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Nonimmigrant Mobility Responses to Immigrant Inflows in Canada: a Panel Data Analysis

Yigit Aydede[1]

Abstract: Population planners in the smaller provinces of Canada have embarked on a mission to increase their share of the national immigration intake to offset, at least partially, their province's population decline. Within each province, maintaining the population balance through immigration across smaller communities is also a desired objective. Yet, rising immigration may also lead to out-migration of the native-born population from a community in the short run. If there is a strong link between immigrant inflows and native outflows, immigration may in fact magnify demographic problems across Canada instead of mitigating them. In this paper, I expand the current Canadian literature on immigration in three new directions. First, I question whether a linkage exists between immigration and native-born outflows as spatial scales get finer. Second, I employ a multi-regional framework where I analyze net internal migration on a national scale. Lastly, to control for fixed-regional and time-specific effects, I apply panel estimations using four population censuses that cover the period 1991-2006 at two geographic levels: metropolitan areas and census divisions. The results indicate that locations that receive high levels of immigration are less likely to retain or receive nonimmigrant residents in Canada.[2]

Keywords: Displacement, Immigration, Native-born Mobility
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Based on its medium-growth scenario, Statistics Canada projects that the natural growth rate of Canada's population will be negative by 2028 (Statistics Canada, 2005). Without immigration, not only will the Canadian population shrink, it will also age faster than it is today. Smaller provinces and rural areas, for example, have already begun to face population decline as they also experience the out-migration of the local population. As a result, provincial leaders are now adopting initiatives to increase their share of annual Canadian immigrant inflows and retain new arrivals (Akbari, 2009).

However, rising immigration levels in an area may also result in the out-migration of the native born. This is likely if 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 experience some degree of social avoidance to immigrants.

The economic and social impacts of immigration have been the strongest in Montreal, Toronto, and Vancouver, where 80 percent of immigrants land each year. Hou and Bourne (2004) and Ley (2007) have studied the effect of rising immigrant inflows on the departure of both the native born and established immigrant population for these cities. However, there are currently no studies have been published on whether these internal migration effects of immigration also apply in smaller areas in Canada. The literature does include some discussion on the situation in the United States, where the "crowding out" effect of immigration is a phenomenon not only in large metropolitan areas but also in smaller ones (Card, 1997; Wright et al., 1997). If this is also true in Canada, although recent policies toward regionalization of immigration may address the current demographic problems faced by smaller provinces and their communities, they may also lead to some corresponding social problems in the near future by the rising separation of immigrants from the rest of the population across regions.

The present study investigates the mobility responses of native-born Canadians to immigrant inflows in three different ways. First, I question the existence of spatial regularities between native-born mobility responses and immigration when spatial scales get finer. As the gravity model of migration hypothesizes (Zipf, 1946), the volume of migration between two places is inversely proportional to the distance between the two. This implies that migration is primarily a phenomenon of neighbouring localities, and it is therefore, highly likely that we can no longer observe population flows between two locations when they get larger. Hence, the use of data at census division levels that are consistent across censuses conducted between 1991 and 2006 is a unique aspect of this study.

Second, even though about 80 percent of immigrants arrive and live in major cities, large agglomerations are distinctive in their form and recent evolution. Thus, the native born may be leaving the large cities because of particularities in them, rather than because of immigration to them (Wright et al., 1997). In this paper, a multi-regional framework is employed in which nonimmigrant mobility responses are analyzed on a national scale by using data from smaller areas across Canada.

Third, people may move internally due to spatial differences in amenities, climate, crime levels, and culture, as well for as time-specific reasons, such as a transition in the national economy that might make some local populations more susceptible to moving than others due to the specifics of the local economy (Jackman and Savouri, 1992). I pooled the panel data for over four population censuses covering the period from 1991 to 2006, which enabled me to apply appropriate techniques to remove those unobserved fixed effects so that the estimations would accurately identify the impact of immigration on the mobility of the native born.

Although the results indicate no immediate negative spatial correlation between native-born and immigrant population flows in Canada, locations with high levels of immigration are eventually less likely to retain or receive nonimmigrant residents in the long run. The rest of the paper proceeds as follows: Section 1 presents a summary of the literature, and Section 2 presents the data. Section 3 shows the estimation results of a model first applied by Borjas et al., (1997). A spatial equilibrium framework is developed and estimated in Section 4 by using population growth rates and in Section 5 by using mobility measures. The final section includes the author’s interpretation of the results.

1. LITERATURE REVIEW

Although the displacement effect of immigrants in gateway cities may be lifestyle driven,[3] researchers increasingly agree 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 new immigrants for the less-skilled, lower-paying jobs. Every year, Canada receives about 225,000 immigrants, and over the last two decades, more than three million immigrant workers have entered local labour markets. Despite Canada's being one of the major immigrant-receiving countries in the world, studies investigating the effects of immigration on local markets are scarce. Studies that *have* analyzed labour market outcomes of immigration have produced mixed results. For example, using national data on industries and occupation,[4] Akbari and Aydede (2010), Islam (2009), Akbari and DeVoretz (1992), and Roy (1987, 1997) concluded that there is an imperfect substitution between immigrant and native-born workers. On the other hand, Aydemir and Borjas (2007) found a strong negative impact of immigrant inflows on labour market outcomes in Canada. Unlike others, they questioned how much of the disparity in the outcome of different skill groups of native-born workers can be attributed to immigration that shifts the relative demand and supply of different skills at the national level. In the immigration literature, this approach

[3] People may want to live in communities where they fit in better.

