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Abstract

Migration between the Canadian provinces generally followed a declining trend over the period 1971-2004. In this paper, taking Ontario a case study, we seek to explain these patterns using recent panel cointegration methods that are robust to cross-section dependence. Estimation of heterogenous models suggests that the determinants of migration vary across provinces. Overall, unemployment differential and income in the sending province appear to be the most important ones, with income and federal transfer differentials playing only a minor role.

Keywords: Internal migration; panel cointegration; bootstrap; Canada.
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1 Introduction

Canada has a long tradition of internal migration stretching from 1950s to the present day. Research over the last four decades on the spatial pattern of migration in Canada brings

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out the following stylized facts. Throughout this long period three economically strong provinces–Ontario, British Columbia, and Alberta–have been the principal net gainers from interprovincial migration; whereas the remaining seven economically weak provinces tended to be the consistent net losers through internal migration. Stronger economic growth and better employment opportunities are often taken to be the explanation of the regional pattern of migration between the two sets of provinces.[1] Persistent change in internal population redistribution over the 1951-2001 period has helped Ontario, British Columbia, and Alberta to increase their share of Canadian population by an average of 4.6%, while the average share for the remaining seven provinces decreased by nearly 3.3% (Liaw and Xu, 2005). In the last three decades this trend somehow has slowed down, so that the 1971-2004 averages are respectively 2.8% and -1.2% (see Table 1).

Since the impact of internal migration on the distribution of population across provinces has been substantial[2] it is not surprising that there is by now a fairly substantial literature on interprovincial migration based on both micro and aggregate data.[3]

Although internal migration has been an integral part of Canada's population dynamics, interestingly very little has changed in terms of specific migration flows among the provincial origins and destinations. More than 30 years later we see that the spatial pattern of migration in Canada is very much in line with the observation first made by Stone (1974). In other words, Ontario still is the most favorite destination for out-migrants from Quebec and the four Atlantic[4] provinces. On the other hand, out-migrants from the western provinces (Alberta and British Columbia) mainly choose other western provinces as their most favored destinations. Importantly, there are no large streams originating in either British Columbia or Alberta and ending east of Ontario; whereas Ontario ranked second as a destination for out-migrants from the two westerly provinces. In this respect, Ontario is like a sort of "buffer zone" inhibiting strong flows between the eastern and western regions of Canada (see Stone, 1969, 1974).

Nevertheless, an important observation that, to the best of our knowledge, has been overlooked in the literature is the fact that gross migration flows from both eastern and

[1] See, for instance, Table 1 in Liaw and Xu (2005).

[2] The annual average (between 1971-2004) migration flow from one province to another is approximately 31,279 persons, or 1.17% of the total population (excluding the territories).

[3] For review of pre-1975 work see Stone (1974). Day and Winner (1994) extensively cover pre-1995 literature, while Finnie (2004) offers a long list of related work on internal migration.

[4] The four Atlantic provinces, all located east of Ontario, are Newfoundland and Labrador, Prince Edward Island, Nova Scotia, and New Brunswick.

western provinces to Ontario have been declining steadily over the years (see Figure 1). Save for British Columbia, the logarithm of gross migration flows to Ontario is mostly trending downward, with only some temporary upward swings around the early 1990s. This is interesting, as Ontario has consistently been one of the most attractive provinces in terms of both income and employment opportunities. What is the cause of the declining trend in migration flows to Ontario from rest of Canada, particularly the eastern provinces? Only the narrowing of income gap and unemployment differentials, or have other factors been at play?

Our goal in this article is to empirically examine the long-run determinants of this decline using a panel of interprovincial migration data from 1971-2004. Obviously, there have been several attempts to explain interprovincial migration pattern using long-span data. Courchene (1970) uses family allowance data over 1952-1967 and observes a positive association between rates of migration flows and ratios of earned income per employed person between Canadian provinces. Vanderkamp (1971) observes a stronger positive association between in-migration and earnings than a negative association between out-migration and earnings for Canadian interprovincial flows during the period 1947-1966. Shaw (1986) finds that fiscal variables (e.g. unemployment insurance) matter more over traditional market variables (e.g. job creation or wages) for Canadian migration during the period of 1956-1981. Day and Winer (2006) use individual tax records to construct in-migration and out-migration data over the period 1974-1996 and find that regional differences in (fiscal) policy variables do not have significant influence on interprovincial migration. Finally, both Day and Winer (2006) and Coulombe (2006) point the importance of structural factors such as regional differences in earnings, employment prospects, and labor productivity as the main drivers of interprovincial migration in Canada.

