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Effects of Immigration on the Canadian Housing Market

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Introduction[1]

The purpose of this study is to assess the impact of immigration on housing prices in Canada. Econometric analysis is performed using data at census division levels based on 1996, 2001, and 2006 population censuses to assess the impact of immigration on prices of new privately owned dwellings.

The study is important for several reasons. First, Canada admits about 225,000 immigrants each year. Based on a housing demand projection study conducted by Canada Mortgage and Housing Corporation (1992), we estimate that a one percent rise in immigrant population causes housing demand in the country to rise by about 0.66 percent. However, a systematic analysis of the impact of immigrants' housing demand on housing prices is lacking.[2]

Second, housing sector in Canada accounts for about 41 percent of its construction industry which accounts for 6 percent of national industrial value-added. About 8 percent of employed Canadians work in construction industry. Canadian Housing and Renewal Association (2009) estimates that for each new home built in Canada, 4-6 person years of employment is generated. Thus, any impact of immigration on the demand for housing can also be viewed as an important component of the overall economic contribution of immigrant population in Canada.

Third, most Canadian immigrants go to Ontario which receives the highest number of them on per capita basis. However, in recent years, annual immigrant inflows per thousand population have fallen in Ontario and increased in other provinces. Manitoba, and tiny Prince Edward Island, experienced a remarkable rise in their immigrant inflow in 2007 on a per capita basis (Chart 1). This changing geographical distribution of immigrant arrivals in Canada is partially attributed to new regional policy and community initiatives adopted outside of Ontario to attract and retain immigrants, in order to reverse population declines in those regions. These initiatives also aim at increasing immigration in smaller, and sometimes rural, areas within each province to prevent an economic decline that could accompany a population decline. By analyzing data at the level of census divisions over three census years, the present paper can help shed some light on the effect of changing geographical distribution of immigrant population on housing prices in the country.

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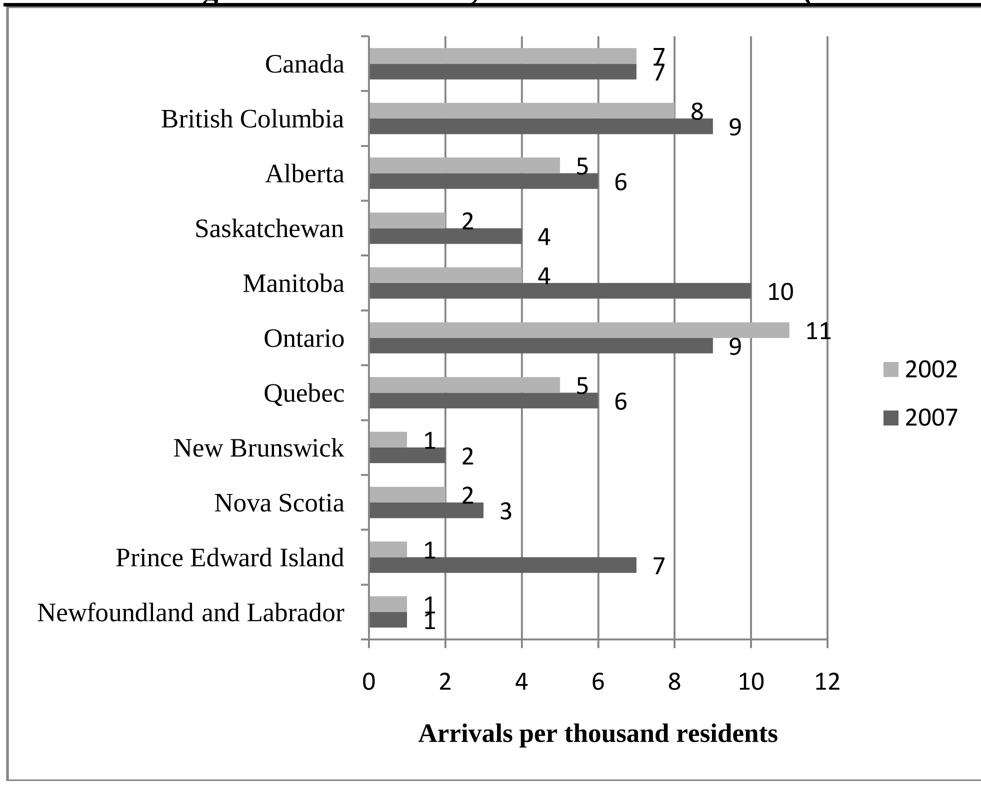
[2] Studies that have analyzed immigrants' purchase of accommodation in Canada have usually focused on the probability of buying a house using census micro data. For example, see Edmonston (2004).

Chart 1: Immigrant Arrival Rates, Canada and Provinces (2002 & 2007)



Source: Authors' calculations based on Citizenship and Immigration Canada (2007) and Statistics Canada (2009).