[4] There are also two recent studies on housing markets (Akbari and Aydede, 2009; Ley and Tutchener, 2001) that indicate a weak negative linkage between immigration and local housing prices.

(*skill-cell approach*) is justified by the fact that if substantial native-born outflows take place in response to immigrant inflows, a “naïve” spatial empirical study may even find a positive impact of immigration on local market outcomes for the native born. One proposed solution to this bias in the estimates of immigrant impact is to test directly the reactions of native-born populations to immigration.

Many studies conducted on native-born mobility responses to immigration in the United States have obtained mixed evidence.[5] 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. However, Frey’s displacement hypothesis was 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 workers from metropolitan areas is probably a function of those cities’ population size and industrial restructuring rather than of 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

[5] For an excellent literature review, see Hou and Bourne (2004).

the internal migration of the native born is a significant adjustment process that accounts for as much as 60 percent 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 due not to the exit of native-born workers but to 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 1981-2000. 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 percent increase in immigration reduces net in-migration of the native born by 0.064 percent, implying that immigration induces native-born residents to relocate to other cities.

No Canadian study has reviewed native-born mobility responses to immigration on a national scale. Two recent studies, those Hou and Bourne (2004) and Ley (2007), found that the growth in recent immigration co-varies 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 microdata 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). Hou and Bourne compared the effects of economic restructuring, housing market conditions, and immigrant in-inflows in a CMA on the trends in internal migration from and to that CMA. They found a significant correlation between growth in the recent immigrant population and an increased out-migration rate among low-skilled Canadians born in Toronto and Vancouver. However, this association becomes insignificant across CMAs, which implies that immigration may not be the major source of out-migration of the native born in gateway cities.

2. DATA

The 1991, 1996, 2001, and 2006 population censuses were used to build panel data at two geographic levels: census metropolitan areas (CMA) and census divisions (CD).[6] One problem with pooling censuses is that the geographic coverage of some CDs and CMAs changes over time. Therefore, the concordance tables, provided by Statistics Canada, were used to drop the regions whose boundaries have changed significantly (more discussion provided in the Appendix). Although only four censuses with 31 CMAs and 238 CDs were pooled to limit the number of regions dropped from the panel, the period this

[6] To conduct a population census, Statistics Canada divided the country into 289 provincially legislated census divisions (CDs) or smaller communities that are intermediate geographic areas between the province/territory level and the municipality (census subdivision). A CMA refers to the main labour market area of an urbanized core having a population of 100,000 or more. CMAs are created by Statistics Canada and are usually known by the name of the urban area forming their urbanized core.

study covers is peculiar for 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 away from Europe to countries in Asia and other, third world, countries. Immigrants coming from these regions are distinguished as “visible minorities” in the censuses after 1991.

The present analysis investigates local population flows that include all people, not just labour flows, mainly because data on the mobility of nonimmigrant workers by skill groups or labour force status are not available to us at the CD level and the cost of obtaining such data is too high. Although labour market effects on relocation decisions will be diffused in this setting, these data do have the advantage of including people who are not in the labour force but who move to other localities due to housing market outcomes and for reasons of social avoidance. [7]

Lastly, data on net population growth rates are used, so the results cannot distinguish in- and out-migration responses to immigration. Although population growth rates may be a blunt measure of mobility, they reveal spatial regularities between the net growth rates of nonimmigrants and immigrants in the local population. To address this issue, the mobility of the local population was used to test the sensitivity of local population mobility to changes in the immigrant

[7] This is similar to Frey’s models that cover all people.

density across censuses and regions.[8] The resulting five-year interval between censuses provides a reasonable time window for individuals to respond to changes in local conditions.

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

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 (PE)	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 (QC)	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: Author's calculations based on the data.

Table 1 gives a general indication of provincial differences about the provincial differences in population growth rates between 1971 and 2006. It shows that although British Columbia, Ontario, and Quebec each 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 each experienced a decline in their immigrant population during the same period.

Table 2 shows the provincial distribution of immigrant and native-born populations before and after 1996, the year smaller provinces started to more

[8] Borjas et al. (1997), Filer (1992), Walker et al. (1992), and Frey (1994, 1995) used net migration models.

actively attract and try to retain immigrants to offset their population decline and the effects of aging. Except for the Atlantic provinces, the shares of immigrant and native-born populations tended to move in opposite directions, with an accelerated rate particularly after 1996. While British Columbia, Ontario, and Quebec experienced large-scale immigration after 1996, the opposite trend was observed in the Prairie provinces.

