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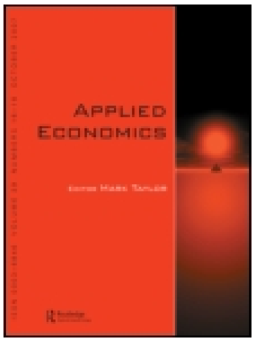

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The impact of trade liberalization and the fiscal equalization transfer policy on provincial income disparities in Canada: an application of GMM estimation

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This article uses the Solow growth model and the panel data method to examine the effect of trade liberalization and the federal equalization transfers on income convergence among Canadian provinces between 1981 and 2006. Estimation problems of weak instruments and endogenous regressors are addressed by the use of a system Generalized Method of Moment (GMM) estimator. The results from the empirical analysis indicate that the current rate of convergence of Personal Income (PI) in Canada is 4.41% per year. This rate is considerably higher than the range of 1.80 and 2.41% per year that previous studies using least-square estimators have reported. The findings from the policy analysis show that the launching and expansion of the North America regional integration have de-accelerated the convergence speed for Canadian provinces by 3.99 and 3.15% per year, respectively. However, consistent with the results from previous studies, the fiscal transfers, which are part of the federal equalization programme, have accelerated the convergence speed for Canadian provinces.

I. Introduction

Regional income disparities have long been the concern of economists and national governments of both developed and developing countries. The recent growth literature, mainly based on the neo-classical growth theory, has focused on the factors that drive growth rate differentials among countries or regions in the same country. Unfortunately, the role economic policies play in the process has received little attention in the literature. This article attempts to fill this gap by using Generalized Method of Moment (GMM) estimators and recent Canadian data set to investigate the effect of trade liberalization and fiscal equalization transfers on regional income disparities in Canada. Specifically, the article seeks to answer two policy questions. First, has the increasing regional integration in North America since the late 1980s led to accelerated or de-accelerated convergence of income among Canadian provinces? Second, have the inter-provincial fiscal transfers, that are part of the overall federal fiscal equalization programme in Canada, helped poorer provinces catch up with the richer ones? To answer these questions, we analyse

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the evolution of provincial disparities of Personal Income (PI) and Personal Income Less Government Transfers (PIT) over the pre- and post-regional integration periods.

According to the neoclassical growth theory, if different regions are at different points relative to their steady state growth paths, then poorer regions will grow faster than the rich ones (the so-called β-convergence). Barro (1991), Barro and Sala-i-Martin (1991, 1992) and Sala-i-Martin (1996) have used cross-section regression analysis to provide empirical evidence of this 'catch-up' effect. They found that convergence occurred for the US States and the regions of Europe and Japan at a rate of about 2% per year. Quah (1993, 1997) criticized the cross-section methodology and offered a more general approach that takes advantage of the dynamic evolution of the distribution of growth rates. This methodology has recently been advanced and refined by Maasoumi *et al*. (2007). Other studies by Ben-David (1993, 1996), Sachs and Warner (1995), Ranjan (2003), Ben-David and Kimhi (2004) and Sohn and Lee (2006) have linked trade to income convergence and concluded that international trade helps accelerate per capita income convergence among countries. The theoretical models used in these studies have suggested three possible explanations for how trade catalyzes income convergence. First is the standard Factor Price Equalization (FPE) theorem. The theorem says that under certain circumstances, if a low-factor price country trades with a high-factor price country, factor prices will eventually equalize. Second is the technology transfer explanation. If countries have different levels of technology, trade could be an important channel through which technology is transferred from a high-technology country to a low-technology country, eventually affecting their factor prices. Third is the endowment of factor quantities explanation which proposes that increased trade in capital goods can affect a county's per capita income by changing its endowment of factor quantities. However, Slaughter (1997, 2001) has demonstrated both theoretically and empirically that these theorems may not necessary hold.

The issue of provincial income disparities in Canada is particularly interesting for three important reasons. First, Canada is a large country, characterized by geographic disparities in resource base and industries. The industrial base of the country is highly concentrated in the Great Lakes-Saint Laurent area, with economic activities in the remaining regions largely based on the exploitation of various natural resources. Second, the Constitution Act of 1982, Section 36 spells out federal responsibilities in the area of provincial disparities and equalization. As a result, the federal and provincial governments have embarked on a comprehensive fiscal transfer programme particularly aimed at bridging the gap between the poor and rich provinces. The programme involves unconditional federal transfers to provinces with low per capita revenues in order to raise their fiscal capacities to a standard level (an average of the five representative provinces: Ontario, Quebec, Manitoba, Saskatchewan and British Columbia). Third, the push for trade liberalization in North America in the early 1980s led to the creation of the Canada–United States Free Trade Agreement (CUSFTA) in 1989. After just 5 years, an expanded free trade area was created with the implementation of the North American Free Trade Agreement (NAFTA), which includes Mexico. Studies by Cox (1995), Kohoe and Kohoe (1995), Wall (2003) and Romalis (2007) have evaluated the effect of CUSFTA and NAFTA on member countries and have concluded that they have led to increase in trade and welfare among the member countries. Depending on the nature of the distribution of these positive benefits, income disparities within a member country could ameliorate or worsen. Another interesting issue related to regional integration and income disparities within a member country is how firms change their location within a country as a response to joining the free trade area.[1]

