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Jiong Tu

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Jiong Tu
Human Resources and Skills Development Canada – Labour Program
and IZA

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IZA

P.O. Box 7240
53072 Bonn
Germany

Phone: +49-228-3894-0
Fax: +49-228-3894-180
E-mail: iza@iza.org

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ABSTRACT

The Impact of Immigration on the Labour Market Outcomes of Native-Born Canadians

Although immigration has become a major growth factor for Canadian labour force, there is little economic research on the effect of immigration on native-born Canadians' labour market performance. This paper examines the relationship between changes in the share of immigrants by sub-labour markets (categorized by skill types and geographic areas) and changes in native wage growth by a two-stage regression analysis, using 1991, 1996 and 2001 Canadian Census data files. After accounting for biases due to native mobility, endogenous location of immigrants and labour demand shifts, the estimated effects of immigration are consistently insignificant or significantly positive. The results are robust to various specifications of sub-labour markets at city, provincial and national levels, suggesting that there is no evidence for a negative impact on native wage growth rate from the large immigrant influx during the 1990s.

JEL Classification: F22, J15, J31, J61

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Corresponding author:

Jiong Tu
Research and Data Development
Human Resources and Skills Development Canada – Labour Program
165 Hotel de Ville
Gatineau, QC K1A0J2
Canada
E-mail: jtu@lakeheadu.ca

1 Introduction

Since the introduction of the points system to Canadian immigration policy, a large number of highly educated working-age immigrants arrived in Canada and substantially changed the size and demographic composition of the labour force. It is therefore natural to investigate how native-born Canadians' labour market outcomes are affected by immigration. However, there is no easy answer to this question: on the one hand, the increased labour supply tends to put downward pressure on wage rates; on the other hand, immigrant consumption also helps to raise the aggregate demand, and hence encourages demand for labour inputs. Since immigration is likely to cause both labour supply and demand curves to shift out, its net effect on the equilibrium wage is theoretically ambiguous in sign and is thus an empirical question.

Although there has been substantial research attempting to address this issue using data from the U.S. (Altonji and Card 1991, Borjas 1991, 2003, Card 1990, LaLonde and Topel 1991, and Ottaviano and Peri 2006), Australia (Addison and Worswick 2002, and Maré and Stillman 2007) and Europe (Bonin 2005, D'Amuri et al. 2009, and Zorlu and Hartog 2005), no consensus has been reached on the net effect of immigration. In Canada, despite its large foreign-born population, there is a paucity of research on this question. As a complement to the existing literature, I will investigate in this paper the effect of immigration on native-born Canadians' wages during the 1990s. Using both a first difference regression and a two-stage regression approach, I link the changes in native wages to changes in the intensity of immigrants in a sub-labour market categorized by skill type and geographic area. The estimates of the effect of immigration are statistically insignificant or significantly positive, depending on the specification of sub-labour markets. My findings indicate that there is no negative effect of immigration on the growth rate of native wages.

The remainder of the paper is organized as follows. Section 2 reviews literature on the effect of immigration on natives. Models of first difference and two-stage regressions are developed in the Section 3. Section 4 discusses the data and tests native geographic mobility. I present my estimation results in Section 5. In Section 6, I discuss the potential bias caused by the endogeneity of the measure of immigrant intensity and use instrumental variable regression to solve this problem. Section 7 concludes.

2 Literature Review

A number of studies addressing the impact of immigration on natives use American data. In his influential paper, Card (1990) compares the Miami labour market to other areas in the U.S. after the 1980 Mariel boatlift. His difference-in- differences analysis shows that the employment opportunities and wages of Miami natives are not adversely affected by the Mariel boatlift. Butcher and Card (1991) extend the study to 24 major American cities and also find little negative wage effect from immigration during the 1970s and 1980s. They explain that the large immigrant inflows raised the cost of living of the immigrant-intensive cities and increase the earnings of high-wage workers. Lalonde and Topel (1991) use the 1970 and 1980 U.S. Censuses and analyze immigrants' quality and assimilation by arrival cohorts. They find little negative effect on native wages, including those of young minority workers.