Table 2: Distribution of Population by Immigrant Status (in percent)

	Canada		BC		QC		ON	
	N	M	N	M	N	M	N	M
1971	84.7	15.3	77.3	22.7	92.2	7.8	77.8	22.2
1981	84.0	16.0	76.9	23.1	91.8	8.2	76.4	23.6
1986	84.4	15.6	77.9	22.1	91.8	8.2	76.9	23.1
1991	83.1	16.1	76.9	22.3	90.7	8.7	75.0	23.7
1996	82.0	17.4	74.7	24.5	90.0	9.4	73.7	25.6
2001	80.9	18.4	72.9	26.1	89.5	9.9	72.3	26.8
2006	79.3	19.8	71.3	27.5	87.9	11.5	70.8	28.3
	MB		SK		AB		Atlantic	
	N	M	N	M	N	M	N	M
1971	84.7	15.3	88.0	12.0	82.7	17.3	96.4	3.6
1981	85.8	14.2	91.3	8.7	83.7	16.3	96.2	3.8
1986	86.4	13.6	92.8	7.2	84.2	15.8	96.4	3.6
1991	86.8	12.8	93.8	5.9	84.3	15.1	96.5	3.3
1996	87.3	12.4	94.4	5.4	84.4	15.2	96.4	3.4
2001	87.5	12.1	94.7	5.0	84.5	14.9	96.4	3.4
2006	86.0	13.3	94.5	5.0	83.0	16.2	95.9	3.8

Source: Author’s calculations based on the data. N and M represent native-born and immigrant populations. The Atlantic provinces are NL, PE, NS, and NB.

Annualized population growth rates (not shown here) also were calculated. An increasing immigrant population after 1986 coincided with a declining growth in the native-born population in British Columbia, Ontario, and Quebec. Finally, the author also reviewed the distribution of native-born and immigrant populations for 36 CMAs[9] and calculated components of their population growth rates for the period 1981-2006 (not shown here). It was interesting to observe, for example, that although Vancouver experienced a 123 percent increase in its immigrant population compared to a 40 percent increase in its native-born population, all major cities in British Columbia experienced the opposite trend. In fact, this was almost the same for all three immigrant-receiving provinces: all neighbouring CMAs to Montreal, Toronto, Calgary, and Ottawa experienced more growth in their native-born populations than in their immigrant populations. Although the data shown in Tables 1 and 2 are instructive and reflect immigration policy changes, using a provincial-level classification can mask possible crowding out patterns across metropolitan areas and/or smaller neighbouring regions within the same province.

3. *FIRST* AND *DOUBLE DIFFERENCE* REGRESSION MODELS

Borjas et al. (1997) compared native interstate migration to immigration in the U.S. during 1970-1990 by estimating *first difference* and *double difference* models. To understand whether nonimmigrant individuals adjust to the impact of immigration in an area by moving their labour or capital to other localities, I begin with a *first difference* model, which takes the following form:

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[9] Each subsequent census records new additions to its current CMA list as some census agglomerations (CA) reach 100,000 or more in population. Hence, to construct a balanced panel, we removed all new additions after 1981.

$$n_r(t,t+j) = \alpha + \beta m_r(t,t+j) + e_r, \qquad (1)$$

where e_r is the stochastic error and n_r and m_r are the annualized population growth rate contributions of the native born and immigrants in region r, respectively, as computed below:

$$n_r(t,t+j) = \frac{N_{rt+j} - N_{rt}}{jP_{rt}} \text{ and } m_r(t,t+j) = \frac{M_{rt+j} - M_{rt}}{jP_{rt}},$$

where N and M are the numbers of native-born and immigrant individuals living in region r in year t and P is the population. Equation (1) correlates the annual growth rate of the native-born population in region i to the growth rate of immigrants in the same region relative to the region's total population. Hence, the coefficient β measures the immediate effect of one additional immigrant arriving in region r between $t + j$ and t on the change in the number of native-born individuals living in that region during the same period.[10]

As pointed out by Borjas et al., one problem with (1) is that the first difference specification implicitly assumes each region could have the same growth rate in the native-born population absent of immigration. However, even if there were no immigration, many CDs (CMAs) probably would have different

[10] Some bias in the coefficient estimate is expected because these calculations do not reflect changes in the number of second generation immigrants born in Canada and the death rates among native-born individuals.

growth paths in any two given subsequent periods. As seen in Table 2, the surge in immigrant inflows after 1996 has changed the distribution of the population, particularly in the three major immigrant-receiving provinces. Thus, having contrasted two periods—the "pretreatment" period from 1991-1996 and the "post-treatment" period after 1996—allows us to isolate the impact of immigration on the change in the native-born population. To address the problem with the *first difference* specification, I apply a difference-in-difference comparison as specified below:

$$n_r(96{,}06) - n_r(91{,}96) = \alpha + \beta[m_r(96{,}06) - m_r(91{,}96)] + e_r, \qquad (2)$$

where the coefficient β now reveals what would have happened to the region's native-born population growth after 1996 if the immigrant supply "shock" had not occurred between 1996 and 2006.[11] Table 3 shows the estimated effect of changes in the number of immigrants on the number of native born by *first and double difference* specifications.

[11] Borjas et al. (1997) used the same method.