Studies using Canadian provinces as units of analysis have provided a stock of empirical evidence that establishes the nature of the evolution of provincial income disparities and some evidence on how the federal government policies have affected it. Coulombe and Lee (1995, 1998) and Coulombe (1996, 2000) found that there is convergence among Canadian provinces for different measurements of per capita output and income. According to their empirical analysis, the terms of trade and governments transfers and taxes are two key factors that have helped the provinces to converge at a rate similar to earlier studies on regions within the US, Japan and Europe. Lee and Coulombe (1995) approached the issue differently. They analysed the convergence pattern in earnings, labour productivity and unemployment rates and concluded that there has been provincial convergence in earnings and labour productivity, but not in the unemployment rates. They also found that the per capita measurement of the

[1] See Hanson (1998) and Krugman (1998) for a discussion on the effect of North American integration on firms' optimal location decision.

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variables converged at a much slower rate than the per hour measurement of the variables, reflecting some sustained differences in hours of work and unemployment in the provinces. They then suggested that the changes in the transfer payments that occurred during the period, especially the equalization payments in 1957 and the changes in the unemployment insurance in 1971 had contributed to the reduction of dispersion of PI among the provinces. Wakerly (2002) conducted a different analysis. Using the evolving distribution approach first used by Quah (1993), and data on provinces and industries, he found that the growth process in Canada has not reduced income disparities among the provinces.

Few studies have discussed the role international trade plays in income convergence in Canada. Though Gunderson (1996) discussed the effect of trade on convergence in Canada, it was Coulombe (2003) who first conducted an empirical analysis of the effect of trade on convergence among Canadian province. In his empirical analysis, he included both interprovincial and international trade variables in the convergence equation and concluded that international trade has a long-run positive effect on Gross Domestic Product (GDP) per capita and employment, but interprovincial trade has only a long-run positive effect on employment.

All the studies discussed above and others such as Coulombe (1997) and Coulombe and Day (1999) used least-square estimators for their estimate of the convergence parameter. These estimation techniques have some potential econometric problems which may lead to biased estimates of the convergence parameter. For instance, the techniques pay little or no attention to the problem of unobserved province specific effects which has considerable implications with regards to the estimation of unbiased convergence rate. In addition, the techniques neglect the potential problem of endogeneity of regressors in the growth equation which may give rise to dynamic bias problems. The panel data analysis by using both cross-sectional and time series variability is well-equipped to deal with the above problems. Ralhan and Dayanandan (2005) is the only Canadian study that used panel data and GMM estimation techniques to examine the issue of income convergence in Canada. They used random effect, fixed effect, and Difference GMM (DIFF-GMM) estimators and concluded that Canadian provinces converged at a rate of about 6.5% per year. Unfortunately their analysis did not examine the validity of instruments used in the DIFF-GMM estimation and did not empirically examine the finite sample performance of the DIFF-GMM estimator.

The goal of this article, therefore, is to examine in a comprehensive and transparent manner, the speed of income convergence among Canadian provinces, and how the regional trade liberalization in North America and the government fiscal transfers have affected it. In contrast to earlier Canadian studies, we use the Arellano and Bond (1991) and Blundell and Bond (1998) linear GMM estimators, which address the potential endogeneity of the regressors, and incorporate, albeit implicitly, fixed effects. To evaluate the finite sample performance of the GMM estimators, we use the Ordinary Least Squares (OLS) and the Within Group (WG) estimators to establish an upper and lower bound for the convergence parameter. This article contributes to the literature on provincial income disparities in Canada in two important aspects. The first relates to the econometric techniques used. Unlike earlier studies, we use the DIFF-GMM and the System GMM (SYS-GMM) estimators which allow us to eliminate unobserved province specific effects and correct for the problem of endogenous regressors. Second, it is the first study that analyses how the benefits of the increasing North American integration are distributed among Canadian provinces within the convergence framework.

Our findings are relatively easy to report. First, we find that the SYS-GMM estimator yields the best results, in terms of finite sample performance, because it addresses the estimation problems of weak instruments and endogenous regressors. Second, the launching and expansion of the North America regional integration have de-accelerated the convergence speed for Canadian provinces by 3.99 and 3.15% per year respectively. Third, federal government transfer, which is part of the overall equalization programme in Canada, has accelerated the speed of convergence in income in Canada.

