Spatial regularities between non-immigrant and immigrant numbers in Canada

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Abstract
The present study intends to reveal spatial regularities between non-immigrant and immigrant numbers in two different ways. First, it questions the existence of those regularities when spatial scales get finer. Second, it uses pooled data over four population censuses covering the period from 1991 to 2006, which enabled us to apply appropriate techniques to remove those unobserved fixed effects so that the estimations would accurately identify the linkage between local immigrant and non-immigrant numbers. The results provide evidence about the existence of negative spatial regularities between non-immigrant and immigrant numbers in Canada at national scale.

Keywords: Displacement, immigration, residential mobility, crowding-out, migration.

Introduction

Based on its medium-growth scenario, Statistics Canada projects that the natural growth rate of the population of Canada will be negative in 2028 (Statistics Canada, 2005). Smaller provinces and rural areas have already begun to face population declines as they also face the out-migration of the local population. As a result, provincial leaders are now adopting initiatives to increase their shares in annual Canadian immigrant inflows and to retain the new arrivals (Akbari, 2009).

However, rising immigration levels in an area may result in the out-migration of the native born. This is likely to happen if the immigrants displace the native-born workers in employment, bid down wages, or cause housing prices to rise through increased demand for shelter. Besides these economic reasons, the native born may also have some degree of social avoidance toward immigrants.

The economic and social impacts of immigration have been the strongest in Montreal, Toronto, and Vancouver where about 80 per cent of the immigrants land each year. Hou and Bourne (2004) and Ley (2007) have studied the effect of rising immigrant inflows on the departure of the native-born and established immigrant population for these cities. However, there are currently no studies available on whether these internal migration effects of immigration also apply in smaller areas in Canada. The literature does contain some discussion with regard to the United States where the "crowding-out" effect of immigration is not only a phenomenon of large metropolitan areas but also smaller areas (Card, 1997; Wright et al., 1997). If there is a significant displacement effect of immigration in Canada, the recent policies towards re-

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gionalization of immigration may not address the current demographic problems faced by the smaller provinces and their communities.

The objective of this study is neither to discover the crowding-out effect of immigration nor to identify a set of possible channels that immigration would lead to displacements of non-immigrant local populations. The present study rather intends to reveal possible spatial regularities between non-immigrant and immigrant numbers in Canada from 1991 to 2006 when local housing and labour market conditions are controlled and unobserved regional differences are removed. The research in this paper that looks for evidence on the displacement effect of immigrants is conducted in two different ways. First, we question the existence of those regularities when spatial scales get finer. Second, people may move internally due to spatial differences in amenities, climate, crime levels, culture, as well as time-specific reasons, such as a transition in the national economy that might make some local populations more susceptible than others due to the specifics of the local economic structure. We pooled the data over four population censuses covering the period from 1991 to 2006, which enabled us to apply appropriate techniques to remove those unobserved fixed effects so that the estimations would accurately identify the linkage between local immigrant and non-immigrant numbers.

The results provide evidence about the existence of negative spatial regularities between non-immigrant and immigrant numbers in Canada at national scale. The rest of the paper is as follows: Section 1 presents a summary of the literature. Section 2 explains the data used and estimation results. The interpretation of the results is at the end.

Literature review

Although the displacement effect of immigrants in gateway cities may be lifestyle driven, there is growing agreement among researchers that the lower-income, less-educated, native-born population is sensitive to immigrant inflows because this group most likely will be in direct competition with the new immigrants for the less-skilled and lower-paying jobs. Every year, Canada receives about 225,000 immigrants. Over the past two decades, more than three million immigrant workers entered local labour markets.

Many analysts have conducted studies on native-born mobility responses to immigration for the United States and have obtained mixed evidence.[1] Frey (1994, 1995, 1996, 2002) found strong native-born mobility responses leading to the "demographic balkanisation" of U.S. cities. Borjas et al. (1997) reported consistent evidence confirming the substantial out-migration of the native born in response to immigrant inflows on a national scale. Frey's displacement hypothesis has been challenged by White and Imai (1994), Wright et al. (1997), and Harrison (2002) who found that net in-migration of the native born is either positively related or unrelated to immigration in metropolitan areas. In fact, their results indicate that the net loss of unskilled native-born

[1] An excellent literature review can be found in the Hou and Bourne (2004) paper.