Table 3: Estimated Effect of Immigrants on the Population Growth Rate of the Native Born

CD	3(a)		3(b)		3(c)	
	β	*SE*	*β*	*SE*	*β*	*SE*
m (91,06)	1.060	0.503				
m (96,06)-*m* (91,96)			-1.046	0.006		
n (91,96)					0.051	0.058
m (96,06)					-1.057	0.019
m (91,96)					0.087	0.095
# of observations	234		234		234	
R^2	0.332		0.975		0.990	
CMA						
m (91,06)	0.456	0.494				
m (96,06)-*m* (91,96)			2.006	0.605		
n (91,96)					0.557	0.244
m (96,06)					2.499	0.531
m (91,96)					-3.169	0.743
# of observations	31		31		31	
R^2	0.268		0.229		0.407	

Notes: (1) For all estimations, the ordinary least square (OLS) method is used. (2) Robust standard errors (SE) are calculated and adjusted by provincial clusters for the CD level and regional clusters (ON, BC, QC, Atlantic, and Prairie) for the CMA level. (3) The dependent variables are *n*(91,06) for 3(a), *n*(96,06)–*n*(91,96) for 3(b), and *n*(96,06) for 3(c), as defined in the text above.

The first column in Table 3 reports the results of (1), estimated for CDs and CMAs separately. A positive coefficient of β rejects a negative link between the growth rates of immigrant and native-born populations. Because both n and m are scaled by the same denominator, β quantifies the effect of an immigrant arriving to the CD (CMA) on the change in the native-born population during the same period. On average, for each arriving immigrant, the native-born population increases by 1.06 persons in CDs and by 0.46 persons in CMAs.

Although the second column in Table 3 reports the results of *double difference* estimations specified by equation (2), the third column shows the unrestricted version of (2) that relaxes the restrictions on the coefficients. Controlling the region's pre-1996 population growth rate changes the sign of the coefficient β from positive to negative at the CD level, implying that a 1 percent increase in the growth rate of the immigrant population reduces the native-born population's growth rate by 1 percent (the coefficients are not significantly different from –1). At the CMA level, however, the results do not indicate a negative link. One reason for this could be a potential simultaneity problem that worsens, particularly for CMAs. Because arbitrary adjustment costs rise as the distance between source and receiving regions goes farther across CMAs, the nonimmigrant population may not adjust contemporaneously in response to immigrant inflows. In addition, specifications (1) and (2) do not control regional fixed effects efficiently, a problems that worsens as spatial differences become more significant across CMAs as opposed to across neighbouring regions.

Moreover, as CMAs have much larger labour markets than CDs, it is possible to see offsetting population moves across different skill groups, resulting in an overall positive correlation between immigrant and native-born population growth across CMAs. Similar to the method applied by Card and DiNardo (2000), I used PUMF (Public Use Micro Files) to estimate the net impact of immigration inflows on the relative skill distributions of different CMAs based on the following specification:[12]

$$(\Delta N_{sr} / P_{sr} - N_r / P_r) = \alpha + \beta(\Delta M_{sr} / P_{sr} - M_r / P_r) + e_{sr}, \qquad (3)$$

where the terms in parentheses are relative growth rates (*RGR*) of native-born and immigrant workers in skill group *s*, respectively.[13] Hence, equation (3) connects the response of native-born workers to changes in the relative supply of immigrants in their own skill group. Table 4 shows the estimated response of native-born workers to changes in the relative supply of immigrants in their own skill group across CMAs by *first* and *double difference* specifications.

[12] Only the CMA-level geographic classification is available in PUMFs. Hence, we do not have CD-level estimations. However, as the number of skill cells increases, a CD-level classification may result in insufficient or missing data for many CDs.

[13] How to build skill cells is an open question. The literature shows several different skill classification systems, but the most comprehensive method is that used by Card and DiNardo (2000). They define three skill cells and find the probabilities of being in one of those cells based on each individual's human capital characteristics. Because applying a similar method is beyond this paper's scope, we created skill groups in five education levels (1-high school dropouts or lower degree, 2-secondary, high school or non-university diploma, 3-university diploma under a bachelor degree, 4- bachelor degree, and 5- graduate degree) and ignored additional classification by years of experience.

Table 4: Estimated Responses of Native Born to Immigrants

Education Level (2)	**4(a)**		**4(b)**	
	β	*SE*	*β*	*SE*
RGRm (91,06)	-1.000	0.237		
RGRn (91,96)			0.473	0.648
RGRm (96,06)			-1.068	0.238
RGRm (91,96)			-3.730	1.751
# of observations	17		17	
R^2	0.268		0.440	
Education Level (5)				
RGRm (91,06)	0.1393	0.0845		
RGRn (91,96)			0.271	0.370
RGRm (96,06)			0.168	0.053
RGRm (91,96)			-0.192	0.263
# of observations	17		17	
R^2	0.180		0.335	

Notes: (1) For all estimations, the weighted least square (WLS, weight=total labour force) method with robust standard errors (SE) is used. (2) The dependent variables are *RGRn*(91,06) for 4(a) and *RGRn*(96,06) for 4(b), as defined in the text above, where *n* and *m* represent native-born and immigrant workers, respectively.

Table 4 contains the estimation results of (3). I display only two education levels where the results are significantly different from zero. As seen in the two columns, *first* and *double difference* (unrestricted version) methods do not produce significantly different outcomes, which indicates that in large labour markets, it is quite likely to have offsetting population movements across different skill groups. Again, an unbiased estimation of (3) requires that the allocation of immigrants across regions is random, which is unlikely. Hence, the reliability of these results depends on the magnitude of simultaneity and unobserved heterogeneity problems despite the fact that relative growth rates and double differencing provide some level of remedy.