The rest of this article is organized as follows. Section II presents an overview of the Solow neoclassical growth model. Section III discusses the DIFF-GMM and SYS-GMM estimators in conjunction with the problem of endogenous regressors and weak instruments. This section also discusses problems associated with applying least-square estimators to panel data set. Section IV conducts the empirical analysis of the effect of regional trade liberalization and the fiscal equalization transfers on conditional income convergence among Canadian provinces. Section V concludes this article.

II. The Solow Growth Framework

The Solow growth model provides the theoretical basis for a large number of studies on income convergence (Barro and Sala-i-Martin, 1992;

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Mankiw *et al.*, 1992; Quah, 1993; Coulombe and Lee, 1994, 1995; Islam, 1995; Coulombe and Tremblay, 2001; Weeks and Yao, 2003). Although most of these studies have used either pooled time series and cross-section or simple cross-sectional data approach to estimate the convergence rate, Islam (1995) and Weeks and Yao (2003) have provided a good background to a panel data approach to the convergence hypothesis. In this section, we use a summarized version of their models. Using standard notation, we assume a Cobb–Douglas production function with labour augmenting technological process

$$Y(t) = K(t)^{\alpha}[A(t)L(t)]^{1-\alpha} \tag{1}$$

where $0 < \alpha < 1$, Y is output, K is capital, L is labour and A is the level of technology. Labour force and technology are assumed to grow exogenously at the rate n and g, respectively

$$L(t) = L(0)e^{nt} \tag{2}$$

$$A(t) = A(0)e^{gt} \tag{3}$$

Define $\hat{y} = Y/AL$, $\hat{k} = K/AL$, δ as a constant rate of depreciation, and s as a constant fraction of output that is saved and invested. Then the dynamic equation for $\dot{\hat{k}}$ is given by

$$\dot{\hat{k}}(t) = s\hat{y}(t) - (n+g+\delta)\hat{k}(t) \tag{4}$$

$$= s\hat{k}(t)^{\alpha} - (n+g+\delta)\hat{k}(t) \tag{4$'$}$$

From Equation 4′ $\hat{k}$ converges to its steady state value

$$\hat{k}^{*} = \left(\frac{s}{n+g+\delta}\right)^{\frac{1}{(1-\alpha)}} \tag{5}$$

Substituting Equation 5 into Equation 1 and taking logs, the steady state income per capita is

$$\ln\left[\frac{Y(t)}{L(t)}\right] = \ln A(0) + gt + \frac{\alpha}{1-\alpha}\ln(s) - \frac{\alpha}{1-\alpha}\ln(n+g+\delta) \tag{6}$$

Most conditional convergence studies, such as Mankiw *et al.* (1992), Barro and Sala-i-Martin (1995) and Coulombe (2000, 2003) have paid little attention to $\ln A(0)$ and gt. Specifically, they have almost invariably relegated them into the error or the constant terms of their regression models with the assumption that they are independent of the s and $(n+g+\delta)$ variables. Our main argument is that a panel data framework that explicitly controls for the technological shift term $\ln A(0)$ is the appropriate approach. Following Weeks and Yao (2003), we write an autoregressive form of the growth model (Equation 6) as

$$\ln y(t_2) = \zeta \ln y(t_1) + (1-\zeta)\ln A(0) + g(t_2 - \zeta t_1) + (1-\zeta)\frac{\alpha}{1-\alpha}\ln s - (1-\zeta)\frac{\alpha}{1-\alpha}\ln(n+g+\delta) \tag{7}$$

where $y(t) = Y(t)/L(t)$ is the per capita income and $\zeta = e^{-\beta_1}(t_2 - t_1)$. Equation 7 represents the transitional growth dynamics of an economy towards its steady state income path and represents the general dynamic framework within which income convergence is examined. The equation is a dynamic panel data model with $(1-\zeta)\ln A(0)$ as the time invariant individual unit/regional effect term and the $g(t_2 - t_1)$ as the time specific effect. Using standard notation of the panel data literature and adding a disturbance term we may re-write Equation 7 as

$$y_{it} = \gamma y_{it-1} + \sum_{j=2}^{3}\beta_j x_{it}^{j} + \eta_t + \mu_i + \nu_{it} \tag{8}$$

where $y_{it} = \ln y(t_2)$, $y_{it-1} = \ln y(t_1)$, $\gamma = e^{-\beta_1}(t_2 - t_1) = \zeta$, β_1 measures the rate of convergence, $\beta_2 = (1-\zeta)\alpha/(1-\alpha)$, $\beta_3 = -(1-\zeta)\alpha/(1-\alpha)$, $x_{it}^{1} = \ln(s)$, $x_{it}^{2} = \ln(n+g+\delta)$, $\mu_i = (1-\zeta)\ln A(0)$, $\eta_t = g(t_2 - t_1)$, and ν_{it} is the usual transitory error term that varies across units/regions and time periods and has mean equal zero.