However, all these empirical studies share a serious methodological weakness: the problem of non-stationarity of the data is never taken into account. In fact, the issue of non-stationarity seems to be largely ignored even in the most recent contributions to the empirical literature on interprovincial migration in Canada (e.g. Day and Winer, 2006; and Coulombe, 2006). Therefore, the objective of this paper is to fill the gap in the literature by applying recent advances in panel data econometrics to examine the long-run determinants of interprovincial migrations in Canada over the period 1971-2004. Although we focus our attention to the spatial migration patterns in Ontario, our framework can clearly be used to study the migration behavior in other provinces.

Our paper is closely related to the strand of literature that examines regional disparity in income across Canadian provinces. As such, convergence in income is likely to reduce the economic motivation for migration, while absence or lack of convergence, on the opposite, is likely to trigger migration from poorer to the richer provinces. Using cross-sectional methods Coulombe and Lee (1995) and Coulombe and Day (1999) show that poor provinces tend to grow faster than rich ones in per capita terms during the period 1961-1991. Based on unweighted standard deviation of relative income across Canadian provinces Helliwell (1996) concludes that there was no significant differences in the rate of income convergence between 1926-1960 and 1960-1990. Afxentiou and Serletis (1998) applying cointegration approach are unable to find any support for convergence from 1961 to 1991. By contrast, DeJuan and Tomljanovich (2005) find overwhelming evidence of economic convergence for most part of Canada over the period 1926 to 1996. In particular, they observe very strong convergence evidence particularly in the Atlantic and Plains[5] provinces.

In light of these findings it seems that while complete convergence (equilibrium) for all provinces may not been achieved, there seem to be a rather general consensus on the existence of a process of convergence. The increased efforts by the federal government to establish interprovincial redistribution programs (i.e. equalization payments) may have facilitated such convergence process (e.g. Kaufman et al., 2003; Rodriguez, 2006). Therefore, one could argue that a possible explanation for the steady decline in migration flows into Ontario is the increasing disappearance of regional income disparity within Canada. We agree to this view and incorporate these ideas in our analysis.

The rest of the paper is organized as follows. Section 2 describes data and empirical model. Section 3 briefly reviews the econometric methodology used. Section 4 presents the empirical results. Section 5 concludes. All additional materials are presented in the Appendix.

[5] Manitoba and Saskatchewan.

2 Modeling Internal Migrations: Data and Models

The primary source of our data is Statistics Canada[6], which records annual migration streams by province of origin and destination. The advantage of using migration data with a *defined origin* and *defined destination* is that it will allow us to see how the provinces of origin react to economic fundamentals when choosing Ontario as their favorite destination. In particular, we will be able to ascertain why one region respond to more of a certain kind of attributes than another. The disadvantage is that the data do not come by age or sex group which makes the approach aggregate in nature. However, given our goal to understand the long-run determinants of the declining migration inflows to Ontario, the benefits of using origin-destination migration stream outweigh the costs.

Let us now discuss the variables used in some detail. First of all, the dependent variable. Since our focus is on the destination region, Ontario, we normalize the migration flows with the population in this province. Letting the home area h and the destination area i, $m_{ht} = Pop_t^{-1} M_{ht}$, where M_{ht} is the total migration flow from h to Ontario and Pop_t is the total population in Ontario, at time t.

The explanatory variables are those typically used in previous studies, such as Courchene (1970), Shaw (1986), and Osberg et al. (1994). According to standard models the key determinant of migrations is the expected income differential, which, assuming for simplicity static expectations, is a function of current unemployment and income differentials. To capture the growth in the ability to support the unemployed population we include in the model log income per capita in the home region as well. In both cases we use disposable income, thus taking into account the regional differences in income tax rates. In addition, we have also included per capita federal transfer differentials between origin and destination provinces. Federal transfers, which includes the equalization payment[7] and unemployment benefits, are an important determinant of the so-called "fiscally-induced migration" (e.g. Day and Winer, 2006). Formally, the log differentials between h and Ontario are defined as $x^d_{hit} = x_{ht} - x_{it}$, $x = y, u, g$, with the symbols y, u, and g indicating log of disposable income per capita, unemployment rate, and federal transfer per capita, respectively, and $i =$ Ontario.

Finally, we include a measure of migration chain effect. Migrants are known to move

[6] A detailed data source list is provided in Appendix A.

[7] The idea behind the equalization program is to provide similar levels of public services among the provinces.

with a higher probability to destinations where people from the same area have moved to in the past, as it is easier both to obtain information and to receive material support when settling down.[8] Considering that the probability of accessing information and support is proportional to that of a contact with a past migrant, we define as a measure of the migration chain effect the total migrations from the home area to the destination area over the previous three years[9], divided by the total population of the home area: more precisely, $c_{ht} = Pop_{ht}^{-1} \sum_{s=1}^{3} M_{h,t-s}$.

