-0.035^c
British Columbia	3.752^b	1989:3	-0.389	-4.193^c	0.069^c	0.018^c
Manitoba	0.769	1987:1	-2.149^c	1.437^c	-0.003	-0.025^c
New Brunswick	0.923	1990:3	3.755^c	-2.127^c	0.007	0.019
Newfoundland	2.799^c	1983:3	7.471^c	3.386^c	-0.067^c	0.003^c
Nova Scotia	2.537^c	1984:3	2.452^c	1.435^c	-0.023^c	-0.014
Ontario	3.614^b	1990:2	-0.571^c	2.075^c	-0.040^c	0.004^c
PEI	0.972	1989:1	1.667^c	3.387^c	0.035^c	-0.044^c
Quebec	3.965^b	1982:3	1.725^c	-1.148^c	0.053^c	-0.004^c
Saskatchewan	1.196	1990:3	-4.511^c	-2.485^c	0.062^c	0.030^c

^{a}, ^{b}, ^{c} and ^{d} show significance at 1%, 2.5%, 5%, and 10% level of significance, respectively.

Table 7. Summary of Results using Volgelsang (1998) and Perron and Yabu (2009a, 2009b)

Province	Vogelsang						Yabu	
	$t-PS_T$		$t-PS_T$		$T^{-1/2}t_y$		W_{RQF}	
	$I(0)$ Errors Assumed		Robust to $I(1)$ Errors		Robust to $I(1)$ Errors		Robust to $I(1)$ Errors	
	Pre-break	Post-break	Pre-break	Post-break	Pre-break	Post-break	Pre-break	Post-break
Alberta	d	c	u	c	d	u	c	C
British Columbia	c	C	u	E	u	u	c	C
Manitoba	c	D	c	u	d	u	d	C
New Brunswick	d	d	d	u	d	u	d	c
Newfoundland	c	d	c	d	c	d	C	D
Nova Scotia	c	c	c	c	c	u	C	c
Ontario	D	d	u	E	u	u	D	D
PEI	d	C	d	c	u	c	D	C
Quebec	d	c	u	u	d	u	D	D
Saskatchewan	C	c	c	u	c	u	C	C

Table 8. Results using Bai and Perron (1998, 2003)

Statistics	Alberta	BC	Manitoba	NB	NF	NS	Ontario	PEI	Quebec	Saskatchewan
UD_{MAX}	248.248^{a}	474.371^{a}	70.803^{a}	75.874^{a}	60.010^{a}	76.882^{a}	291.746^{a}	175.221^{a}	47.552^{a}	110.638^{a}
WD_{MAX}	396.022*	930.092*	102.582*	121.039*	88.647*	105.483*	572.021*	206.132*	90.680*	132.228*
$\sup F_T(1)$	69.456^{a}	442.603^{a}	10.454^{d}	19.115^{a}	55.658^{a}	16.829^{a}	84.442^{a}	134.016^{a}	36.540^{a}	16.947^{a}
$\sup F_T(2)$	241.268^{a}	64.336^{a}	25.538^{a}	35.846^{a}	50.541^{a}	38.732^{a}	103.881^{a}	175.221^{a}	27.765^{a}	110.638^{a}
$\sup F_T(3)$	112.647^{a}	255.582^{a}	70.803^{a}	39.906^{a}	63.010^{a}	76.882^{a}	99.120^{a}	83.132^{a}	31.758^{a}	96.375^{a}
$\sup F_T(4)$	248.248^{a}	240.970^{a}	64.304^{a}	75.874^{a}	55.468^{a}	63.020^{a}	207.213^{a}	65.726^{a}	47.552^{a}	67.788^{a}
$\sup F_T(5)$	200.624^{a}	474.371^{a}	48.805^{a}	59.799^{a}	45.212^{a}	44.803^{a}	291.746^{a}	67.467^{a}	46.249^{a}	56.284^{a}
$\sup F_T(2\|1)$	24.862^{a}	41.604^{a}	35.450^{a}	34.389^{a}	23.225^{a}	50.859^{a}	39.805^{a}	13.547^{c}	18.549^{a}	93.789^{a}
$\sup F_T(3\|2)$	1.822^{d}	80.208^{a}	18.474^{a}	18.566^{a}	8.505	3.623	69.349^{a}	4.57	26.431^{a}	43.756^{a}
$\sup F_T(4\|3)$	16.808^{b}	82.495^{a}	24.572^{a}	6.816	16.776^{b}	2.658			26.431^{a}	6.506
$\sup F_T(5\|4)$		32.167^{a}		1.453						
BIC	5	5	4	3	3	3	4	3	4	4
LWZ	4	4	3	3	1	3	4	2	3	3
Sequential	2^{a,b,c,d}	3^{a,b,c,d}	0^{a,b,c} 3^{d}	3^{a,b,c,d}	2^{a,b,c,d}	2^{a,b,c,d}	3^{a,b,c,d}	2^{a,b,c} 3^{d}	4^{a,b,c,d}	3^{a,b,c,d}