In our empirical analysis, we will allow provinces to have differences in the initial state of technology $A(0)$, and assume that g(technological growth rate) is homogeneous across provinces. Hence Equation 8 becomes[2]:

$$y_{it} = \gamma y_{it-1} + \sum_{j=2}^{3}\beta_j x_{it}^{j} + \mu_i + \nu_{it} \tag{9}$$

It must be recalled that one of our main goal is to investigate the effect of regional integration on income convergence among provinces in Canada. We therefore need to distinguish the conventional income convergence from the income convergence driven by regional integration. This will allow us to investigate how the steady state varies due to the nature of the distribution of the benefits from CUSFTA and NAFTA. To capture the effect, an interactive dummy variable (Dy_{it-1}) is added to Equation 9 as follows:

$$y_{it} = \gamma y_{it-1} + \theta D y_{it-1} + \sum_{j=2}^{3}\beta_j x_{it}^{j} + \mu_i + \nu_{it} \tag{10}$$

[2] The $A(0)$ term reflects not just technology but resources endowments, climate and institutions. As mentioned earlier these factors differ across the Canadian provinces.

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where $D=0$ for all t before the formation of the regional integration, and equals 1 if the regional integration is formed in year t. The coefficient on the interactive term, θ, measures the accelerating/de-accelerating effect of the regional integration, while γ measures the conventional convergence. We use γ to calculate the rate of convergence during the pre-integration period, and use $\gamma+\theta$ to calculate the rate of convergence during the post-integration period. If the coefficient θ is negative in sign and statistically significant, then the regional integration, as a trade policy, has resulted in accelerated convergence among Canadian provinces. On the other hand, if the coefficient θ is positive in sign and statistically significant, then the regional integration has reduced the rate at which Canadian provinces were converging before the trade policy. With respect to the effect of the fiscal transfers on the speed of convergence, we will follow Coulombe and Lee (1995) by estimating the convergence equation for both PI and PIT. If the transfers are truly equalizing, then the convergence rate for PI must be greater than PIT.

III. Panel Data Estimation Issues and Data

Earlier studies on income convergence in Canada have mostly used least-square estimators. These studies have paid little or no attention to the problem of the unobserved province specific initial technology levels. This has considerable implications for the estimation of the unbiased convergence rate. It is also well-known in the convergence literature that y_{it-1} (the lagged dependent variable) is endogenous to the fixed effect μ_i. Hence, OLS estimation of Equation 10 without the fixed effect gives rise to dynamic panel bias.[3] This is because the lagged dependent variable is positively related to the fixed effect which violates an assumption necessarily for the consistency of OLS. In particular, OLS inflates the coefficient estimate of the lagged dependent variable by attributing predictive power to it that actually belongs to the fixed effect.[4] One of the nonpanel data approach to solving the endogeneity problem is the WG estimator. The technique partials the cross-section fixed effect from the data by applying a mean deviation transform to each variable, when the mean is calculated at the cross-section unit level. However, this approach does not eliminate the 'dynamic panel bias'. According to Bond (2002) the lagged dependent variable under the WG estimator becomes $y^*_{it-1}=y_{it-1}-1/T-1(y_{i2}+\cdots+y_{iT})$ while the error term becomes $v^*_{it}=v_{it}-1/T-1(v_{i2}+\cdots+v_{iT})$. The problem is that the y_{it-1} term in y^*_{it-1} correlates negatively with $-1/T-1v_{it-1}$ in v^*_{it}. Hence, the coefficient on the lagged dependent variable will be biased downwards. Finally, in the case of conditional convergence studies, the problem of potential endogeneity of other variables, such as savings/investment rate is also neglected.

The panel data analysis by using both cross-sectional and time series variability is well-equipped to deal with the above problems. One such method is the DIFF-GMM. As suggested by Arellano and Bond (1991), the endogeneity problem of the lagged dependent variable can be corrected by first differencing the data and under the assumption of serially uncorrelated level residuals, the second and third untransformed lags are used as instruments.[5] This implies the following moment condition $E(y_{it}\Delta\mu_{it})=0$ for all $t=3,\ldots,T$. At the same time differencing the data addresses the problem of unobserved fixed effect. Applying the transformation to Equation 10 gives

$$\Delta y_{it}=\gamma\Delta y_{it-1}+\theta D\Delta y_{it-1}+\sum_{j=2}^{3}\beta_j\Delta x^j_{it}+\Delta v_{it} \quad (11)$$