4. A SPATIAL EQUILIBRIUM FRAMEWORK

To identify the key channels by which immigration affects out-migration of nonimmigrants, I introduce a simple spatial equilibrium model in which population flows are explained by regional housing and labour market conditions, the quality of local amenities, and the presence of social avoidance and/or self-selected ethnic segregations.[14] To understand how relocation decisions can be made by nonimmigrant individuals, I start with the following separable utility function:

$$U_{ir} = \left[\alpha \frac{h^{1-\theta} - 1}{1-\theta} - R_r h\right] + w_r(P_r) + A_{ir} - \lambda M_r . \qquad (4a)$$

The term in brackets captures the net income effect of housing, where h is the value received from housing services that the nonimmigrant individual i consumes and R represents the housing rent in region r so that the optimal housing can be expressed as $h_{ir} = (\alpha/R_r)^{1/\theta}$. The individual earns and consumes the current regional wage w_r, which is a function of the local population (P_r)—the sum of nonimmigrant (N_r) and immigrant (M_r) residents. Assuming that all

[14] This model builds on the framework used by Saiz (2007) to identify the relationship between immigration and local housing markets in the U.S.

individuals are in the labour force, the demand for labour can be expressed by $w_r = \bar{w}_r - \rho(N_r + \varepsilon M_r)$, where ρ measures the impact of population growth on local wages and ε reflects the degree of substitutability between immigrant and nonimmigrant workers with $0 \le \varepsilon \le 1$.[15]

The value of regional amenities for the individual is represented by the term *A*, which is heterogeneous among established nonimmigrants and new immigrants. Hence, $A_{ir} = T_r - aN_r$ provides a linear approximation of the congestion effect, where *T* is the total capacity of local amenities, and *a* is the individual share of amenities that each nonimmigrant consumes on average.

Finally, immigration has a direct negative impact on the well-being of nonimmigrants expressed by λM, where λ captures the degree of self-segregation and/or social avoidance. This shows that the native-born and established immigrants may want to live in communities with other households who have similar cultural and social values. This topic has been the subject of discussions particularly in the U.S. Frey used terms such as “demographic balkanisation” and “white flight” in his earlier articles investigating racial segregation across cities in the U.S. Further, Filer (1992) found that although white wages are affected less by low-skilled immigration than are black wages in the U.S., mobility responses were stronger among whites, which implies that something other than direct labour market effects influence the native born’s migration.

[15] We assume that $\rho > 0$.

At the steady state, for a spatial equilibrium to hold, U_{ir} must be equal to a reservation utility level denoted by $\underline{U}$. In other words, the marginal nonimmigrant will be indifferent between staying and leaving the region if $U_{ir} = \underline{U}$, where I normalize the utility level outside of the region to $\underline{U}$.[16] From this spatial equilibrium condition, we can derive the supply of nonimmigrant residents in region *r* as follows:

$$N_r = \Phi - \frac{1}{a+\rho}\left((\underline{T} - T_r) + (\underline{\overline{w}} - \overline{w}_r) + \sigma(\underline{R}^{\frac{\theta-1}{\theta}} - R_r^{\frac{\theta-1}{\theta}}) \right) - \left(\frac{\varepsilon\rho + \lambda}{a+\rho} \right) M_r, \quad \text{(4b)}$$

where $\Phi = (a+\rho)^{-1}(\varepsilon\rho+\lambda)\underline{M} + \underline{N}$ and $\sigma = \theta\alpha^{1/\theta}/(1-\theta)$. I assume that immigration to a region from abroad is exogenously determined by conditions in source countries and previous immigrant inflows.

A number of observations can be made based on (4b). First, unfavourable spatial differences (*e.g.*, in amenities, income levels, and housing costs) expressed in the first parenthesis has a negative effect on the number of nonimmigrant residents.[17] Second, the effect of immigration on nonimmigrant mobility is not

[16] To incorporate moving costs, we can discount $\underline{U}$ by μ; hence, the equilibrium condition becomes $U_{ir} = \mu\underline{U}$, where $0 \leq \mu \leq 1$ depending on the distance. Although discounting does not change the fundamental results, it implies that spatial differences and immigration have weaker effects on relocation decisions as the distance increases.

[17] The list of spatial differences that may affect the supply of nonimmigrant residents should be longer than what we have in (4b). For example, as Ley (2007) pointed out, in the last two decades, global cities have been experiencing significant shifts in their economic structure with new types of capital growth and polarized labour demand. This ongoing economic restructuring has created a new demand for managerial and professional occupations with the resulting decreasing importance of primary jobs in some disappearing manufacturing and service industries.

independent of the parameter of ρ—the impact of immigration on local wages. For example, if ρ is "large" in the case of imperfect substitution, the effects of immigration on internal migration are diffuse even if the direct disutility from immigration is substantially larger than zero. Third, in this framework, the labour market is clear, and there is no unemployment. If wages do not adjust, internal migration will be determined by the inter-regional differences in unemployment rates. Lastly, the coefficient on M is biased downward because the relation between local housing rents and immigration is absent in the above setting.[18]

Note that by using (4b), we can also express differences in the supply of nonimmigrant residents between region j and r as a function of spatial differences and the divergence in the size of immigrant populations as follows:

$$N_j - N_r = \frac{\Omega}{a+\rho} + \left(\frac{\varepsilon\rho + \lambda}{a+\rho}\right)(M_j - M_r), \tag{4c}$$

where Ω represents the term in the first parenthesis of (4b).[19] Based on (4c), the estimating framework takes the following form:

As a result, blue collar native-born workers have had to move to different places or upgrade their skills, whereas new immigrant workers are willing to work for low wages (Sassen, 1995).