$^{a, b, c}$ and ^{d} show significance at 1%, 2.5%, 5%, and 10% level of significance, respectively; * denoets significance at 5% level of significance.

Table 9. Estimates using Bai and Perron (1998, 2003)

	Alberta	BC	Manitoba	NB	NF	NS	Ontario	PEI	Quebec	SAS
μ_1	-3.392	0.437	-2.433	4.616	7.471	2.452	-0.992	1.519	1.725	-3.513
	(-16.793)	(1.691)	(-13.714)	(21.068)	(23.467)	(13.427)	(-10.464)	(4.394)	(15.131)	(-18.306)
β_1	-0.007	-0.012	0.029	-0.055	-0.067	-0.023	-0.002	0.080	0.052	0.012
	(-0.594)	(-0.803)	(2.445)	(-5.046)	(-3.865)	(-2.648)	(-0.4)	(3.577)	(7.376)	(0.89)
μ_2	-1.896	5.634	-4.043	4.891	8.478	4.911	-2.254	-4.898	3.478	-8.551
	(-3.453)	(6.744)	(-6.738)	(5.937)	(20.247)	(9.614)	(-8.417)	(-5.681)	(9.209)	(-25.971)
β_2	0.044	-0.059	0.046	-0.011	0.006	-0.045	-0.006	0.188	-0.046	0.152
	(3.39)	(-3.105)	(2.754)	(-0.616)	(0.845)	(-4.624)	(-1.093)	(8.897)	(-4.789)	(19.945)
μ_3	-0.475	2.597	2.349	3.518	15.822	5.841	-1.471	9.228	2.170	-2.731
	(-1.511)	(2.616)	(7.046)	(2.658)	(7.825)	(12.209)	(-3.725)	(18.832)	(6.179)	(-3.954)
β_3	-0.019	-0.048	-0.057	-0.022	-0.065	-0.036	0.011	-0.044	-0.003	-0.004
	(-5.507)	(-3.435)	(-11.931)	(-1.196)	(-3.37)	(-7.216)	(2.04)	(-7.88)	(-0.504)	(-0.45)
μ_4		2.455	-1.509	6.197			-4.386		8.103	-0.723
		(1.88)	(-1.416)	(6.061)			(-8.27)		(7.828)	(-0.679)
β_4		-0.020	-0.006	-0.031			0.035		-0.065	-0.010
		(-1.623)	(-0.585)	(-3.081)			(6.849)		(-5.839)	(-0.984)
μ_5									-0.142	
									(-0.09)	
β_5									0.013	
									(0.932)	
$\bar{R}^2$	0.846	0.801	0.648	0.666	0.644	0.449	0.9	0.8	0.709	0.839
F	109.038^a	60.349^a	28.015^a	30.270^a	36.414^a	16.855^a	134.361^a	79.625^a	29.693^a	77.823^a

Table 9 (continued). Estimates using Bai and Perron (1998, 2003)