Though the fixed effect is expunged, and the endogeneity problem is solved by first differencing the data, Blundell and Bond (1998) demonstrate that if y is persistent (close to random walk) then DIFF-GMM performs poorly because past levels convey little information about future changes. Hence, untransformed lags are weak instruments for transformed variables. This is referred to as the 'weak instrument problem' of the DIFF-GMM estimator.[6] The SYS-GMM developed by Blundell and Bond (1998) addresses the weak instrument problem of DIFF-GMM.[7] The approach comprises two equations. The first is the usual DIFF-GMM which uses lagged levels as instruments for equations in first differences. In the second equation, instead of

[3] This is also known in the literature as the endogeneity problem.
[4] Hsiao (1986), Sevestre and Trongnon (1996) and Weeks and Yao (2003) make the same argument.
[5] Another form of transformation known as 'forward orthogonal deviation' or 'orthogonal deviation' is commonly used.
[6] Blundell and Bond (1998) used Monte Carlo simulation to demonstrate that the weak instrument problem can result in large finite-sample biases when using DIFF-GMM estimator to estimate autoregressive models with relatively short panels.
[7] Blundell and Bond (1998) have also demonstrated that under a random effect model, the DIFF-GMM estimator can suffer from serious efficiency losses. This is because there are potential informative moment conditions that are ignored in the DIFF-GMM approach.

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differencing the data to expunge the fixed effect, it takes the first difference of the variables to make them exogenous to the fixed effect and use them as instruments in the level equation. This amounts to adding another moment condition, $[E(\Delta w_{it}\mu_i) = 0]$, for all i and t, where Δw_{it} is the instrument and μ_i is the fixed effect. By exploring more moment conditions, the SYS-GMM estimator is more efficient asymptotically and in finite sample properties than the DIFF-GMM estimator that uses only a subset of linear moment conditions. The efficiency gain from imposing the level moment condition comes with some potential problems with the SYS-GMM estimator. There is the need for additional assumptions, which if not satisfied will lead to bias in the estimates. For instance, if the unit specific effects are correlated, then some of the level moments conditions will not be valid and the SYS-GMM estimates will be inconsistent. It is, therefore, important to conduct specification test to justify the use of additional level moment conditions. This can be evaluated by the Sargan-difference test for instruments validity. We also used the Stock and Yogo (2001) approach to test for the presence of weak instruments.[8] To evaluate the finite sample performance of both the DIFF-GMM and SYS-GMM estimators, we used the OLS and the WG estimators to establish an upper and a lower bound for the autoregressive parameter (y_{it-1}).

The data set

All the data used for the study were obtained from the online database of Statistics Canada: CANSIMM II. The period of analysis is from 1981 to 2006. Since we want to determine the effects of regional integration and federal fiscal transfers on provincial income convergence, we used annual data on per capita real PI, with and without government transfers for all the 10 provinces. Other variables for which we collected data are: real investment, labour force growth for working age population (between 15 and 65 years).[9] To render our results comparable with other Canadian studies, we defined all our variables as relative to the Canadian average.[10] By adopting this approach, our results are less likely to be influenced by business cycle fluctuation. For regional integration, we tested the effect of the CUSFTA which was formed in 1989 between Canada and the US, and was later deepened and expanded to include Mexico in 1994 to form the NAFTA. Hence, the CUSFTA dummy takes the value '0' for all periods before 1989 and '1' otherwise. Similarly, the NAFTA dummy equals '0' for all periods before 1994 and '1' otherwise.

IV. Empirical Investigation and Results

The regression results of the test of the effect of the regional integration and the federal government transfer systems on income convergence in Canada are reported in Tables 1–4. As discussed in the previous section, the OLS level estimators, by omitting the unobserved unit-specific effects in a dynamic panel data model, yield estimates that are biased upwards and inconsistent due to the positive correlation between the lagged dependent variable (y_{it-1}) and the fixed effects (μ_i). On the other hand, the WG estimator produces a downward bias with the extent of attenuation increasing when exogenous variables are added to the model. In this section, we used the OLS and the WG estimators to establish the upper–lower bounds for the coefficient of the lagged dependent variable. A good estimate of the true parameter should therefore fall within the range established by these two estimators. Hence, we used the estimates from the OLS and WG estimators to judge the unbiasedness of the DIFF-GMM and SYS-GMM estimators. Columns 1 and 2 of the tables report, respectively, the results of the OLS and WG estimators. Columns 3 and 4 report, respectively, the parameter estimates using the DIFF-GMM and SYS-GMM estimators. In all the regressions, we used the coefficient of the interactive dummy variable to test for the effect of the regional integration on the speed of convergence. As earlier mentioned, a negative coefficient indicates an accelerating convergence and a positive coefficient implies a de-accelerating convergence.