The starting model for our empirical analysis is then the following:

$$m_{hit} = \beta_{0h} + \beta_{1h} y_{hit}^{d} + \beta_{2hi} u_{hit}^{d} + \beta_{3hi} g_{hit}^{d} + \beta_{4h} y_{ht} + \beta_{5h} c_{hit} + \varepsilon_{ht} \tag{1}$$

where h = home = NL, PE,...,BC, i = destination = ON, with $h \neq i$; $t = 1974, ..., 2004$, as some initial observations are needed to initialize the migration chain variable.

Anticipating the sign of slope parameters, an increase in y^d, g^d, and y in the home area is likely to reduce the (out) migration flows to Ontario. Therefore, we expect the slope parameters to obey $\beta_1 < 0$, $\beta_3 < 0$, and $\beta_4 < 0$. While, it is expected that $\beta_2 > 0$ and $\beta_5 > 0$ as higher unemployment differentials and higher values of the migration chain variable are expected to foster migration.

3 Econometric Methods

Since we are interested in the long-run patterns we need to carry out cointegration analysis, first testing if (1) describes an equilibrium relationship, and, provided this is the case, estimating the values of the coefficients. However, with only 31 time periods, the power of conventional cointegration tests is known to be very low. Fortunately, a solution is readily available: considering each origin-destination pair as a "unit" our dataset is naturally seen as a panel, with the number of units $N = 9$ and that of time observations $T = 31$. We can thus obtain higher power by applying some panel cointegration tests.

By now there exists a burgeoning literature on non-stationary panels suggesting nu-

[8] An obvious example is the frequent case of male heads of families migrating alone first, with their wives and children joining them after some time.

[9] Using a longer time span would have in principle been desirable, but it was feasible because of the very short time sample available.

merous panel estimators and tests for panel cointegration.[10]. The advantage of using panel data approach over standard time series methods is that by combining the information coming from both the cross-section and the time dimensions, the power of the tests can be increased, even without imposing any homogeneity assumption. When dealing with panel data, it is important to keep in mind the issue of cross-section dependence due to common trends and cycles in output across Canadian provinces (e.g., Wakerly et al., 2006). Ignoring the cross-section correlation is known to cause severe size distortion (e.g. Banerjee et al., 2004), so that the power gain delivered by the panel dimension, which is the very reason for its use, is entirely fictitious.

In this paper, we apply the bootstrap panel cointegration tests recently proposed by Fachin (2007). This test is robust to both short- and long-run dependence across units and delivers good small sample performances, which is important since in our case $N = 9$. The basic principle of this test[11] is to compute a summary statistic (say, G) of the no cointegration tests for the individual units on the empirical data set and on a large number of pseudo-datasets constructed under the null hypothesis of no cointegration (say, G^*). The no cointegration hypothesis is rejected if the empirical statistic G falls in the tail of the distribution of the $G^{*\prime}s$; in the classical Engle-Granger cointegration test, if the bootstrap p-value $p^* = prop(G^* < \widehat{G})$ is small. Clearly, a key point of the procedure is the construction of the series under no cointegration. These are obtained applying Paparoditis and Politis (2001) Continuous-Path Block Bootstrap (CBB); more details, which are beyond the scope of this paper, are given in Fachin (2007). Natural choices of summary statistics are the mean ($\overline{G} = N^{-1}\sum_{i=1}^{N} ADF_i$, where ADF_i is the ADF statistic computed on the residuals of the $i - th$ cointegrating regression) corresponding to Pedroni (1999) Group t-statistic, and the median, $G^{me} = Median(ADF_1, \ldots, ADF_n)$. In both cases the null hypothesis is 'cointegration in no units', against the alternative hypothesis 'cointegration in a large number of units'. Hence, in case of rejection we are *not* implying that in the specific time sample at hand cointegration holds in *all* units. Rather, we are implicitly taking what we may define a democratic stance: since the units are reasonably homogenous, if a long-run equilibrium exists in most of them, the exceptions are regarded as due to temporary conditions, which will vanish asymptotically.

As mentioned above, if cointegration is found to hold we can proceed to estimate the

[10]See Breitung and Pesaran (2007) and the references therein for an overview of the field.
[11]A more detailed description is given in Appendix B.

coefficients of the long-run relationships. From model (1) letting $h = NL, \ldots, BC$, we obtain a system of nine equations. This system is likely to be characterised by strong correlation of the disturbances and cointegration of the explanatory variables across equations. The first point would suggest that efficiency gains may be obtained applying SUR estimation methods (such as FM-SUR by Moon, 1999, or DGLS by Mark, Ogaki and Sul, 2005). Unfortunately, all these methods require the inversion of the long-run covariance matrix of the system, singular under cointegration across units, and are thus not feasible. However, the issue is not as serious as it might appear: according to simulation results by Moon and Perron (2005), the efficiency gains delivered by SUR estimators with respect to single-equation estimators such as FM-OLS are in fact essentially negligible. We can thus safely proceed to separate estimation by FM-OLS of each individual equation.