Finally, the sudden rise in house prices in Canada, the United States, and in Western Europe at the turn of present century led many economists to hold that in these countries housing sector was in a "bubble", meaning that the house prices were not sustainable.[3] Many attribute this bubble to the low interest rates, resulting from high liquidity, which attracted many low income buyers into the housing markets. The rate for rental accommodations remained flat during this period. In the United States, and in the United Kingdom, many public circles have also pointed out that immigration also contributed to the creation of housing bubble as many new immigrants purchased homes from their past savings and not through the money earned through employment in the host country. Many economists, for example, Paul Krugman (2007), predicted bursting of this bubble and this prediction turned out to be correct towards the end of the year 2008. Home buyers who had borrowed funds to pay for their homes have ended up with a mortgage value that is higher than the value of their homes. Some similar effect has been reported in the news media about Canadian housing market.[4] Since Canada is a major immigrant receiving country, and new immigrants represent a large pool of potential home buyers, it is important to investigate if immigrants in Canada also contributed to the creation of

[3] In 2001, the price of an average house in Canada was about 3.5 times more than the median household income; in 2006 it was about 5 times more. The average rent and average monthly payment for an owner occupied dwelling remained almost constant during this period. Interest rates remained below five percent. In the United Kingdom, the average house price rose to about 8 times the median income in 2007.

[4] For example, MacClean's magazine (February 14, 2009)

housing bubble. Results of present study will help shed some light on this effect of immigration in Canada.

Patterns of home ownerships in Canada

The last two Canadian census data show that home ownerships in Canada increased among both immigrants and non-immigrants (Table 1). Homeownership among immigrants who have stayed in the country for more than 20 years was higher than among non-immigrants. Fastest increase was observed among those immigrants who have stayed in the country between six to ten years. These findings corroborate with those of Edmonston (2004), who based his analysis on the 1991 and 1996 censuses.

Table 1: Percentage of persons in private occupied dwellings owned by a household member by number of years since immigration

Number of years since immigration	2001	2006	Difference
	percentage		percentage points
Canadian-born population	73.1	75.3	2.2
Total immigrants	68.2	71.6	3.4
5 years or less	39.9	45.7	5.8
6 to 10 years	58.9	66.7	7.8
11 to 20 years	68.2	72.7	4.5
21 to 30 years	77.8	78.9	1.1
31 to 40 years	83.8	84.2	0.4
More than 40 years	84.4	85.0	0.6

Source: Statistics Canada (2006).

Table 2 provides data on housing prices in seven selected Census Metropolitan Areas (CMAs). Also provided in the same table are data on their immigrant concentration, and homeownerships. The first three CMAs are traditional destinations of immigrants where about 80 percent new immigrants settle. The next two CMAs are among the new emerging destinations of immigrants, as identified by previous authors. To the list, we also add two more CMAs, Halifax and Winnipeg which have also experienced significant increase in their immigrant inflows since the turn of present century. For comparison purposes, we have also provided data for Canada in the same Table.

Table 2: Housing market and immigrant concentration in selected Census Metropolitan Areas (CMAs) and Canada, 2006.

CMA & Canada	Average price of a privately owned dwelling	Percentage of immigrants in total population	Population per privately owned dwelling	Construction workers per 100 privately owned dwellings
Montreal	244,417	21	4.41	11
Toronto	403,112	45	4.19	12
Vancouver	520,937	40	3.95	14

Calgary	381,866	24	3.48	17
Edmonton	265,030	19	3.65	20
Halifax	212,942	7	3.72	12
Winnipeg	168,255	18	4.08	10
Canada	263,369	20	3.67	12

Source: Statistics Canada, 2006 Census Community Profiles.

Several interesting observations can be made from these data. First of all, while Montreal, Toronto, and Vancouver are among traditional immigrant recipients, about 24 percent of Calgary's resident population is foreign born which exceeds that of Montreal where it is at 21 percent. More immigrants to Canada have declared Calgary and Edmonton as their homes since late 1990s due to the booming economy of Alberta. Except for Montreal, housing prices in major immigrant receiving CMAs (with immigrant population exceeding the national average of 20 percent) are above the national average. Population per dwelling data show that it is easier to find a house in Calgary, Edmonton and to some extent, in Halifax than it is nationally. It is also easier to find a construction worker in Calgary, Edmonton and in Vancouver where the number of construction workers per dwelling exceeds the national average. In sum, these results show some relationship between immigrant concentration by geography and housing prices in Canada. We aim to study this relationship more formally in an econometric model.

Empirical Analysis

Econometric estimates of a housing price model can be based on time-series and / or cross-sectional data. While time-series data can relate price variation to its sources over time, cross-sectional data can explain price differences among various housing markets by the differences in regional characteristics. The relevance of cross-sectional data arises from the fact that house prices are unlikely to experience the arbitrage of tradable divisible commodities, and so it is unlikely that the law of one price holds. However, even though differences in local market characteristics may explain differences in housing prices, single cross-sectional data sets may not provide unbiased and consistent econometric estimates if there are unobserved interregional differences inherent in comparisons of different regions. A panel data, on the other hand, allows us to control unobserved heterogeneity among regions by including time-series or cross-sectional data at the same time.

Our dataset is based on the last three population censuses of Canada conducted in 1996, 2001, and 2006. To conduct the population census, Statistics Canada divides the country into Census divisions (CD) or smaller communities which are intermediate geographic areas between the province / territory level and the municipality (census subdivision). There are 289 CDs across the country, which are provincially legislated. Each CD may be viewed as a group of neighbouring municipalities joined together for the purposes of regional planning and managing common services (such as police or ambulance services). Based on the information collected from population census, Statistics Canada releases on its web site, the socio-economic and demographic profiles of each CD. The present paper is the first one to use these community based data to construct a panel dataset in order to analyze the impact of immigration on housing markets in Canada.