The coefficients on the conditional variables ($(n+g+\delta)$, I/GDP) are mostly not significant and carried the wrong signs. Table 1 reports results from PIT. The estimated coefficients for y_{it-1} for all four estimation methods are very significant. The implied speed of convergence from the OLS result is 3.47% which is significantly lower than the 11.08% speed of convergence implied by the WG estimation. The parameter estimate of the y_{it-1} for the

[8] According to the Stock and Yogo (2001) procedure, a group of instruments is weak if the bias of the Instrumental Variable (IV) estimator relative to the bias of the OLS could exceed a certain threshold, say 10%. The null of weak instrument is rejected if the calculated test statistic ($g_{\min}$) is greater than the critical value.
[9] The exact definitions of variables are available from the author on request.
[10] Specifically, the provincial economic variables Q_{it} (like y_{it}, x_{it}) are measured as the logarithmic deviation from the cross-sectional mean at time t. $Q_{it} = \log(Q_{it}/\sum_{i=1}^{N}\frac{1}{N}Q_{it})$.

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Table 1. Panel test for conditional convergence of PIT: CUSFTA

	1	2	3	4
	OLS	WG	DIFF-GMM	SYS-GMM
$\ln(y_{t-1})$	0.9653***	0.8892***	0.8704***	0.9259***
	(0.0124)	(0.0327)	(0.0442)	(0.0165)
$\ln(n+g+\delta)$	0.0022	0.0039	0.0269	0.0157
	(0.0156)	(0.0186)	(0.0227)	(0.1564)
$\ln(I/\text{GDP})$	0.0098	0.0032	−0.0415	0.0148
	(0.0097)	(0.0164)	(0.0279)	(0.0116)
CUSFTA $*\ln(y_{t-1})$	−0.0019	−0.0276	0.0080	0.0314*
	(0.0157)	(0.0189)	(0.0233)	(0.0174)
Constant	−0.0056	−0.0221	NA	−0.0297
	(0.0284)	(0.0341)		(0.0283)
Sargan test (p-value)			0.024	0.188
$g_{\min}$			8.16	14.73
Number of observations	250	250	200	210

Notes: SEs in parentheses. All reported SEs are corrected for heteroscedasticity. The parameter estimates and the SEs reported from the GMM are one-step estimators. The figure reported for the Sargan test is the p-value of the null hypothesis of valid instruments. The $g_{\min}$ is calculated using three instruments for the DIFF-GMM and six instruments for the SYS-GMM. The desired maximum bias of the IV estimator relative to OLS is 10%.
*** and * denote significance at 1 and 10% levels, respectively.

Table 2. Panel test for conditional convergence of PI: CUSFTA

	1	2	3	4
	OLS	WG	DIFF-GMM	SYS-GMM
$\ln(y_{t-1})$	0.9520***	0.8767***	0.8566***	0.9058***
	(0.0132)	(0.0309)	(0.0432)	(0.0179)
$\ln(n+g+\delta)$	0.0060	0.0095	0.0291	0.0186
	(0.0129)	(0.0153)	(0.0187)	(0.0130)
$\ln(I/\text{GDP})$	0.0061	0.0032	−0.0371	0.0078**
	(0.0081)	(0.1366)	(0.0230)	(0.0036)
CUSFTA $*\ln(y_{t-1})$	0.0067	−0.0315	0.0128	0.0399**
	(0.0176)	(0.0214)	(0.0262)	(0.0193)
Constant	−0.0123	−0.0289	NA	−0.0348
	(0.0235)	(0.0275)		(0.0236)
Sargan test (p-value)			0.035	0.111
$g_{\min}$			7.69	14.33
Number of observations	250	250	200	210

Notes: Refer footnote of Table 1.
*** and ** denote significance at 1 and 5% levels, respectively.

DIFF-GMM estimator, though significant, falls out of the upper and lower bound (0.9653−0.8892) established by the OLS and the WG estimators. This weak performance of the DIFF-GMM estimator is likely due to the weakness of the instrument set. The Stock and Yogo (2001) test procedure for weak instruments was conducted. The displayed statistics of 8.16 is less than the critical value of 9.8 at the 5% significance level, indicating that the instruments for the lagged dependent variable for the DIFF-GMM estimator are weak. The p-value of the Sargan test also suggests that the moment conditions for the DIFF-GMM estimator are not valid. The convergence coefficient of (y_{it-1}) for the SYS-GMM estimator (0.9259) is significant and falls between the upper and lower bounds (0.9653−0.8892).[11] Therefore, the estimator is likely to be unbiased. The test statistics of the test for weak instruments (14.73) is greater than the critical value of 11.12 at the 5% significant level. The Sargan test

[11] It is important to note that these numbers are themselves point estimates with associated confidence intervals.