3.1 Testing for Cross-section Dependence in Panels

Recent developments in the literature offer the possibility of testing for the presence of cross-section dependence among individuals. Pesaran (2004) presents a simple test of error cross-section dependence that is valid asymptotically under very general conditions and can be applied to both stationary and non-stationary panels. The test statistic based on the average of pair-wise Pearson's correlation coefficients $\hat{p}_j$, $j = 1, 2, \ldots, n$, $n = N(N-1)/2$, of the residuals obtained from the estimation of autoregressive (AR) regression models. The CD statistic in Pesaran (2004) is given by

$$CD = \sqrt{\frac{2T}{n}} \sum_{j=1}^{n} \hat{p}_j \rightarrow N(0,1) \tag{2}$$

This statistic tests the null hypothesis of cross-section independence against the alternative hypothesis of dependence. Simulation results in Pesaran (2004) show that the statistic has reasonably good finite sample performance in terms of size and power.

4 Empirical Results

We first examined the time series properties of the series with individual ADF unit root tests, reported in Table 2. Consistently with our expectations, the migration rates are clearly I(1); the same holds for income level in the sending provinces. In these two cases the deterministic kernel has been chosen on a priori grounds, with a linear trend respectively included in the latter and excluded in the former. In the the remaining cases (income, unemployment and federal transfer differentials) we followed the selection procedure proposed by Ayat and Burridge (2000), including a deterministic trend when significant.[12]

As to be expected, this criterion leads always to tests with constant only in the case of the unemployment differential, while for the other two differentials both cases are present. For the income differential the I(1) hypothesis is never rejected except one case, at the 5%, while this happens in three cases for the unemployment differential. For both variables there are no rejections at 1%. The federal transfer differentials appear somehow more stationary, with three rejections at 5% and two at 1% (Quebec and Newfoundland and Labrador). We decided to take a rather conservative stance, excluding from the initial models only the variables found stationary at 1%. Also, given the results on the migration rates we always included a migration chain variable.

To get a feeling of the size of the cross-section dependence problem in the data, we have computed some cross-correlation tests. The results are summarized in Table 3. The short-run cross-section correlations are based on the regression residuals obtained from (1). At first sight, there seem to be only moderate evidence of cross-correlation in the data, as only 16 out of 36 cross-correlations are significant according to the usual approximate critical bounds. However, the average absolute correlation[13] between all the cross-section units is 0.382, which is not negligible. Further, Pesaran's (2004) CD statistic strongly rejects the null hypothesis of no cross-section dependence at least at the 1% level of significance. Hence, overall there seems to be enough evidence suggesting the presence of cross-section dependence in model (1), confirming the need to apply a robust testing procedure.

The results of panel cointegration tests with fully heterogenous specification (fixed effects, heterogenous slopes) are reported in the top panel of Table 4, and FM-OLS estimates of the individual equations in the bottom part of the same table. The bootstrap

[12]Although the idea of a linear trend in a differential may appear not plausible, this is not actually the case. If the linear trend enters the univariate Data Generating Processes of a variable with different coefficients in the home and destination regions the differential will have a linear trend also, with coefficient the difference of the regional coefficients.

[13]This is a useful measure in cases where the sign of the cross-correlations is alternating, thus pushing the correlations near to zero.

algorithm used 1000 redrawings and block length fixed at 4. The p-values for both the mean and median of the individual ADF statistics are all very small[14], indicating that that the models specified are cointegrating relationships. We can then examine the FM-OLS estimates of the cointegrating coefficients. For brevity, we report only the final specifications[15].

The plots of the series and FM-OLS estimates, shown in Figure 2, reveal that the models manage to capture rather accurately both the main trends and some local swings. The income differential (y^d) enters only in the equations for three Atlantic provinces (NL, PE, NB) but with a very large elasticity. Compared to this, home income (y) might seem to be a more important explanatory variable, as it enters in five equations. However, its elasticity is always rather small (average around 0.46 in absolute value).

Save for Alberta, we see that migrations originating in the westerly provinces are more likely to respond to changes in home income, while the case is quite mixed for provinces located east of Ontario with three out five home areas responding strongly to income differential. As argued in Faini et al. (1997), a higher home income implies that it may have become easier for households to finance protracted period of unemployment of some of their members causing migration flow to decrease.[16]

The unemployment differential (u^d) enters in seven equations with a near average of unit elasticity. As can be seen, migrations originating in the western provinces respond comparatively more strongly than those of the eastern provinces, in particular the Atlantic ones. This is not surprising given the more generous unemployment benefits existing in the maritime provinces compared to other provinces[17]. It is instead somehow puzzling to find that Manitoba and Saskatchewan react strongly to the unemployment differential, since their condition compared to the average of the provinces has improved or remained approximately constant (see Table 1).