workers from metropolitan areas is probably a function of those cities' population size and industrial restructuring rather than immigrant inflows to them. Moreover, Card and DiNardo (2000) estimated the net impact of immigration inflows on the relative skill distribution of different cities in the U.S. and found that increases in the immigrant population in specific skill groups led to small increases in the population of native-born individuals of the same skill group. In a recent study, Borjas (2006) showed that the internal migration of the native born is a significant adjustment process that accounts for as much as 60 per cent of the difference between wage effects of immigration estimated by skill-cell and spatial correlation approaches. Federman et al. (2006) tested for native-born responses to the arrival of Vietnamese immigrants in the manicurist occupation in California and concluded that the displacement effect was not due to the exit of native-born workers but fewer new entries of native-born manicurists.

Hatton and Tani (2005) reviewed migration patterns across 11 regions of the United Kingdom using annual data for the period from 1981 to 2000. They found a strong negative link between immigration flows and native-born mobility responses. More specifically, for all 11 regions, their results showed that a 1 per cent increase in immigration reduces net in-migration of the native born by 0.064 per cent, implying that immigration induces native-born residents to relocate to other cities.

Unlike the U.S., Canada has a point system that targets skilled immigrants to reduce the labour shortages in specific markets. In that sense, it could be expected that selective immigration policies may bring more "complementary" new immigrant workers in Canada rather than a stream of "substitute" foreign labour that competes with native-born workers for the existing jobs with lower wages as arguably being the case in the U.S. However, although the Canadian immigration system largely aims at skilled workers, less than fifty per cent of the new immigrants come through the point system. No Canadian study has reviewed native-born mobility responses to immigration on a national scale. Two recent studies, Hou and Bourne (2004) and Ley (2007), found that the growth in recent immigration covaries with out-migration rates among the less-educated native born in Toronto and Vancouver, which are traditional immigrant destinations in Canada. While Ley compared Sydney (Australia) and Toronto by using time-series data between 1977 and 2002, Hou and Bourne calculated in- and out-migration rates by using multivariate logistic regression techniques on a sample of micro data drawn from five censuses from 1981 to 2001 for the working population aged between 25 and 64 living in three CMAs (Toronto, Montreal, and Vancouver). To our knowledge, this study is the first of its kind that reports the association between regional immigrant and non-immigrant populations at a national scale by using four population censuses covering 15 years with two different and consistent geographical classifications in Canada.

Data and estimation results

We use the 1991, 1996, 2001, and 2006 population censuses to build panel data at two geographical levels: census metropolitan areas (CMA) and census divisions (CD).[2] One problem with pooling censuses is that the geographical coverage of some CDs and CMAs changes over time. Therefore, we used the concordance tables, provided by Statistics Canada, to drop the regions whose boundaries have changed significantly. The period this study covers is peculiar because of two reasons. First, between 1991 and 2000, Canada received around 2.2 million new immigrants, which is the largest decadal inflow in the past 100 years. Second, during the 1990s, the major source of immigration shifted further away from Europe to Asia and other third-world countries. Immigrants coming from these regions are distinguished as "visible minorities" in the censuses after 1991.

It should be noted that changes in the number of local residents could result not only from population flows but also by differences in natural growth patterns in the region.[3] To address this issue, we also use the mobility of the local population to test the sensitivity of local population mobility to changes in the immigrant density across censuses and regions. Lastly, the resulting five-year interval between censuses provides a reasonable time window for the individuals to respond to changes in local conditions.

Table 1: Components of Population Growth (1971–2006) (in percentages)

Provinces	Total	Native Born	Immigrants
Canada	44.85	35.66	87.74
Newfoundland and Labrador (NL)	-4.12	-4.35	-6.32
Prince Edward Island (PEI)	20.21	19.66	29.15
Nova Scotia (NS)	14.47	13.66	21.51
New Brunswick (NB)	13.41	13.07	11.25
Quebec (QE)	23.36	17.57	81.60
Ontario (ON)	56.16	41.97	99.06
Manitoba (MB)	14.70	16.46	-0.01
Saskatchewan (SK)	2.98	10.49	-56.49
Alberta (AB)	100.04	100.82	86.72
British Columbia (BC)	86.50	72.06	125.35

Source: Authors' calculations based on the data. We extend time period in this table to 1971-2006 in order to show the long-term trend between the growth rates.