One criticism of the above spatial approach is the bias toward zero due to native migration. Borjas, Freeman and Katz (1996 and 1997) argue that since the spatial approach fails to consider native migration across cities in response to immigrant inflows, it will underestimate the impact of immigration on the host country.[1] They suggest analyzing the labour market at national level, in the assumption that immigrants affect the destination country as a whole. Their time-series analysis of the U.S. 1980-88 Current Population Survey show that immigration increased the American less-skilled workforce and is partially responsible for the declining wages and job opportunities of native high school dropouts. Borjas (2003) groups natives by measure of skill types, such as education and work experience, instead of geographic area, and finds that immigration has substantially worsened native earnings.

In response to Borjas's criticism on native migration, Card and Dinardo (2000) explore the correlation between changes in immigrant shares and changes in native skill distribution across cities during 1970-90. They find that native population of a skill group in a city actually slightly increases when the number of immigrants rises. Their results are consistent with those of Butcher and Card (1991) in that natives' intercity migration is positively correlated with inflows of recent immigrants. Therefore, they conclude that the estimates of immigration impact with spatial approach are not contaminated by out migration of natives. Card confirms the results in his 2001 and 2005 papers.

There are also studies using data from Australia and European countries. Chang (2004) calibrates a dynamic inter-temporal model on the 1990 Australian Census data, and finds that immigration does not significantly shift the aggregate average wage or unskilled workers' wages. Addison and Worswick (2002) analyze the 1981-1996 Income Distribution Survey of Australia using the spatial approach, and they find modestly positive and insignificant effect of immigration on native wages. Gross (2002) applies a general equilibrium method on the French labour market, and his findings suggest that immigrants have increased the aggregate demand and have created more jobs opportunities than they occupy.

In spite of the large immigrant population in Canada, there are a limited number of studies on this topic. Grant (1998) defines the sub-labour markets by occupation and applies the skill approach on the 1986-1992 Canadian Survey of Consumer Finances. Her findings suggest that immigration has little impact on native wages. Recently, Aydemir and Borjas (2006) extend Borjas's (2003) method to the 1971-2001 Canadian Censuses to show that immigration adversely impacts natives wages and labour supply. However, given the substantial changes in immigration policies and skill composition of immigrants during the three decades (Green and Green 1999), it may be too strong an assumption that the effect of immigration is constant over such a long time span. An alternative is to consider a relatively short period of time to avoid large demographic changes in immigrant population.

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[1] In fact, Card (1990) also noticed that after the 1980 Mariel boatlift, Miami experienced slower population growth, and that fewer natives moved to Miami than to other cities.

3 Methodology

A straightforward method of estimating the impact of immigration on natives is to include in the native wage function the ratio of immigrants to the native-born in a sub-labour market that is combinedly categorized by skill type and geographic area:

$$\ln W_{it} = X_{it}\beta_t + SKILL_{jt}\beta^j_t + AREA_{kt}\beta^k_t + \beta^p_t Y_{pt} + \gamma_t (M/N)_{jkt} + u_{jk} + e_{it} \tag{1}$$

where subscript t stands for the year of observation. W is the weekly wages of a native-born Canadian i. X is a vector of natives' labour market characteristics, such as age, sex, marital status and visible minority. $SKILL$ is a vector of dummies identifying skill type j and $AREA$ indicates geographic area k. Y controls for variations in the labour demand, for example, the unemployment rate in an area p, where p might or might not be the same as k. M is the number of immigrants and N the number of natives with skill j living in area k, so the variable (M/N) measures the intensity of immigrants in a skill-area group. u is a group-specific fixed effect which represents unobserved wage determinants, while e is a random normal error term for individuals.