[14] Note that the FDB p-values, computed through Davidson and MacKinnon's (2000) Fast Double Bootstrap, are very close to the base boostrap values, suggesting these to be very reliable.

[15] Given the small time sample available we followed the model selection procedure discussed in Fachin (2007).

[16] We are assuming that income was always high enough to make the financing effect (higher home income making it easier to finance the costs of migration, resulting a positive income-migration link) irrelevant.

[17] A reverse causality running from higher unemployment insurance inducing people to move to the relatively high unemployment region may not be supported by the data and in line with the findings by Lin (1995) and Day and Winer (2006). In other words, Atlantic region does not attract migration by proving higher unemployment benefits, they have higher unemployment rates and thus do not get any migration.

Next, federal transfer differential (g^d) is seen to be relevant for the two most western provinces, while the migration chain (c) measure is relevant only for the three Atlantic provinces. The latter indicates that family and friends who have previously migrated from the Atlantic region to Ontario may have provided important information about their present location which may have made the social transition easier for new migrants from their former locality.

Summing up, our estimates, in line with previous work on interprovincial migration in Canada (e.g. Helliwell, 1996; Day and Winer, 2006; and Coulombe, 2006), suggest the existence of a very strong link between unemployment and migrations. The evidence on income effect is mixed, with the level of home income appearing more important than relative income differentials.

5 Conclusions

We started our study from the observation that migration from the Canadian provinces towards Ontario mostly followed a declining trend over the period 1971-2004. In view of the process of convergence in both income and labour market conditions which, though by no means complete, did however took place during this period this is not surpring. However, important questions are open: is the convergence process really the explanation for these migration trends? Have income and labour market conditions been equally important? Given the small time sample studying this phenomenon using conventional cointegration methods would be difficult; further, strong short- and long-run cross-section dependence make the most popular asymptotic panel cointegration tests unfit. The solution is provided by the bootstrap panel cointegration test recently proposed by Fachin (2007). Applying this test we have been able to reach rather clear conclusions. First of all, a rather small and natural set of determinants (income, labor market and fiscal differentials, income in the sending provinces) is able to explain the observed declining trends. Second, estimation of heterogenous models suggests that the determinants of migration vary across provinces, with unemployment differential and income in the sending province the most important explanatory ones. Income and federal transfer differentials appear to play only a minor role.

Hence, the general implication of our results is that interprovincial migration has been significantly reduced by shrinking differentials in the labour market and income growth in the sending provinces. At this point, an obvious remark is that our macro approach leaves open a natural and important question: if unemployment in a province tends to increase the outmigration of labor force from that province, who is more likely to move? The unemployed, or, rather, those already employed? Unfortunately data limitation does not permit us to shed light on this important question. Lack of data also prevent us to analyze movements in intraprovincial migration which plays even more important role in labour market adjustment than interprovincial migration (e.g. Lin, 1995). With the availability of more data future research should focus on this area.

6 Appendix A. Data

All data are annual and come from Statistics Canada's E-STAT database. All income figures are expressed in Canadian dollar, while per capita figures are obtained by normalising by the population. The data is available from the corresponding author on request.

- *Migration (persons)*: Interprovincial migration flow data by province of origin and destination, 1971/72 to 2005/06, E-STAT Table 051-0019. For convenience, we treat annual years as 1971 to 2004.
- *Population (persons)*: Total population by province, 1971 to 2006, E-STAT Table 051-0001.
- *Disposable income (dollar)*: Personal disposable income by provinces, 1981 to 2004, E-STAT Table 384-0013. Data for the period 1971-1980 is available from E-STAT Table 384-0035.
- *Unemployment rate (percent)*: Data for 1976-2004 is taken from E-STAT Table 282-0002. Remaining data for 1971-1975 come from E-STAT Table 384-0035.
- *Federal transfer (dollar)*: Federal government current transfer to persons, 1981-2004, E-STAT Table 384-0004. Remaining data for 1971-1980 come from E-STAT Table 384-0022.