Table 1 gives some idea about the provincial differences in population growth rates between 1971 and 2006. It shows that although British Colum-

[2] There are 289 provincially legislated census divisions (CDs). A CMA refers to the main labour market area of an urbanized core having 100,000 or more population. CMAs are created by Statistics Canada and are usually known by the name of the urban area forming their urbanized core.

[3] Most of the studies mentioned in the literature review use the same measure: percentage changes in the number of local residents. See Borjas et al. (1997), Filer (1992), Walker et al. (1992), and Frey (1994, 1995).

bia, Ontario, and Quebec experienced large increases in their immigrant population relative to the increase in the native-born population the opposite was true in Alberta, Saskatchewan, and Manitoba. Manitoba and Saskatchewan experienced a decline in their immigrant populations during the same period.

To identify correlations between the non-immigrant and immigrant numbers, we conceptualize the relationship by a simple spatial equilibrium model similar to one used by Saiz (2007). More specifically we can express changes in supply of non-immigrant (*N*) residents in region *r* as a function of changes in regional-housing (*h*) and labour-market conditions (*L*), the presence of social avoidance and/or self-selected ethnic segregations (*M*), and some local fixed-effects (*Z*) such as the quality of local amenities, climate, regional differences etc. as follows:

$$\Delta \ln N_{r,t} = f(\Delta \ln h_{r,t}, \Delta \ln L_{r,t}, \Delta \ln M_{r,t}, Z_r) \qquad (1)$$

A number of observations can be made on (1). First, unfavourable spatial differences (e.g., in income levels and housing costs) might have a negative effect on the growth rate of non-immigrant residents. Second, the effect of immigration on non-immigrant mobility is not independent of the impact of immigration on local labour and housing markets. Lastly, the coefficient on fixed-effects (*Z*) cannot be identified in the above setting.

Based on (1), the estimating framework takes the following form:

$$\Delta \ln(N_{rt}) = \beta \Delta \ln(M_{rt-1}) + \delta_k \mathrm{X}_{rt-1} + \alpha Z_r + u_r + \tau_t + e_{rt}, \qquad (2)$$

where the dependent variable is the growth rate of local non-immigrant residents measured by the change in the log of *N*. The adjustment of the non-immigrant residents may not be contemporaneous in response to immigrant inflows and changes in local conditions. We use lagged values for the growth rate of immigrant residents, $\Delta \ln(M_{rt-1})$, and other explanatory variables, X_{rt-1}, which is a 1 x *k* vector of variables that vary over region and time. Lastly, Z_r is a vector of variables that varies only over regions; u_r and τ_t are the unobserved region-specific and time-specific effects, respectively, and e_{rt} is the idiosyncratic disturbance term that satisfies $e_{rt} \sim$ i.i.d. $(0,\sigma_e^2)$ for all *t* and *r*.[4]

Following the literature, first, we predict that, all things being equal, people tend to move to regions with lower housing costs. As proposed by Poterba (1991), in equilibrium, the expected cost of owning a house should equal the cost of renting. Hence, we use two alternative variables to control regional

[4] As expected, the test results strongly reject the spatial independence so that the measured growth in the native-born population in one location may be correlated to those in neighbouring locations. we address this problem in the estimations by using clustered robust standard errors.

housing costs: gross average monthly rents for residential properties and average housing prices. Second, we control the linkage between labour-market outcomes and population mobility by changes in the region's unemployment rate. Our prediction is that rising unemployment will be associated with declining growth in the non-immigrant population. Finally, to remove unobserved spatial differences, we apply a fixed-effect model that includes a full set of year dummies.