One problem with using the above data is that the coverage of some CDs changed between 1996 and 2001 censuses. To solve this problem, we had two options: to file a special request with Statistics Canada to provide consistent data since 1996 census on all CDs, or to drop the CDs whose boundaries changed. Due to the exorbitant cost of the first option, we resorted to the second option which led us to drop 31 CDs from our analysis. Hence, our analysis is based on 258 CDs across Canada for 3 census years which cover 10 years starting in 1996. We believe that the cost of reduction in the periodicity by the use of five-year censuses is somewhat offset by the fact that the time gap between censuses allows house prices to fully adjust to changes in economic and demographic conditions .

The objective our empirical analysis is to understand the relationship between housing prices and immigration. More specifically, we want to answer the following two questions: (1) what is the expected difference in the average housing price between two locations when they differ in the composition of immigrants in their total populations? (2) What is the expected change in the price of an average house when the composition of immigration in total population changes over time? While the answer to the first question uses cross-sectional information and reveals the long-run relationship between average housing prices and immigration, the answer to the second question uses time-series information and tells us how much of the increase in average real housing prices over the last ten years can be attributed to changes in immigration.

In most empirical analyses, house price movements are modeled on the assumption that markets clear quickly and prices equate the demand for housing with the existing stock. Thus, equilibrium price levels can be determined by the housing stock and demand instruments, such as user cost, demographic characteristics, permanent income, opportunity cost of owning a dwelling, and other exogenous variables. The supply of housing, on the other hand, is assumed to exogenously depend on the level of construction cost. Hence, prices in housing markets can be expressed as a function of the existent housing stock and determinants of demand and supply of housing. Recent studies also incorporate expected future housing price inflation (Egert and Mihaljek 2008) as well as hedonic price determination methods (Bover and Velilla 2002) in their empirical analyses.

The effect of immigration on housing prices can be different than on rents. A large body of research has shown that most newcomers are tenants and immigrants who have lived in Canada more than a decade have above-average rates of homeownership (CMHC, 2007). In general, the impact of immigration on a market comes from a shift in demand. Two factors make this impact in rental markets larger in the short-run than in the long-run: first, it takes time for housing supply to respond to an unexpected higher demand; second, local population may not quickly adjust to a rising housing cost by moving out from the region. New supply of housing and a potential displacement of local population can lessen this impact in the long-run. In addition, as they settle in their destination cities, immigrants become home buyers and leave rental housing markets in the long-run. If this is true, the effect of immigration on average housing prices can be expected to be smaller in the short-run than in the long-run.

Given the pooled cross-section time-series data, we can define several models that arise from the following most general linear representation:

$$y_{it} = \propto + \sum_{k=1}^{k} x_{kit}\beta_{kit} + \epsilon_{it},\ i = 1,\dots,\mathrm{N},\ t = 1,\dots,\mathrm{T}$$

where N is the number of locations and T is the number of periods. It is a reduced-form linear model that underlines our empirical work where all variables are in logarithmic forms. The dependent variable is the average price of an owner occupied dwelling (*price*). The main independent variable is the ratio of total immigrant residents in a CD to that CD's population (*immratio*). The incidence of homeownership among immigrants varies with their duration of stay in the country. Hence, prices of average homes can differ across otherwise two identical regions if one has more settled immigrants than the other. To control for differences in immigrants' duration of stay among regions, we include two additional variables: the ratio of recent immigrants — the number of people immigrated in the last five years – to total resident immigrants (*imm5*), and the ratio of new immigrants — the number of people immigrated last year over the total immigrants *(newimm)*.

Immigration can also have a crowding out effect on the original residents of a CD through its effect on house prices, which in turn could weaken immigration's impact on housing demand.[5] Therefore, to control this possible displacement effect on original residents, we include two variables: the ratio of movers who were in a different CD a year ago to non-movers (*mobility*) and the ratio of movers who were in a different CD five years ago to non-movers (*mobility5*).

We use three labour market indicators in our model. These include, per capita income before taxes (*income*), unemployment rate (*un*), and labour force participation rate (*lfpr*). Home buyers form their expectations of future income, to ascertain future payments towards maintenance of their house, based on current income. Unemployment controls the effects of consumption smoothing and uncertainty at the same time. Therefore, its ex-ante impact on housing prices is undetermined. Since labour force participation decision reflects lifetime income expectations, it is also a part of permanent income. However, a variation in LFPR can also reflect the difference in the distribution of population between the seniors and the young members. A lower LFPR, for example, can be attributed to

[5] This displacement effect can also occur through labour markets: depending on the level of substitution between immigrant and local workers, an inflow of newcomers can reduce local real wages and force native workers to move out from the region.

higher income expectations or a retired community. Therefore, we do not have prior expectations on the sign of LFPR either.