Assuming the β's and γ are time-invariant, first-differencing will cancel out the fixed effect and yield the following equation:

$$\Delta\ln W_{it} = \Delta X_{it}\beta_t + \Delta SKILL_{jt}\beta^j_t + \Delta AREA_{kt}\beta^k_t + \beta^p_t\Delta Y_{pt} + \gamma_t\,\Delta(M/N)_{jkt} + \varepsilon_{it} \tag{2}$$

where "Δ" represents the change in the value of a variable during a period τ, for example, $\Delta\ln W_{it} = \ln W_{it} - \ln W_{i(t-\tau)}$, and the error term $\varepsilon_{it} = e_{it} - e_{i(t-\tau)}$. The coefficient γ then measures the net effect of a change in the immigrant-to-native ratio on the change in native wages. However, this method requires panel data that trace a person over time, and it cannot be applied on cross-sectional data at the individual level. Given that my dataset is cross-sectional, I aggregate the observations into skill-area groups and use the means of the dependent and independent variables to construct pseudo-panel data. Then the wage function with aggregated data becomes:

$$\overline{\ln W_{jkt}} = \overline{X_{jkt}}\beta_t + SKILL_{jt}\beta^j_t + AREA_{kt}\beta^k_t + \beta^p_t Y_{pt} + \gamma_t (M/N)_{jkt} + u_{jk} + e_{it} \tag{3}$$

where $\overline{\ln W_{jkt}}$ is the mean log weekly wage of skill group j in area k in year t; $\overline{X_{jkt}}$ is the vector of mean values of X variables in the relative group. By definition, Y, M/N and u are already average values. The first difference equation of aggregated data under the assumption of time-invariant β, β^p_t and γ, but time-variant β^j_t and β^k_t, will be as follows:

$$\overline{\Delta\ln W_{jkt}} = \Delta\overline{X_{jkt}}\beta_t + SKILL_{jt}\beta^j_t + AREA_{kt}\beta^k_t + \beta^p_t\,\Delta Y_{pt} + \gamma_t\,\Delta(M/N)_{jkt} + e_{it} \tag{4}$$

If I relax the restriction on time invariance of coefficients β and γ, the first difference equation should then include interactions between time indicators and all the right-hand-side variables of equation (3). Empirically, it requires a large sample size, or number of sub-markets jk, to estimate all the coefficients. However, the limited number of labour force groups defined by skill types and areas will either make the fully interacted equation inestimable, or yield oversized standard errors.

This problem with limited sample size can be solved by a two-stage regression method developed by Borjas et al. (1996), Grant (1998) and Addison and Worswick (2002). In the first stage, means of the logarithm of native weekly wages are calculated for each year separately after controlling for effects of X variables. The difference in the adjusted means between the two surveys is then regressed on the change in the immigrant to native ratio in the second stage. The process can be illustrated as follows:

Stage 1: Run the following regression at the individual level on each cross-sectional dataset:

$$(5) \qquad \ln W_{it} = X_{it}\beta_t + \theta_{jkt}\,(SKILL_{jt}*AREA_{kt}) + v_{it}$$

where the interaction ($SKILL*AREA$) indicates a native worker's skill group j and resident area k, and v_{it} is the residual. Then the average logarithm of weekly wages of skill-area group jk observed in year t is:

$$(6) \qquad \overline{\ln W_{jkt}} = \overline{X}_{jkt}\beta_t + \theta_{jkt} + v_{jkt}$$

where the coefficient estimates θ_{jkt} can be treated as average wages of each skill- area group adjusted for effects from the X variables. Substituting (6) into (3), I obtain the following relationship:

$$(7) \qquad \hat{\theta}_{jkt} = SKILL_{jt}\,\beta^{j}_{t} + AREA_{kt}\,\beta^{k}_{t} + \beta^{p}_{t}Y_{pt} + \gamma_t\,(M/N)_{jkt} + u_{jk} + e_{it}$$