7 Appendix B. Bootstrap algorithm

Denoting by G a summary statistic of the no cointegration tests for the individual units (e.g. $G = N^{-1}\sum_{i=1}^{N} ADF_i$, or $G = Median(ADF_1, \ldots, ADF_n)$, where ADF_i is the ADF statistic computed on the residuals of the $i-th$ cointegrating regression) the proposed bootstrap procedure includes five simple steps:

1. compute the Group statistic $\widehat{G}$ for the data set under study,
   $\{X_1X_2 \ldots X_N, Y_1Y_2 \ldots Y_N\}_{t=1}^{T}$;
2. construct separately by CBB two sets of N pseudo-series,
   $\{X_1^*X_2^* \ldots X_N^*\}_{t=1}^{T^*}$ and $\{Y_1^*Y_2^* \ldots Y_N^*\}_{t=1}^{T^*}$;
3. compute the Group statistics G^* for the pseudo-data set,
   $\{X_1^*X_2^* \ldots X_N^*, Y_1^*Y_2^* \ldots Y_N^*\}_{t=1}^{T^*}$;
4. repeat steps (2) and (3) a large number (say, B) of times;
5. compute the boostrap significance level; assuming that the rejection region is the left tail of the distribution, $p^* = prop(G^* < \widehat{G})$.

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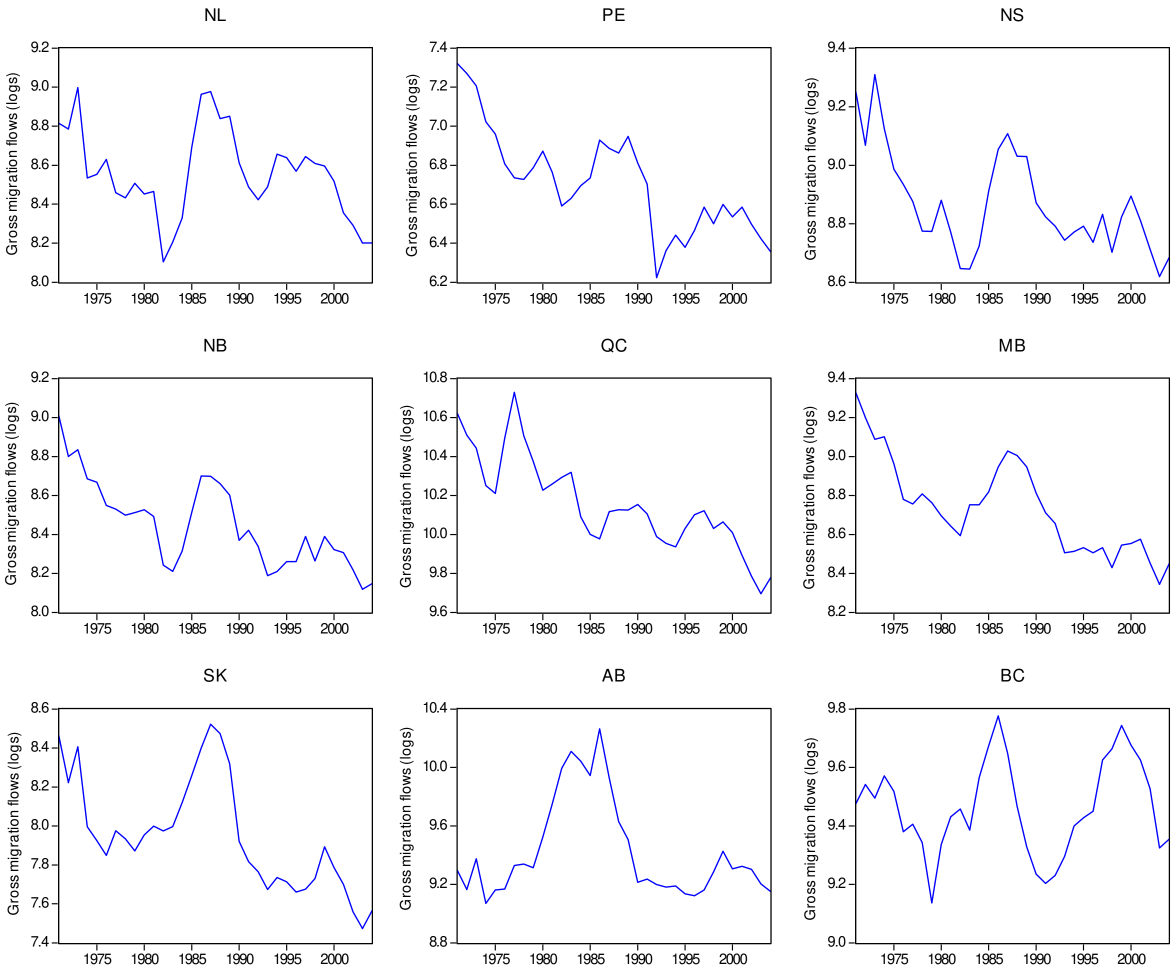