[18] From (4b), it is obvious that an increase in R reduces N. It can also be shown that at any given level H (= h x P) and N, an increase in M raises R in the short run.

[19] See Hatton and Tani (2005) for a nonimmigrant labour-supply function in terms of growth rates. Moreover, (4c) puts a constraint on the coefficient of (M_j–M_r) by assuming that the degree of substitutability (ε) and the disutility from immigration are identical across regions, which may not be an unrealistic assumption for neighbouring regions. When we relax it, (4c) becomes

$$N_j - N_r = \Omega(a+\rho)^{-1} - [(\varepsilon_j\rho + \lambda_j)M_j - (\varepsilon_r\rho + \lambda_r)M_r](a+\rho)^{-1}$$

$$\Delta \ln(N_{rt}) = \beta \Delta \ln(M_{rt-1}) + \delta_k \mathrm{X}_{rt-1} + \alpha \mathrm{Z}_r + u_r + \tau_t + e_{rt}, \qquad (5)$$

where the dependent variable is the growth rate of nonimmigrant residents measured by the change in the log of *N*. Because I expect that the adjustment of the nonimmigrant population may not be contemporaneous in response to immigrant inflows and changes in local conditions, I use lagged values for the growth rate of immigrant residents, $\Delta\ln(M_{rt-1})$, and other explanatory variables, $\mathrm{X}_{rt\text{-}1}$, 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*.[20]

Following the literature, first, I predict that all things being equal, people tend to move to regions where housing costs are lower. As proposed by Poterba (1991), in equilibrium, the expected cost of owning a house should equal the cost of renting.[21] Hence, I use two alternative variables to control regional housing costs: gross average monthly rents for residential properties and average housing prices. Second, I control the linkage between labour market outcomes and population mobility by changing in the region's unemployment rate. My

[20] 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.

[21] User cost = $R = P(r + \tau + \lambda - \pi)$, where R is the imputed gross rent, P is the housing price, r is the cost of foregone interest that the homeowner could have earned on an alternative investment, τ is the property tax rate, λ is the recurring holding cost consisting of depreciation, maintenance, and the risk premium on residential property, and π is the expected capital gains (or loss).

prediction is that rising unemployment will be associated with declining growth in the nonimmigrant population. Finally, to remove unobserved spatial differences, I first apply the following fixed effects model that includes a full set of year dummies:

$$\Delta \ln N_{rt} - \overline{\Delta \ln N_r} = \beta(\Delta \ln M_{rt-1} - \overline{\Delta \ln M_r}) + \phi_k(\Delta \ln X_{rt-1} - \overline{\Delta \ln X_r}) + \tau_t + (e_{rt} - \overline{e}_r), \quad (6)$$

where variables with bars represent panel-level averages for their respective periods. By its construction, (6) has explanatory power only if the variations in $\Delta\ln N$ around its mean are significantly correlated with the variations in $\Delta\ln M$.

The estimation results of (6) with and without lagged explanatory variables are provided in Table 5. The estimations in the first column, 5(a), which use non-lagged explanatory variables show counterintuitive results, and, when compared with the results in 5(b), which use lagged variables, appear sensitive to the use of lagged covariates, implying that the contemporaneous link between native-born and immigrant population growth rates can be subject to a simultaneity problem. Moreover, test results confirm that unobserved regional fixed effects are not random; therefore, although I use growth rates in (5), the use of specification (6) instead of a random effect model is justified.

Except for 5(a), the fixed effect estimators consistently show statistically significant and negative relationships between the growth rates of nonimmigrant 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. This could be explained by the fact that the dependent variable represents the growth in the nonimmigrant population, not in the number of 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 nonimmigrant residents. By definition, (4a) 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. For example, consider two neighbouring regions with identical amenities, one of which has annual wages of $40,000 with annual housing costs of $10,000, while the other has annual wages of $60,000 and annual housing costs of $30,000. As seen in this example, even though the housing cost is higher in the second region, higher wages offset the difference, and therefore in both regions, people earn the same income ($30,000, net of housing cost). This interpretation leads to a downward bias in the coefficients of the housing variables in the estimations.[22]

[22] See Glaeser (2008) for more details.