As commonly accepted in the literature (see Poterba, 2000), housing markets represent two sources of demand at the same time: needs for accommodation and investment. When people purchase a house, the return on investment, which is the ratio of the imputed gross rent to the price of dwelling, becomes a function of the nominal interest rate (i), the property tax rate (τ), the depreciation rate on housing capital (δ), maintenance cost (m), the expected rate of inflation in housing prices (π^e), the owners marginal tax rate (θ), and the required risk premium (α) as expressed below:

$$Rent/Price = f(i, \theta, \delta, \alpha, \pi^e, m, \tau).$$

It is evident from the above expression that the price should respond to changes in the average rent and/or in the variables on the right hand side. We include gross average rents (*rent*) into our estimations. In order to control expected appreciation in housing prices, we use the ratio of number of new dwellings to total stock (*ndwell*). The census collects information on minor and major renovation needs of a house. Hence, we use the ratio of the number of houses that need minor renovations to those that need major renovations as a proxy for the maintenance cost (*maincost*) of the house. We also use the average number of rooms per dwelling to control for differences in property taxes in each census division (*room*).

When the market clears, the equilibrium price is determined by the amount of housing stock, for which we use the sum of owner occupied dwellings, rental properties, and band houses (*stock*). On the supply side, we use the ratio of the number of construction workers to the total labor force to proxy construction cost (*cwork*) in each CD. In addition, we include density of the region (*density*), which is the number of people living in a square kilometer, to control for supply factors related to land availability as well as demographic differences among regions. The age composition of the population is also an important factor that captures the differences in housing prices (Mankiw and Weil, 1989). We calculate the ratio of people in their prime home-buying ages (19-34) to the total population (*ratio2034*) for each CD. To control special legal circumstances due to aboriginal population in some divisions, we use the share of aboriginal population (*abor*). Finally, we deflated all dollar values (for *price*, *rent*, and *income*) to 1996 prices by using gross domestic product (GDP) deflators.

As stated earlier, we want to identify the changes in average housing price as caused: (1) by the changes in its determinants within each CD from 1996 to 2006 reflected in the time-series information and (2) by the changes in its determinants across CD reflected in the cross-sectional information. To detect time series variations, we specify a within fixed-effect model based on the following one-way fixed-effect structure:

$$y_{it} = X_{it}\beta_k + u_i + \epsilon_{it}, \qquad (1)$$

where X_{it} is a 1 x k vector of variables that vary over CD (subscript i) and time (subscript t), β is the k x 1 vector of coefficients on X, u_i represents division-specific time-constant unobserved heterogeneity, ϵ_{it} is the idiosyncratic disturbance term, which satisfies $\epsilon_{it} \sim IID(0,\sigma_\epsilon^2)$ for all t and i.[6] We assume that the division-level fixed effects are correlated with the included variables and allow for heterogeneity across CD's but limit the same to the intercept terms of the relationship.

A great advantage of the within fixed effect model (WM) is that the estimates are not susceptible to bias due to omitted panel-level covariates. Since our panel has only three periods, instead of using a conventional two-way within fixed effect model, we estimate a one-way within fixed effect model with the use of year-dummies (τ_t) to control time-specific effects.[7] The underlying model can be obtained by removing panel-level averages, $\bar{y}_i = (\frac{1}{T})\sum_{t=1}^{T} y_{it,}\ \bar{x}_i = (\frac{1}{T})\sum_{t=1}^{T} x_{it,}\ \bar{\epsilon}_i = (\frac{1}{T})\sum_{t=1}^{T} \epsilon_{it}$, from each side of (1) as follows:

$$y_{it} - \bar{y}_i = (X_{it} - \bar{X}_i)\beta + (u_i - u_i) + \tau_t + \epsilon_{it} - \bar{\epsilon}_i. \qquad (2)$$

The WM captures the source of variation within a census division over the period from 1996 to 2006 when unobserved fixed year-effects for each census are controlled. There is a price of applying (2), though: it removes not only unobserved but all time-invariant effects and uses the variation within covariates over time.

In the absence of random-effect model, the way to examine the variations in prices between census divisions by using the cross-sectional information in a panel data, though, is more subtle. Although the between-effect model (BM),

$$\bar{y}_i = \propto + \bar{X}_i\beta + u_i + \bar{\epsilon}_i, \qquad (3)$$

which uses panel-level means of dependent and independent variables could have an intuitive interpretation for β as the long-run elasticity of average housing prices, given the nature of our analysis, the estimator (3) produces biased results. This is because we expect a high correlation between unobserved fixed effects and regressors.[8] However, it is worth comparing the within and between estimators as it is apparent from the following model:

$$y_{it} = \propto + \bar{X}_i\beta_1 + (X_{it} - \bar{X}_i)\beta_2 + u_i + \tau_t + \epsilon_{it}, \qquad (4)$$

where changes in the average value of X for a census division can have a different effect from temporary departures from the average. In other words, the between estimator (3)

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[6] We will test the assumption that $E[\epsilon_{it}\epsilon_{is}] = 0$ for all $s \neq t$ for contemporaneous correlations across cross-sectional units as well as for serial correlation within cross-sectional units.
[7] Fixed-year effects take care of macrofactors such as changes in the rate of GDP growth, tax regulations, interest rates, economic and political conditions over the period between 1996 and 2006.
[8] We have applied the Hausman test to see the appropriateness of fixed-effect models. The test results strongly reject the null hypothesis that the random effect estimators are consistent, which is also supported by a high correlation between X and u, as shown in Table 1.

estimates β_1 by using cross-sectional information and the within estimator (2) estimates β_2 by using time-series information and neither estimates the other.