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Table 3. Panel test for conditional convergence of PIT: NAFTA

	1	2	3	4
	OLS	WG	DIFF-GMM	SYS-GMM
$\ln(y_{t-1})$	0.9655***	0.8919***	0.8313***	0.9417***
	(0.0116)	(0.0351)	(0.0529)	(0.0135)
$\ln(n+g+\delta)$	−0.0017	0.0041	0.0468	0.0157
	(0.0165)	(0.0208)	(0.0277)	(0.0172)
$\ln(I/\text{GDP})$	0.0097	0.0016	−0.0485	0.0198*
	(0.0097)	(0.0166)	(0.0385)	(0.0114)
$\text{NAFTA} * \ln(y_{t-1})$	0.0089	−0.0198	−0.0267	0.0264**
	(0.0168)	(0.0211)	(0.0248)	(0.0139)
Constant	−0.0016	−0.0208	NA	−0.0298
	(0.0301)	(0.0386)		(0.0313)
Sargan test (p-value)			0.048	0.146
$g_{\min}$			8.08	14.62
Number of observations	250	250	200	210

Notes: Refer footnote of Table 1.
***, ** and * denote significance at 1, 5 and 10% levels, respectively.

Table 4. Panel test for conditional convergence of PI: NAFTA

	1	2	3	4
	OLS	WG	DIFF-GMM	SYS-GMM
$\ln(y_{t-1})$	0.9534***	0.8874***	0.8327***	0.9244***
	(0.0125)	(0.0332)	(0.0525)	(0.0149)
$\ln(n+g+\delta)$	−0.0007	0.0058	0.0378	0.0148
	(0.0137)	(0.0169)	(0.0223)	(0.0144)
$\ln(I/\text{GDP})$	0.0060	0.0042	−0.0393	0.0074**
	(0.0081)	(0.0138)	(0.0273)	(0.0036)
$\text{NAFTA} * \ln(y_{t-1})$	0.0228	−0.0119	−0.0153	0.0315**
	(0.0193)	(0.0241)	(0.0288)	(0.0153)
Constant	−0.0001	−0.0198	NA	−0.0280
	(0.0248)	(0.0308)		(0.0261)
Sargan test (p-value)			0.058	0.146
$g_{\min}$			7.82	13.76
Number of observations	250	250	200	210

Notes: Refer footnote of Table 1.
*** and ** denote significance at 1 and 5% levels, respectively.

(p-value $= 0.18$) suggests that the moment conditions of the SYS-GMM are valid. In addition, the interactive regional integration dummy ($\text{CUSFTA} * \ln y_{it-1}$) is positive and significant at the 10% level. This implies that the formation of the CUSFTA between the US and Canada in 1989 reduced the speed of convergence among Canadian provinces. That is, rich provinces benefited more from CUSFTA than the not-so-rich provinces. The results suggest that the rates of convergence of PIT were 7.41 and 4.27%, respectively, for the period before and after the formation of the regional integration.

Table 2 reports the results for per capita PI. The pattern of the parameter estimates is similar to those of Table 1. The convergence parameter for the DIFF-GMM estimator falls out of the upper–lower bounds (0.9520–0.8767) established by the OLS and the WG estimators. The weak instruments and the Sargan tests indicate that the instruments used in the DIFF-GMM are weak and the moment conditions are not valid. On the other hand, the SYS-GMM estimate falls within the bound. The null hypothesis of weak instruments is rejected and according to the Sargan test, the instruments used are valid. The coefficient of the regional integration dummy and the investment rate variables are positive and significant only for the SYS-GMM estimator. The implied convergence rate for the estimator is 9.42% before

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the integration and 5.43% after the integration, suggesting that the regional integration benefited the rich provinces more than it did for the not-so-rich provinces. These results present strong evidence that the launching of CUSFTA de-accelerated income convergence among Canadian provinces. On the other hand, the federal government fiscal transfer programme has led to acceleration of income convergence in Canada. The convergence rate of PI is 9.42% and that of PIT is 7.41%. These rates are for the period before the formation of CUSFTA. The corresponding rates after CUSFTA are 5.43 and 4.27, respectively. Whereas the results of the effect of the regional integration is new in the Canadian literature, the results in terms of the income transfer programme is similar to previous findings of Coulombe and Lee (1995, 1998) and Coulombe (2000), which were obtained using least-square estimators.