Table 5: Fixed Effect Estimators with Growth Rates

CD	**5(a)**		**5(b)**		**5(c)**	
	Coefficent	SE	Coefficent	SE	Coefficent	SE
$\Delta \ln M$	0.013	0.012				
$\Delta \ln(un)$	0.037	0.003				
$\Delta \ln(hprice)$	0.113	0.048				
L.$\Delta \ln M$			-0.011	0.004	-0.013	0.005
L.$\Delta \ln(un)$			-0.021	0.008	-0.018	0.012
L.$\Delta \ln(hprice)$			-0.011	0.033		
L.$\Delta \ln(rent)$					-0.038	0.020
No of observations (t, r)	(3, 234)		(2, 234)		(2, 234)	
R2 (within)	0.2052		0.1633		0.1801	
rho	0.7111		0.7993		0.8047	
CMA						
$\Delta \ln M$	0.222	0.040				
$\Delta \ln(un)$	0.043	0.034				
$\Delta \ln(hprice)$	0.104	0.044				
L.$\Delta \ln M$			-0.155	0.049	-0.169	0.070
L.$\Delta \ln(un)$			-0.056	0.021	-0.053	0.025
L.$\Delta \ln(hprice)$			-0.024	0.048		
L.$\Delta \ln(rent)$					-0.016	0.042
No of observations (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 $\Delta \ln N$. (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 (not shown here). (5) L in front of the variable indicates the lagged value of the variable.

5. SENSITIVITY OF INTER-REGIONAL MIGRATION TO INTERNATIONAL MIGRATION

Instead of using population growth rates to understand whether immigration induces nonimmigrant residents to relocate to other regions, I now use inter-regional and international mobility measures for local residents. I estimate the following version of (6):

$$\ln I_{rt} - \overline{\ln I_r} = \beta(\ln E_{rt-1} - \overline{\ln E_r}) + \varphi_k(\ln X_{rt-1} - \overline{\ln X_r}) + \tau_t + (e_{rt} - \bar{e}_r), \qquad (7)$$

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. Hence, specification (7) reveals the sensitivity of inter-regional mobility to immigration when the regional housing and labour market conditions are controlled. If the displacement effect of immigration is significant, I expect that regions receiving more residents from abroad will attract less internal migration.

Table 6 shows the estimation results of (7) with and without lagged explanatory variables 6(a) and 6(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, rising unemployment rates now have significant and negative effects on the number of local residents who moved from a different region in the last five years. This is perhaps because the dependent variable may now include more people 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 due to this bias, as explained before.

Tables 5 and 6 results both indicate significant geographical variations in the observed spatial correlation between the native-born and immigrant growth rates. For example, in Table 5, the coefficients of interest show that a 10 percent increase in the immigrant growth rate across CMAs decreases the native-born population growth rate around 1.5 percent, whereas the same decline is roughly 0.1 percent across CDs. In other words, the results imply that approximately 15 fewer native-born people will choose to live in a particular CMA (CD) in the next five years for every 100 new immigrants entering the CMA (CD) today.[23] Given that almost 80 percent of the immigrant population is located in CMAs, this can be interpreted that the displacement effect of immigration becomes dispersed as the immigration population spreads out more.

[23] Assuming that a 10 percent increase in the immigrant growth rate represents 100 more immigrants.

Table 6: Fixed Effect Estimators with Mobility Measures

CD	**6(a)**		**6(b)**		**6(c)**	
	Coefficient	SE	Coefficient	SE	Coefficient	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
No of observations (*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
No of observations (*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 ln*I*. (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.

6. CONCLUDING REMARKS

This study is the first in the literature to investigate nonimmigrant mobility responses to immigrant inflows in Canada on a national scale. It used data from four population censuses covering the period 1991-2006, a unique period in recent Canadian immigration history. A panel was built over four population censuses by using the concordance tables to obtain stable borders at the CMA and CD levels through time. To my knowledge, no one has yet used the CD-level spatial scale to investigate population flows in Canada. Migration is a phenomenon mostly observed in neighbouring regions. Therefore, using the smallest available geographic classification in a balanced panel is important as larger regional sizes mask local population flows.

First, a model applied by Borjas et al. (1997) was estimated. At the CD level, the results implied that a 1 percent increase in the growth rate of the immigrant population reduces the nonimmigrant population growth rate by 1 percent. At the CMA level, however, the results did not indicate a negative link. This may indicate that immigration may be causing the out-migration of the native born within a CMA and/or skill groups may have offsetting mobility responses in the same CMA. In the second part of the paper, a spatial equilibrium model was introduced that identifies the channels by which immigration may affect nonimmigrant population flows. When panel estimations were applied, the results consistently showed a negative and robust impact on the location decisions of nonimmigrants.

The estimated coefficients on immigration variables were not independent of the elasticity of labour demand, the substitution between immigrant and nonimmigrant workers, and the effect of immigrants on housing costs, even though housing and labour market conditions were controlled and unobserved spatial differences removed. Therefore, a negative coefficient substantiates its existence but cannot provide an unbiased magnitude of social avoidance and/or self-segregation.