Figure 1: Total gross migration flows to Ontario, 1971-2004

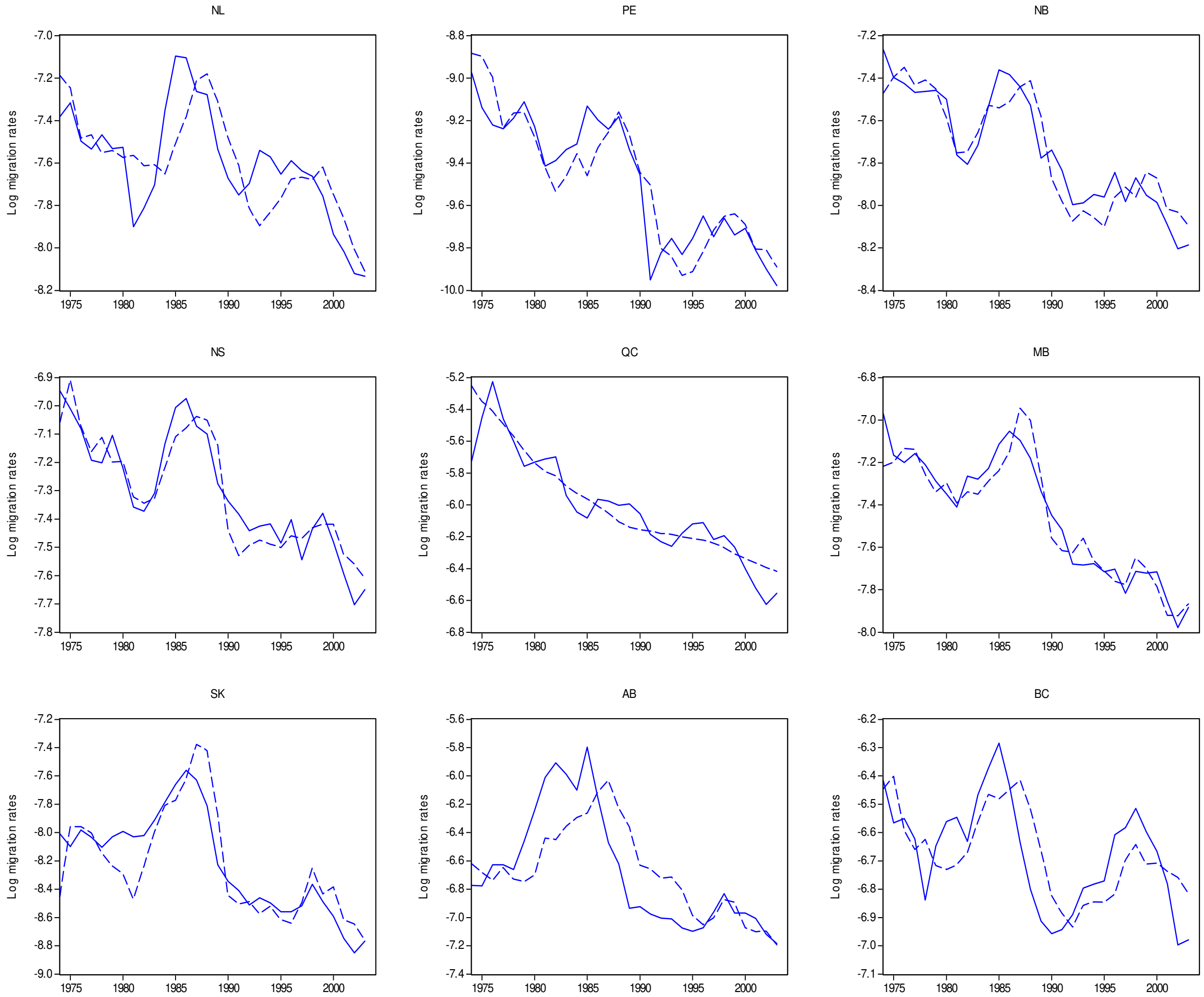

Figure 2: Log migration rates (solid line) and FM-OLS estimates (dashed line), 1974-2004.

Table 1: Population, income and unemployment in the Canadian provinces, 1971-2004

		y		Δu		Population	
Province	Code	1971	2004	1971	2004	1971	2004
Newfoundland and Labrador	NL	67.8	82.3	2.4	**7.4**	2.4	1.6
Prince Edward Island	PE	66.4	85.4	−2.3	**3.0**	0.5	0.4
Nova Scotia	NS	77.9	89.7	**1.0**	**0.5**	3.6	2.9
New Brunswick	NB	74.6	88.4	**0.1**	**1.5**	2.9	2.4
Quebec	QC	91.6	93.0	**1.3**	**0.2**	28.0	23.7
Ontario	ON	**115.3**	**104.5**	−0.6	−1.5	**35.8**	**38.9**
Manitoba	MB	93.9	92.6	−0.3	−3.0	4.6	3.7
Saskatchewan	SK	84.3	91.8	−2.5	−3.0	4.3	3.1
Alberta	AB	96.6	**117.6**	−0.3	−3.7	**7.6**	**10.1**
British Columbia	BC	**105.9**	96.8	**1.3**	−1.1	**10.2**	**13.2**

Notes: (i) y: Disposable income, Canada = 100. Bold face: values greater than 100. (ii) Δu: differential of provincial unemployment rate $\times 100$ from simple mean across provinces. Bold face: values greater than zero. (iii) Population: share (Canada = 100). Bold face: increasing share.