	Alberta	BC	Manitoba	NB	NF	NS	Ontario	PEI	Quebec	SAS
Estimate Break Dates and their 90% confidence intervals										
T_{B1}	1982:4	1983:2	1982:1	1984:2	1983:3	1984:3	1983:2	1982:2	1982:3	1981:4
90% $C.I.$	1982:2-1983:1	1983:1-1983:2	1981:3-1982:2	1983:4-1985:2	1983:1-1984:1	1983:2-1984:4	1982:4-1983:4	1981:4-1982:3	1982:1-1983:4	1981:1-1982:1
T_{B2}	1989:3	1989:3	1987:1	1990:3	1998:2	1992:4	1990:2	1989:1	1988:1	1990:3
90% $C.I.$	1989:1-1991:1	1988:2-1990:1	1986:2-1987:2	1989:4-1990:4	1997:1-1998:4	1992:2-1993:3	1989:4-1990:3	1988:3-1989:2	1987:3-1989:2	1990:1-1990:4
T_{B3}		1997:2	1998:3	1996:3			1997:4		1996:2	1998:2
90% $C.I.$		1996:4-1998:3	1997:4-2003:3	1995:4-1997:1			1997:1-1998:3		1995:3-1996:3	1997:4-1998:4
T_{B4}									2001:2	
90% $C.I.$									2000:3-2001:3	

Table 10. Summary of Results using Bai and Perron (1998, 2003)

Province	Regime 1	Regime 2	Regime 3	Regime 4	Regime 5
Alberta	d	C	d		
BC	c	C	C	C	
Manitoba	C	C	C	u	
NB	C	c	c	C	
NF	C	d	C		
NS	C	C	C		
Ontario	d	d	C	C	
PEI	D	C	C		
Quebec	D	C	C	C	u
Saskatchewan	c	C	d	u	

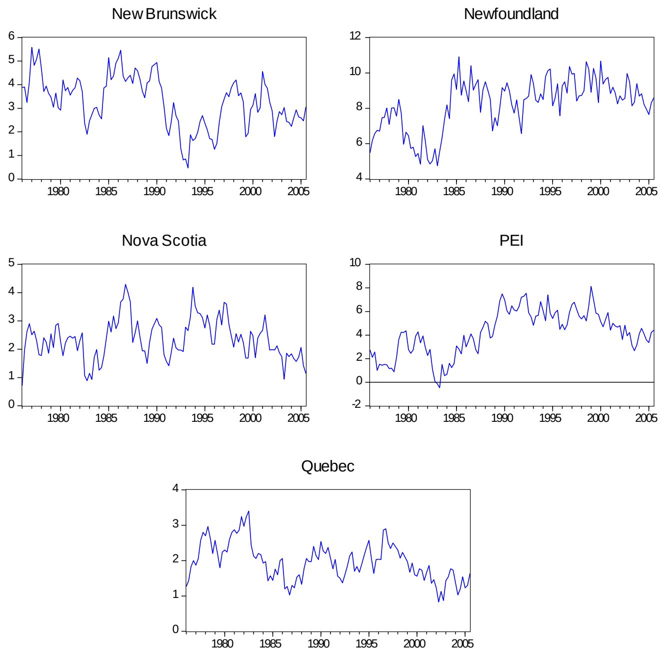

Figure 1. Deviation of the unemployment rates of the Canadian provinces from the unemployment rate of Canada.

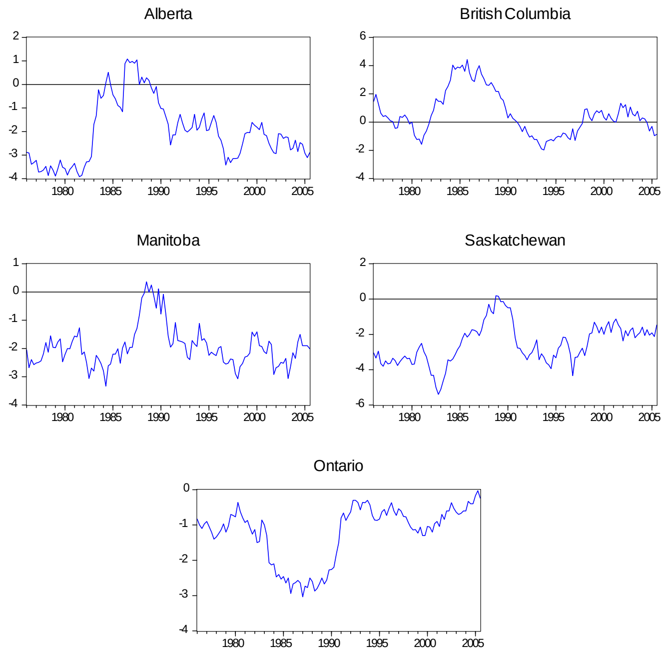

Figure 2. Deviation of the unemployment rates of the Canadian provinces from the unemployment rate of Canada.

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