The estimation results of (2) are given in Table 3, which also includes the results of the first-difference specification and the estimation of (2) with an autoregressive transformation. We use a modified Wald statistics (see Green, 2000, p.235) to test for group-wise heteroskedasticity in the residuals for the within estimator. The test results strongly reject the null hypothesis of homoskedastic errors, as expected. As suggested by Stock and Watson (2006), when serially correlated errors are possibility in fixed-effect regressions applied for panel data with $T = 3$, even if its degree of freedom is adjusted, instead of using the conventional heteroskedasticity-robust variance matrix estimator, clustered (over entities) variance estimators should be used for hypothesis tests on β. Accordingly, we adjust standard errors for 258 clusters (CD's). As a usual practice, error terms in pooled cross-section time-series data models are assumed cross-sectionally independent. We test the assumption of spatial dependence among neighboring divisions and/or omitted unobserved common components for the within estimator by following the method developed by Paseran (2004) for short T and large N panel data sets. The test result suggests that the null hypothesis that census divisions are cross-sectionally independent should be

Table 3: Within and First-Difference Estimators (1996 — 2006)

	Within			Within with AR(1)			First-Difference		
Variables	β	SE	P>|t|	β	SE	P>|t|	β	SE	P>|t|
Per capita income (*income*)	0.394	0.126	0.002	0.374	0.132	0.005	0.194	0.097	0.045
Unemployment rate (*un*)	-0.039	0.028	0.165	-0.076	0.034	0.027	-0.020	0.024	0.404
Labour force participation rate (*lfpr*)	-0.006	0.197	0.978	-0.618	0.247	0.013	-0.051	0.178	0.775
Total stock of dwellings (*stock*)	0.545	0.150	0.000	-0.150	0.337	0.658	0.399	0.136	0.004
Average gross rent (*rent*)	0.287	0.086	0.001	0.405	0.095	0.000	0.186	0.062	0.003
Explicit user cost (*maincost*)	-0.049	0.039	0.216	-0.058	0.042	0.165	-0.012	0.032	0.706
Average number of rooms per dwelling (*room*)	0.758	0.227	0.001	0.995	0.284	0.001	0.707	0.207	0.001
New construction/Total stock of dwellings (*Ndwell*)	0.280	0.032	0.000	0.289	0.042	0.000	0.214	0.031	0.000
Construction workers/Total labour force (*cwork*)	0.191	0.034	0.000	0.212	0.038	0.000	0.133	0.031	0.000
Population/Size (*density*)	-0.062	0.120	0.606	0.763	0.308	0.014	0.092	0.093	0.322
People between 20 and 34/Population (*ratio2034*)	-0.103	0.084	0.224	0.066	0.141	0.642	0.066	0.064	0.303
Aboriginal population/Population (*abor*)	-0.001	0.007	0.883	0.010	0.015	0.481	0.004	0.006	0.518
People in different CD's a year ago/Non-movers (*mobility*)	0.028	0.034	0.412	0.052	0.048	0.280	-0.040	0.038	0.292
People in different CD's 5 years ago/Non-movers (*mobility5*)	-0.065	0.052	0.216	0.056	0.055	0.310	0.129	0.040	0.001
Total immigrants/Population (*immratio*)	0.005	0.022	0.832	0.031	0.029	0.281	0.010	0.017	0.569
Fixed effects —Year	Yes			Yes			Yes		
Fixed effects —CD	Yes			Yes			No		
Fixed effects —Province	No			No			Yes		
AR(1)—*rho*				0.212					
# of observation, groups, observation per group	774, 258, 3			774, 258, 3			516		
F(df, N)	19, 270	144.5		16, 242	119.7		25, 490	80.3	
R2-within, between, overall	0.854	0.558	0.568	0.075	0.937	0.833	0.777		

Notes: (1) All models are estimated by OLS where the dependent variable is average prices for owner occupied dwellings. (2) We used clustered (by CD) standard errors (SE) for the within estimator and robust SE for the first-difference estimator. (3) rho in the within estimators indicates the fraction of the unexplained variance due to differences among CD's. (4) rho in AR(1) transformation is estimated by the Durbin-Watson *d* statistics.

accepted with the probability of 0.3216. In addition, we use the method applied by Wooldridge (2002) to test the assumption that $E[\epsilon_{it}\epsilon_{is}] = 0$ for all $s \neq t$ for each cross-sectional unit i so that $\epsilon_{it} \sim IID(0,\sigma_{i\epsilon}^2)$ for all t for the within estimator. The null hypothesis of no first-order autocorrelation is strongly rejected with and without time-effect dummies. To address a possible spurious regression problem, we apply AR(1) transformation to the within estimator (2).