We next analysed the effect of the expansion of the regional integration on income convergence in Canada, by interacting a NAFTA dummy with the lagged dependent variable. A negative and significant coefficient of the interactive variable will suggest that the expansion of the regional integration to include Mexico may have presented better opportunities to the not-so-rich provinces that have since been benefiting more from the expansion than the rich one. A positive coefficient will imply that the factors that make the rich provinces benefit more from CUSFTA are still in place after the expansion of the regional integration. Table 3 reports results for PIT before and after NAFTA. As before, the SYS-GMM estimator is preferred to the DIFF-GMM estimator. Whereas the convergence parameter for the DIFF-GMM estimator (0.8313) falls outside the upper–lower bounds (0.8919–0.9655), the parameter for the SYS-GMM estimator (0.9417) falls within the bound. The Sargan and the weak instruments tests also indicate that the instruments used in the estimation are only valid and strong for the SYS-GMM estimator. The coefficient of the interactive integration variable is negative for the WG and DIFF-GMM estimators, though not significant. On the other hand, the coefficient is positive for both the OLS and SYS-GMM estimators, but significant at the 5% level only for the SYS-GMM estimator. Hence the implied rates of convergence before and after the expansion of the regional integration are 5.83 and 3.19%, respectively. The coefficient of the investment rate is significant and has the right sign only for the SYS-GMM estimator.

Finally, Table 4 reports results for PI before and after the expansion of the regional integration. The results followed the same pattern as those discussed previously. The coefficient of the lagged dependent variable for the SYS-GMM estimator (0.9244) falls within the upper–lower bounds (0.9534–0.8874) established by the OLS and WG estimators. The *p*-value of the Sargan test and the displayed statistics for the test of weak instruments suggest that the moment conditions are valid and the instruments are strong only for the SYS-GMM. The implied rate of convergence for PI is 7.56% per year. Compared to the convergence rate of 5.83% (reported in Table 3) for PIT, we can once again conclude that the government transfer programme has accelerated convergence among Canadian provinces. The interactive NAFTA coefficient is positive and significant at the 5% level, implying a convergence rate of income with government transfers of 4.41% after the expansion of the regional integration. The coefficient of the investment rate is positive and significant at the 5% level.

Some key results emerge from the estimates presented in Tables 1–4. First, the general insignificance of the coefficient of the population growth variable may imply that inter-provincial differences in productivity and investment rates instead of population growth rate are likely to play an important role in explaining differences in provincial income. Second, the SYS-GMM is the preferred estimator for the convergence equation. Third, the launching and the expansion of the regional integration in North America have reduced the convergence rate of income among Canadian provinces. Finally, consistent with results from earlier studies, the fiscal equalization programme that involves transfers to individuals appears to reduce provincial income disparities substantially.

V. Conclusion

This article extends the analysis of previous studies on provincial income convergence in Canada in two important directions. First, it uses a new methodology which allows us to correct for endogeneity of right-hand side variables and incorporate, albeit implicitly, provincial fixed effects. Second, it is the first study that investigates the effect of regional integration on income convergence among Canadian provinces. Using results from an application of OLS and WG estimators, the system GMM estimator is shown to be the preferred estimation method, in terms of providing consistent and more efficient estimates of the convergence rate.

The central conclusions of this article are as follows. First, the current rate of convergence of per capita PI is 4.41% per year. This is considerably higher than the range 1.8–2.4% per year that previous studies using least-square estimators have reported. The result is also consistent with the claim by Islam (1995) that the speed of convergence parameter based upon panel data studies has in general been considerably higher than the average of 2% reported by the cross-sectional studies. Second, the launching and expansion of the regional integration in North America has reduced the speed of convergence for the Canadian provinces by 3.99 and 3.15% per year, respectively. This reduced convergence is evidently temporal since the provinces may ultimately adjust to the opportunities presented by the regional integration. Also, over time, capital and labour will move faster to provinces with favourable economic environment leading to adjustments in factor prices and ultimately enabling the provinces to converge at a rate faster than the pre-integration rate. National policies should focus on removing remaining bottlenecks that impede capital and labour mobilities. The national government can also hasten the process by addressing some of the concerns of the provinces about the nature of the regional integration. Third, the federal fiscal transfers, which are part of the overall equalization programme, have accelerated the speed of convergence of income in Canada. This implies that the transfers have succeeded in making poorer provinces catch up with the richer ones as earlier studies, such as Coulombe and Lee (1995) and Coulombe (1996) have already demonstrated.

From a policy point of view, this article has highlighted the relative importance of interprovincial fiscal redistribution and regional trade liberalization in the convergence process among Canadian provinces. The methodology used and the results reported have provided sound basis for further research. In the empirical application of the GMM estimators, we assumed that technology progress is homogeneous among Canadian provinces. This is an empirical issue that needs to be tested explicitly. Theoretical trade models have demonstrated that differences in productivity and technological progress play important role in the distribution of the benefits of trade. Therefore, future research should test for the existence of heterogeneity in the rate of technology process and the role it plays in the relationship between trade liberalization and income disparities in Canada. Another direction for future research is to extend the coverage of the study to include all 50 states of the US and 31 states in Mexico.

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