Table 2: Fixed-effect Estimators

CD	**2(a)**		**2(b)**		**2(c)**	
	Coef.	SE	Coef.	SE	Coef.	SE
ΔlnM	0.013	0.012				
Δln(*un*)	0.037	0.003				
Δln(*hprice*)	0.113	0.048				
L.ΔlnM			-0.011	0.004	-0.013	0.005
L.Δln(*un*)			-0.021	0.008	-0.018	0.012
L.Δln(*hprice*)			-0.011	0.033		
L.Δln(*rent*)					-0.038	0.020
# of obs.(*t, r*)	(3, 234)		(2, 234)		(2, 234)	
R2 (within)	0.2052		0.1633		0.1801	
Rho	0.7111		0.7993		0.8047	
CMA						
ΔlnM	0.222	0.040				
Δln(*un*)	0.043	0.034				
Δln(*hprice*)	0.104	0.044				
L.ΔlnM			-0.155	0.049	-0.169	0.070
L.Δln(*un*)			-0.056	0.021	-0.053	0.025
L.Δln(*hprice*)			-0.024	0.048		
L.Δln(*rent*)					-0.016	0.042
# of obs.(*t, r*)	(3, 31)		(2, 31)		(2, 31)	
R2 (within)	0.3892		0.3415		0.3323	
Rho	0.5099		0.785		0.7877	

*Notes: (1) The dependent variable is Δ*ln*N. (2) Standard errors (SE) are robust and adjusted by provincial (for CDs) and regional clusters (for CMAs). (3) rho indicates the fraction of the unexplained variance due to differences across regions. (4) All regressions have a set of dummy variables to control year effects (significant and not shown here). (5) L in front of the variable indicates the lagged value of the variable.*

The estimation results of (2) with and without lagged explanatory variables are provided in Table 2 above. The estimations in the first column 2(a) that use non-lagged explanatory variables show counterintuitive results, and, when compared with the results in 2(b) that use lagged variables, the estimations appear to be sensitive to the use of lagged covariates implying that the contemporaneous link between non-immigrant and immigrant resident growth rates might be subject to a simultaneity problem. Moreover, the test results confirm that unobserved regional fixed effects are not random; therefore, although we use growth rates, the use of a random-effect model is not justified.

Except for 2(a), the fixed-effect estimators consistently show statistically significant and negative relationships between the rates of change in numbers of local non-immigrant and immigrant residents. Although the unemployment rate has a predicted sign, its significance at the CD level is sensitive to the type of housing cost used in the estimations.[5] This could be explained by the fact that the dependent variable includes all non-immigrant residents not only workers. The results also imply that, even if the sign on the coefficient agrees with what the model predicts, rising housing prices or rents have no statistically significant effect on the number of local non-immigrant residents. By definition, (1) abstracts from income effects in housing consumption. Hence, a more sensible approach would be to consider income net of housing costs, not housing costs as measured by average rents or housing prices. This problem leads to a downward bias in the coefficients of the housing variables in the estimations (Glaeser, 2008). The different magnitudes in the coefficient of interest between CMA and CD levels likely come from the fact that very few CDs have more than 80 per cent of new immigrants. Hence using CD as a geographical category reduces the power of regressions as more observation points have big changes in the explanatory variable (growth rates of immigrant residents) but minor variations in the dependent variable.[6]

We now use interregional and international mobility measures for local residents and estimate the following version of (2):

$$\Delta \ln(I_{rt}) = \beta \Delta \ln(E_{rt-1}) + \delta_k \mathrm{X}_{rt-1} + \alpha \mathrm{Z}_r + u_r + \tau_t + e_{rt}, \qquad (3)$$

where I_{rt} is the number of local residents who lived in a different region in Canada five years ago and E_{rt} is the number of residents who lived in a different country five years ago.[7] Hence, specification (3) reveals the elasticity of interregional mobility to immigration when the regional housing and labour-market conditions are controlled. If the displacement effect of immigration is significant, we expect that regions receiving more residents from abroad will attract fewer internal migrations.

Table 3 shows the estimation results of (3) with and without lagged explanatory variables 3(a) and 3(b), respectively. As before, counterintuitive results in the first column imply the presence of a simultaneity problem. The estimations in the last two columns show a statistically significant β coefficient with expected signs both at the CD and CMA levels. Moreover, the type of housing cost used in regressions does not affect the significance of unemployment rates. This is perhaps because the dependent variable may now in-

[5] Removing housing variables from the regressions does not change the magnitude, sign and significance of coefficient of interest. However, the coefficient of unemployment becomes significant.

[6] We also applied WLS (Weighted Least Squares) with populations as weights. The results do not change significantly.

[7] We also used new immigrant densities instead of the number of people who lived in a different country five years ago. The results not reported here show a similar pattern.

clude more people who are in the labour force, assuming that people move mostly when they are younger for better labour-market opportunities. Finally, housing variables have no robust explanatory power on the number of internal migrants possibly due to the bias explained before.