First of all, a high rho value estimated in the WM model show that almost all the variations in average house prices are explained by unobserved regional differences in housing markets, such us schooling quality, access to health care, crime, climate culture, etc. Moreover, a high correlation between fixed-effects and covariates (0.729) justifies the choice of fixed-effect model over the random effect model. The coefficient of the immigration ratio is not statistically different than zero, when unobserved regional characteristics are controlled. By its construction, the WM model explains how the change in y is related to the change in x. In this sense, the results of this model reveal short-run impact of immigration on prices and imply that immigration was not a source of rising average housing prices during the 1996 to 2006 period. This result does not change when we control for the possibility of serial correlation problem in WM model by using a first-order autoregressive transformation.

Changes in the immigration ratio around its mean over the last ten years come from two sources: changes in total population and the number of immigrants. In other words, the WM model actually captures the relationship between variations in average housing prices and recent immigrants. If we consider the fact that most newcomers are tenants and only after ten years they catch up with natives in terms of average homeownership rates, the results confirm that new immigrants do not have strong effects on housing markets in their first decade of stay in Canada. This result is also confirmed by the first-difference estimator, which shows the growth rate in the immigration ratio—which comes from recent immigrants—does not have a significant explanatory power on the growth rate of average housing prices between 1996 and 2006.

If we assume that marginal prosperity to purchase a house for immigrants is not different from that for non-immigrants, one can argue that the insignificance of immigration ratio may come from the crowding-out effect of immigration. This is because when the number of non-immigrants moves out from the census division,[9] the net effect of immigration on population would be smaller—which diminishes the shift in the housing demand. Consequently, even if the immigration ratio increases, average prices do not rise enough. In fact, if this displacement effect is true, non-immigrants may move to less immigrated regions so that the immigration ratio declines and prices go up due to the shift in demand by increasing population. In other words, if there is a perfect displacement, unless non-immigrants move out from the country, the ultimate effect of displacement may even indicate a negative relationship between the immigration ratio and average prices. The magnitude of displacement effect depends on how mobile the

[9] This argument also refers to whether or not immigrants are perfect substitutes of non-immigrants in labour markets, which is one of the longstanding arguments in the labour economics literature (see Borjas, 1994).

local population is. Even though two variables in our regressions, *mobility* and *mobility5*, may control a possible displacement effect in each region, it is important to note that the coefficient of the immigration ratio represents a net value after a possible crowding-out effect takes place.

As the within estimator indicates, the correlation between fixed-effects and dependent variables is high and unless we control unobserved heterogeneity, estimations that use only cross-sectional information will be biased and inconsistent. We attempt to minimize this problem in the between estimator by controlling unobserved fixed effects at the provincial level. As shown in Table 4 below, we specify two different estimations. In the first one, we control the share of recent immigrants in total immigrant population by two variables: new immigrants (*newimm*) and people who immigrated in the last five years (*imm5*). In the second estimation, we use two immigration ratios—the share of people immigrated in the last 10 years in the local population (*imm10down*) and the share of people who immigrated earlier than 10 years in the local population (*imm10up*)—to see whether the impact of recent immigrants on housing market is different than more established immigrants.

The results show that the elasticity of average housing prices with respect to the immigration ratio is 0.122 and statistically significant. When we estimate the same model with two different ratios, *imm10down and imm10up*, the results show that recent immigrants have no impact on average prices and the share of more established immigrants in the population have a significant explanatory power on price differences among regions. This result is also consistent what the WM indicates: the variation in the immigration ratio that comes from recent immigrants does not have an explanatory power on price changes in housing markets.[10]

[10] We also applied a least-square-dummy-variables (LSDV) model to individual censuses by controlling provincial level fixed effects. Consistent with the between estimators, the results show that recent immigrants have no effect on housing prices and the elasticity of housing prices with respect to the immigration ratio for more settled immigrants s around 0.14 and significant.

Table 4: Between Estimators (1996 — 2006)

	1			2		
Variables	*β*	SE	P>|z|	*β*	SE	P>|z|
Per capita income (*income*)	0.596	0.192	0.002	0.579	0.170	0.001
Unemployment rate (*un*)	-0.101	0.039	0.010	-0.094	0.033	0.005
Labour force participation rate (*lfpr*)	-0.823	0.254	0.001	-0.793	0.237	0.001
Total stock of dwellings (*stock*)	-0.013	0.021	0.552	-0.009	0.020	0.647
Average gross rent (*rent*)	0.665	0.183	0.000	0.663	0.128	0.000
Explicit user cost (*maincost*)	0.030	0.066	0.651	0.034	0.085	0.690
Average number of rooms per dwelling (*room*)	-0.322	0.231	0.164	-0.309	0.243	0.205
New construction/Total stock of dwellings (*Ndwell*)	0.179	0.031	0.000	0.190	0.028	0.000
Construction workers/Total labour force (*cwork*)	0.009	0.027	0.739	0.033	0.016	0.040
Population/Size (*density*)	0.024	0.009	0.008	0.025	0.009	0.004
People between 20 and 34/Population (*ratio2034*)	0.028	0.108	0.795	0.029	0.123	0.812
Aboriginal population/Population (*abor*)	-0.006	0.012	0.588	-0.007	0.013	0.564
People in different CD's a year ago/Non-movers (*mobility*)	0.055	0.071	0.441	0.027	0.082	0.743
People in different CD's 5 years ago/Non-movers (*mobility5*)	-0.123	0.058	0.034	-0.083	0.073	0.252
Total immigrants/Population (*immratio*)	0.122	0.024	0.000			
People residing abroad a year ago/Total immigrants (*newimm*)	-0.013	0.014	0.351			
People immigrated last 5 years/Total immigrants (*imm5*)	0.018	0.016	0.237			
People immigrated last 10 years/Population (*imm10down*)				0.015	0.018	0.391
People immigrated earlier than 10 years/Population (*imm10up*)				0.091	0.024	0.000
Fixed effects —years	No			No		
Fixed effects —CD	No			No		
Fixed effects —Province	Yes			Yes		
# of observation, groups, observation per group	774, 258, 3			774, 258, 3		
	Wald (26)		9750.35	Wald (25)		4582.73
R2-within, between, overall	0.075	0.937	0.833	0.140	0.935	0.843