Appendix: A note on boundary changes of CDs and CMAs across the 1991, 1996, 2001, and 2006 censuses

Data from the last four population censuses (1991, 1996, 2001, 2006) were used to generate a panel. In each census, there are three main standard geographic classifications (SGC): provinces or territories (PR), census divisions (CD), and census subdivisions (CSD). Because CD boundaries tend to be more stable over the years than CSDs, CDs were first used to pool all four censuses. To make each CD consistent in each census, CDs whose coverage was affected by boundary changes were identified by using concordance tables and removing them from each census.[24]

In total, 51 CDs were removed. Their codes are 1010, 1011, 1301, 1305, 2419, 2424, 2425, 2426, 2436, 2437, 2442, 2443, 2444, 2449, 2451, 2452, 2455, 2457, 2458, 2459, 2561, 2462, 2477, 2479, 3512, 3514, 3534, 3539, 3548, 3552,

[24] See Statistics Canada, http://www.statcan.gc.ca/subjects-sujets/standard-norme/sgc-cgt/geography-geographie-eng.htm.

3553, 3557, 4615, 4616, 4803, 4812, 4814, 4815, 4816, 4818, 5903 5909, 5911, 5913, 5915, 5923, 5925, 5933, 5939, 5947, and 5949.

After this adjustment, the residual number of CDs was 234.[25] The land areas for each CD in each census also were compared, and I noticed that in the remaining CDs (where I thought no boundary changes had been made), the land areas (square kilometers) changed across censuses. While between 2001 and 2006 differences in size were minor, such was not the case between 1991, 1996, and 2001. Statistics Canada provided the following response to my question regarding area changes:

> Users should note that even when the boundaries of standard geographic areas did not change between censuses, the land areas might differ due to geometry shifts. The shifts are caused by a change in the underlying database architecture and by improvements in the absolute positional accuracy of some of the roads.

Therefore, even though it was possible to control boundary changes in the estimations, geometric shifts within a CD can result in some measurement errors in pooled data.

[25] We also removed CDs from the Northern Territories.

Statistics Canada uses, in addition to the SGC, a number of other standard geographic entities (*e.g.*, census metropolitan areas and census agglomerations, economic regions, health regions, and countries) to provide statistics. The general concept of these standard units is defined as "the urban core, and the adjacent urban and rural areas that have a high degree of social and economic integration with that urban core, as measured by commuting flows derived from Census of Population data on place of work" by Statistics Canada.[26]

To form a census metropolitan area (CMA), the urban core must have a population of at least 50,000 and the area (CMA) a population of at least 100,000. Once an area becomes a CMA, it is retained as a CMA even if the population of its urban core declines below 50,000 or if its total population falls below 100,000.

Since new CMAs emerge in each census, to be able to build a balanced panel, I took the CMA list in 1991 as my base list and removed all new additions from the subsequent censuses. The resulting list became as follows:

[26] Statistics Canada, http://www.statcan.gc.ca/subjects-sujets/standard-norme/sgc-cgt/2006/2006-intro-fin-eng.htm.

Code	CMA	Code	CMA
1	St. John's (N.L.)	559	Windsor (Ont.)
205	Halifax (N.S.)	562	Sarnia (Ont.)
305	Moncton (N.B.)	575	North Bay (Ont.)
310	Saint John (N.B.)	580	Greater Sudbury / Grand Sudbury (Ont.)
408	Saguenay (Que.)	590	Sault Ste. Marie (Ont.)
421	Québec (Que.)	595	Thunder Bay (Ont.)
433	Sherbrooke (Que.)	602	Winnipeg (Man.)
442	Trois-Rivières (Que.)	705	Regina (Sask.)
462	Montréal (Que.)	725	Saskatoon (Sask.)
505	Ottawa - Gatineau (Ont./Que.)	810	Lethbridge (Alta.)
521	Kingston (Ont.)	825	Calgary (Alta.)
529	Peterborough (Ont.)	830	Red Deer (Alta.)
532	Oshawa (Ont.)	835	Edmonton (Alta.)
535	Toronto (Ont.)	915	Kelowna (B.C.)
537	Hamilton (Ont.)	925	Kamloops (B.C.)
539	St. Catharines - Niagara (Ont.)	932	Abbotsford (B.C.)
541	Kitchener (Ont.)	933	Vancouver (B.C.)
543	Brantford (Ont.)	935	Victoria (B.C.)
550	Guelph (Ont.)	970	Prince George (B.C.)
555	London (Ont.)		

Moreover, the following eight CMAs: 205, 521, 529, 543, 575, 602, 810, and 970 were identified, with substantial changes in their land sizes across years. After removing those eight CMAs, we had 31 CMAs.

The list of CMAs in Public Use Micro Files (PUMF) is different from the above list. Similar to the panel data, to be able to build a balanced panel, I took the CMA list in 1991 as my base list and removed all new additions from the subsequent censuses. After removing the same eight CMAs with border changes (listed above), 17 CMAs remained, as listed below:

Code	CMA	Code	CMA
421	Québec (Que.)	555	London (Ont.)
462	Montréal (Que.)	559	Windsor (Ont.)
433	Sherbrooke (Que.)	599	Thunder Bay (Ont.)
505	Ottawa - Gatineau (Ont./Que.)	799	Regina (Sask.)
532	Oshawa (Ont.)	825	Calgary (Alta.)
535	Toronto (Ont.)	835	Edmonton (Alta.)
537	Hamilton (Ont.)	933	Vancouver (B.C.)
539	St. Catharines - Niagara (Ont.)	935	Victoria (B.C.)

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