Table 3: Fixed-effect Estimators with Mobility Measures

CD	**3(a)**		**3(b)**		**3(c)**	
	Coef.	SE	Coef.	SE	Coef.	SE
ln(*E*)	0.0058	0.0031				
ln(*un*)	0.1216	0.0577				
ln(*hprice*)	0.3348	0.1438				
L.ln(*E*)			-0.0087	0.0041	-0.0087	0.0041
L.ln(*un*)			-0.1512	0.0646	-0.1437	0.0658
L.ln(*hprice*)			-0.1212	0.1710		
L.ln(*rent*)					-0.0050	0.2659
# of obs.(*t, r*)	(4, 234)		(3, 234)		(3, 234)	
R2 (within)	0.4207		0.3591		0.3559	
rho	0.9804		0.9824		0.9834	
CMA						
ln(*E*)	0.0725	0.0722				
ln(*un*)	0.0985	0.1009				
ln(*hprice*)	0.2399	0.1249				
L.ln(*E*)			-0.1145	0.0721	-0.1419	0.0624
L.ln(*un*)			-0.2523	0.1621	-0.2580	0.1475
L.ln(*hprice*)			-0.0863	0.1582		
L.ln(*rent*)					-0.1991	0.1474
# of obs. (*t, r*)	(4, 31)		(3, 31)		(3, 31)	
R2 (within)	0.2817		0.1627		0.1633	
rho	0.9583		0.9926		0.9928	

Notes: (1) The dependent variable is lnI. (2) Standard errors (SE) are robust and adjusted by provincial (for CDs) and regional clusters (for CMAs). (3) rho indicates the fraction of the unexplained variance due to differences across regions. (4) All regressions have a set of dummy variables to control year effects (not shown here). (5) "L" in front of the variable indicates the lagged value of the variable.

Concluding remarks

This study investigated spatial regularities between local non-immigrant and immigrant resident numbers in Canada on a national scale. It used data from four population censuses covering the period 1991 to 2006 to build a panel by using the concordance tables at the CMA and CD levels.

The estimations results consistently showed that the growth rates of local non-immigrant and immigrant residents are negatively related when an adjustment period is allowed. After controlling local housing and labour market conditions, on average, a 1-percentage point increase in the growth rate of local immigrant residents is associated with a 0.15-percentage point decline in the growth rate of the non-immigrant local population across metropolitan areas. When the CD level geographical scale is used, which covers the entire

country, the same negative and robust relationship between the growth rates becomes less sizeable: the same decline becomes 0.01 percentage points. It is likely that measuring local population flows by growth rates may lead to this result. Because less than 20 per cent of immigrants live in more than 90 per cent of CDs, using CDs reduces the power of regressions. Instead of growth rates, when local mobility measures are used, the results showed that the elasticity of in-migration with respect to immigration inflows is around -0.14 for big cities. Yet, although it is negative and significant, the same elasticity becomes negligible in magnitude at the CD level.

The model implied that the included controls were actually the mechanism by which inflows would lead to outflows. Yet, the estimated coefficients on immigration variables were not independent of the elasticity of labour demand, the substitution between immigrant and non-immigrant workers, and the effect of immigrants on housing costs and labour market outcomes. Moreover, the model used here assumes that there is no segregation in local labour and housing markets for immigrant and non-immigrant residents. However, we know from the literature (Pedakur and Pendakur, 1998) that wage rates, unemployment incidences, and to some degree housing costs in segregated regions would differ considerably for both groups. Therefore averaging all these local market outcomes in the regressions would reduce their explanatory power.

The results provide evidence about the existence of negative and statistically significant spatial regularities between local non-immigrant and immigrant resident numbers. The results promise the value of expanding the current study in the future. As the data become more available, one way to deal with the problem that the very low concentration of immigrants in small regions reduces the explanatory power of regressions is to estimate bilateral inter-regional migration rates, which is left to a future work.

Acknowledgement

I thank Dr E. Falaris (University of Delaware) for his valuable comments on an initial draft. I also thank Dr Casey Warman for his useful comments on an earlier draft presented at the CEA conference in May 2011 in Quebec City. Responsibility for any shortcomings remains with the authors.

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