Notes: (1) The models are estimated by OLS where the dependent variable is average prices for owner occupied dwellings. (2) We used bootstrap standard errors (SE).

As it is evident from the WM, removing unobserved regional heterogeneity reduces the impact of immigration on housing price. Hence, the coefficients of immigration ratio in the between estimators model actually overestimate the true values. This implies that average prices are not responsive to changes in immigration in the long-run so that a 1 per cent difference in the immigration ratio explains only around 0.10 (0.12 in (1) and 0.09 in (2)) per cent difference in prices among regions.[11]

The between estimators model is rarely used in panel analyses due to its obvious flaw in the presence of unobserved heterogeneity. It, however, provides long-term observation on the subject by using cross-sectional information, which is not offered by the within or first-difference estimator models. In order to see whether immigration, as one of the regional characteristics, explains differences in average housing prices among different locations, we develop the following model which removes unobserved heterogeneity but provides information about the long-run relationship between prices and immigration by using the cross-sectional information in the data:

$$\Delta\ln(price_{it}) = \left(\frac{imm_{it}}{pop_{it-1}}\right)\beta_1 + \ln(\overline{immratio}_i)\beta_2 + \Delta\ln(Z_{it})\beta_3 + \ln(W_{it-1})\beta_4 + \rho_i + \tau_t + \Delta\epsilon_{it}, \quad (5)$$

where the dependent variable is the growth rate of average housing prices, which is a function of growth in immigration with other region-specific attributes. The expression in the first parenthesis on the right hand side of (5) is a proxy for the growth rate of immigration for each CD where *imm* and *pop* are the number of immigrants immigrated in the last five years (between censuses) and the total population, respectively. By its structure, this model resembles to one specified by Saiz (2007) with an addition that we also want to observe whether the differences in the immigration ratio accounts for differences in the percentage change in average prices among regions. Therefore, we consider the second variable, *immratio*— the average of the immigration ratio over the ten years for each CD—as one of the region-specific attributes, like schooling, crime, climate, and so on. *Z* stands for a vector of growth rates of local income and average rent. *W* represents a vector for lag values of *un*, *lfpr*, *stock*, *majorcost*, *room*, *Ndwell*, *cwork*, *density*, *ratio2034*, *abor*, *mobility5*, as defined previously. We use lag values for the reason that prices may not adjust instantaneously to changes in

[11] However, as discussed before, immigration may be associated with an offsetting crowding-out effect. Therefore, the coefficient of the immigration ratio may underestimate its true value.

fundamentals. P and τ are dummies for provinces and years respectively. Taking the first difference in the price series removes unobserved heterogeneity among different locations. The same transformation also yields stationarity in price, income and rent series.

Table 5: Estimation Results of (5)			
Variables	β	SE	P>|t|
Δ Log per capita income (*income*) at T	0.251	2.290	0.023
Δ Log average gross rent (*rent*) at T	0.286	0.096	0.003
People immigrated in the last 5 years at T/Population at T-1 (*imm5/pop*)	-0.008	0.005	0.094
Average of (total immigrants/Population) (*immratio*)	0.036	0.009	0.000
Unemployment rate (*un*) at $T-1$	-0.026	0.019	0.169
Labour force participation rate (*lfpr*) at $T-1$	-0.174	0.111	0.117
Total stock of dwellings (*stock*) at $T-1$	-0.020	0.010	0.045
Explicit user cost (*maincost*) at $T-1$	0.031	0.033	0.360
Average number of rooms per dwelling (*room*) at $T-1$	-0.084	0.133	0.529
New construction/Total stock of dwellings (*Ndwell*) at $T-1$	0.034	0.013	0.009
Construction workers/Total labour force (*cwork*) at $T-1$	0.023	0.008	0.009
Population/Size (*density*) at $T-1$	0.008	0.005	0.108
People between 20 and 34/Population (*ratio2034*) at $T-1$	0.012	0.050	0.816
Aboriginal population/Population (*abor*) at $T-1$	-0.002	0.005	0.753
People in different CD's a year ago/Non-movers (*mobility*) at $T-1$	0.071	0.031	0.022
People in different CD's 5 years ago/Non-movers (*mobility5*) at $T-1$	-0.038	0.031	0.233
Fixed effects — Year		Yes	
Fixed effects — Province		Yes	
# of observations	516		
F (26, 257)	55.120		
R2-Adjusted	0.688		

Notes: (1) The model is estimated by OLS where the dependent variable is growth rate of average prices for owner occupied dwellings. (2) Standard errors are clustered by CD.