Table 2: ADF unit root tests

Province	m	y	y^d	u^d	g^d
NL	-1.41	-3.13^T	-0.88	-3.37^*	-4.12^{**}
PE	-1.65	-3.01^T	-3.97^*	-3.73^*	-3.05^*
NS	-1.80	-2.36^T	-2.67^T	-2.44	-3.24^*
NB	-1.84	-2.57^T	-2.80^T	-3.04^*	-3.55^*
QC	-0.08	-3.04^T	-1.30	-2.99	-4.33^{**}
MB	-1.90	-2.59^T	-3.67	-2.88	-3.01^T
SK	-1.34	-3.09^T	-2.88	-2.00	-2.66^T
AB	-0.85	-3.00^T	-1.29	-2.57	-1.23
BC	-2.32	-3.03^T	-2.42	-3.36^*	-2.05

Notes: m: migration flows/Ontario population; y: disposable income; y^d: disposable income differential; u^d: unemployment differential; g^d: federal transfer differential. All variables in logs, differentials with respect to Ontario. Number of lagged differentials selected on the basis of $t-$tests. T : linear trend included. ** and * significant at 1% and 5%, respectively.

Table 3: Residuals Cross-correlation

Province	1	2	3	4	5	6	7	8	9
NL	1.00								
PE	0.30	1.00							
NS	0.18	0.14	1.00						
NB	0.13	0.38	0.72	1.00					
QC	0.07	0.03	-0.06	-0.11	1.00				
MB	0.06	0.30	0.53	0.47	-0.36	1.00			
SK	0.52	0.44	0.51	0.47	0.16	0.38	1.00		
AB	0.60	0.35	-0.11	-0.02	0.08	0.19	0.54	1.00	
BC	0.40	0.34	0.33	0.45	-0.01	0.24	0.50	0.58	1.00

CD Statistic: 11.387 (0.000)

Notes: The cross-section correlation matrix is computed based on the residuals in (1). The total number of cross-correlations is 36, of which 16 are significant according to the usual approximate critical bounds $\pm 2T^{-1/2}$=$\pm$0.35. The CD statistic refers to the test statistic proposed in Pesaran (2004); p-value in parentheses.

Table 4: Modelling migration rates in Ontario, 1974-2004

Panel cointegration tests			Bootstrap p-values $\times$ 1000			
			Base	FDB_1	FDB_2	
Mean EG	-4.21		0.00	0.10	0.00	
Median EG	-3.74		0.30	0.10	-0.10	
FM-OLS estimates						
Origin	θ	y^d	u^d	g^d	y	c
NL	-5.35 [-10.91]	-2.70 [-7.84]	–	–	–	0.90 [6.48]
PE	-7.50 [-11.62]	-2.72 [-5.23]	0.33 [3.24]	–	–	0.73 [5.65]
NS	-4.29 [-13.30]	–	1.03 [7.99]	-1.06 [2.26]	-0.21 [-7.15]	0.26 [2.48]
NB	-8.79 [-89.56]	-2.74 [-7.09]	0.65 [4.84]	–	–	–
QC	0.88 [0.98]	–	–	–	-0.73 [-7.72]	–
MB	-2.85 [-7.10]	–	1.03 [11.52]	–	-0.47 [-11.26]	–
SK	-0.60 [-1.03]	–	1.51 [15.02]	–	-0.77 [-12.71]	–
AB	-6.55 [-111.43]	–	1.44 [6.22]	-2.38 [-4.00]	–	–
BC	-5.03 [-9.27]	–	0.94 [3.80]	-1.54 [-1.80]	-0.16 [-3.10]	–

Notes: Mean/Median EG: mean/median of the Engle-Granger ADF cointegration tests for the individual equations. The bootstrap algorithm used 1000 redrawings with a fixed length of 4. FDB indicates Fast Double Bootstrap, type 1 and 2 of Davidson and MacKinnon (2000). θ: constant; y^d: log disposable income per capita differential (home-destination); u^d: log unemployment rate differential (home-destination); g^d: log federal transfer per capita differential (home-destination); y: log GDP per capita in home area; c: migration chain. t-statistics are reported in brackets.