Table 5 above shows the results of (6) estimated by OLS with standard errors clustered by CD. It is interesting to see that β_1 is negative and statistically significant at 10 percent level of significance implying that a larger share of new immigrants in the local population causes smaller growth rates in average housing prices. This result is consistent with our earlier finding in the sense that the growth in immigration that comes from recent immigrants over the period between 1996 and 2006 has no statistically significant impact on housing markets. On the other hand, as one of the regional characteristics, the average immigration ratio is a significant explanatory variable suggesting that the growth rate of prices increase around 0.04 percent in the long-run if the share of all immigrants in the total population rises by 1 percent. This result is also consistent with what we find by the between-effect models in Table 3 above.

Even though estimations with (5) solve the problems related to unobserved fixed effects, a potential threat to the estimations may come from a possible endogeneity of total immigration ratios. In general, there are three sources of this problem: omitted variables, measurement errors, and simultaneity. Note that fixed-effect and first-difference estimators solve the omitted variables problem by removing all observed or unobserved fixed effects. In addition, since these estimators involve performing OLS on changes of the variables, changes in the immigration ratio come from inflows of recent immigrants— which usually choose their initial destinations where they have some cultural/ethnic connections. Therefore, as recent immigrant correlated with the total amount of immigrants but not with error terms, we can predict that the results of the within and the first-difference models would be protected from the endogeneity problem by their structure. Even though we remove provincial level fixed-effects, the between estimators have biased and inconsistent results. However, we already know that the OLS estimates of coefficients for the immigration ratio are biased upward. This implies that finding a correct instrumental variable reduces its effect on prices, which is already weak as indicated by the between estimators and our specification in (5). As it comes to the simultaneity problem, we cannot come up with a conceivable explanation how housing prices and immigration are determined simultaneously. One explanation would be that immigrants may choose to settle in places where there are more job opportunities and higher income prospects, which also increase housing prices. However, since we already control income and local labour market characteristics in our estimations, this creates not a simultaneity problem but colinearity between the immigration ratio and the variables that may also determine the share of immigration in the local population.

Conclusions and some directions for future research

One out of every five Canadians is foreign-born (an immigrant). During 2001-06 period, about 1.1 million immigrants settled in Canada, thereby increasing the country's immigrant population by about 13.6 percent - about 4 times higher than the percentage growth in Canadian-born population. In some regions, such as Atlantic Canada, immigration prevented population decline. As new comers demand shelter, immigrants represent a large population of potential home buyers.
Movements in housing prices, as analyzed in this paper, are a reflection of changing housing demand. However, our results indicate that while changes in immigrant composition of total population do affect housing demand in Canada, the magnitude of this effect is small. Prices of private dwellings rise by only about a tenth of a percentage if the share of immigrants in total population rises by one percent. As Table 1 data showed, immigrants' demand for owned housing in Canada rises slowly with their length of stay and catches up to non-immigrants only after they have stayed for 20 years in the country. In short, immigrants who arrived in Canada over the past ten years did not contribute to the creation of "housing bubble" in Canada which has recently burst.

On the one hand, these results indicate potentially small contribution of immigrants to the economy through the housing sector. On the other hand, these results could also indicate that immigrants are not (1) contributing towards unsustainable growth in the Canadian economy and (2) causing a displacement of native born residents from an area due to house price effect. This later implication can be tested in a separate paper.

Several positive social externalities are associated with home ownership. These include stable, more law-abiding neighbourhoods, better performance of children of homeowners at school, greater engagement in local democracy and, because homeowners must pay off their mortgages, housing supposedly encourages people to save more than they otherwise would. These externalities warrant an extensive investigation of lower participation of immigrants in housing markets if public policy is to use immigration to strengthen the social and economic fabric of Canadian society.

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Appendix:

We used three censuses, 1996, 2001, 2006 to generate a panel data. In each census, there are three main geographical categories: provinces or territories (PR), census divisions (CD), and census subdivisions (CSD). Since CD bounders tend to be relatively more stable over years, we used CD's to pool all three censuses. In order to make each CD consistent in each census, we first identified CD's affected by boundary changes and then removed them from each census.

We used Catalogue 12-571-XIE for changes between 2001 and 2006 and the information from Statistic Canada for changes between 1996 and 2001. The total removed CD's are 25 and their codes are: 1010,1011, 2419, 2424, 2425, 2426, 2436, 2437, 2441, 2442, 2443, 2444, 2449, 2451, 2455, 2456, 2457, 2458, 2459, 3512, 3514, 4615, 4616, 5909, 5933.

After this adjustment each census has 258 CD's. We compared the land areas for each CD in each census and noticed that in the remaining CD's (where we thought no boundary changes were made) the land areas (square kilometers) have changed across censuses. While between 2001 and 2006 differences in size are minor, the variation becomes higher between 1996 and 2001. Statistic Canada explains it by the following response to our question:

"Users should note that even when the boundaries of standard geographic areas did not change between censuses, the land areas may 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 we are able to control boundary changes in our estimations, there can be some measurement errors in pooled data.

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