Stage 2: The change in the adjusted average wage, $\Delta\hat{\theta}_{jkt} = \hat{\theta}_{jkt} - \hat{\theta}_{jk(t-\tau)}$, is calculated and used as the dependent variable in a first difference regression, assuming γ and β^{p}_{t} are time invariant while β^{j}_{t} and β^{k}_{t} are time-variant:

$$(8) \qquad \Delta\hat{\theta}_{jkt} = SKILL_{jt}\,\beta^{j}_{t} \quad + AREA_{kt}\,\beta^{k}_{t} \quad + \beta^{p}_{t}\Delta Y_{pt} + \gamma_t\,\Delta(M/N)_{jkt} + \upsilon_{jkt}$$

where $\upsilon_{jkt} = v_{jkt} - v_{jk(t-\tau)}$. The skill and area dummy variables are included to allow for changes in their specific effects over time.

4 Data

An important policy change happened in the late-1980s when arranged employment was no longer a prerequisite for applicants under the skilled worker class immigration. The subsequent rapid flux of Asian immigrants has greatly altered the demographic composition of the Canadian labour market. It is then reasonable to expect that, after the policy change, immigration would have impacted the native born differently than before. In this paper, I focus my study on the period after the policy change, and use the 1991, 1996 and 2001 Canadian Census Public Use Microdata File (PUMF). Another important reason for me to choose these three most recently available censuses is that the definitions of some key variables, such as city and occupation, are consistent in these data files, whereas those in the previous censuses might be different. I restrict my sample to full time (30 hours or more per week) and full-year (worked for 50 weeks or more per year) paid workers, aged 16 to 64.

In Table 1, I compare wages and demographic characteristics of native-born Canadians and immigrants. All wages are deflated by Consumer Price Index (CPI) based on the 1992 Canadian dollar. Native weekly wages have slightly increased over the three censuses, while immigrants experience a fall in their wages. On average, immigrants are three years older than natives and are more likely to be married. There is a large disparity in the visible minority status: merely 2 percent of natives are non-whites, whereas the share of immigrants who are members of visible minority rises from 37 percent in 1991 to 51 percent in 2001. This is not surprising as the major source of recent immigrants has been Asia.

Given the strong tie between education and productivity, educational attainment is used as a measure of skill type. Individuals are categorized into four education groups: less than high school education, high school diploma, postsecondary certificate and university degrees. Both native and immigrant educational distributions have significantly changed over the decade. For natives, the shares of lower educational levels have declined while the proportion of university degree holders increases by 5 percentage points.[2] On the other hand, immigrants initially have larger proportions than natives at both ends of the educational distribution. But the share of immigrants with less than high school education has fallen by 6 percentage points while that with university education have substantially increased by 7 percentage points.

Immigrants' preference for large census metropolitan areas (CMA) is clearly shown in the table: more than 60 percent of immigrants choose to live in Toronto, Vancouver or Montreal, while the total proportion of natives in these three areas is less than 30 percent. About one third of native-born Canadians live in non-CMAs, whereas only 10 percent of immigrants do so. Therefore, urban natives are more likely to be affected by immigrant inflows. Interestingly, the geographic distributions of natives and immigrants have not significantly changed over the three census years. There seems to be no direct evidence for native mobility as a result of immigrant inflow.

The occupational distributions of natives and immigrants are not consistent with immigrants' lead in educational attainment. Over 22 percent of immigrants work in low-paid jobs, such as manual workers, sales and service personnel, whereas this proportion of natives is only 18 percent. Since wage is usually compatible with occupation, occupation is used as an alternative measure of skill type in my empirical analysis.

5 Regression Results

Before I move on to the discussion of my regression results, it is important to tackle the issue of native mobility. If natives, in response to labour market competition

[2] Checking the educational composition by age cohorts, I find that almost half of natives aged between 60 and 64 in 1991 Census have less than high school education, while the average share of the rest of the natives is only 21 percent. This oldest native group is over 65 five years later and leaves the 1996 Census sample. As a result, the share of natives who have less than high school education declines by 4 percentage points, indicating an improvement in native educational attainment.

from immigrants, move to less immigrant-intensive areas, the increase in local labour force from immigration will be offset by such native mobility, and hence the impact of immigration on wages will be underestimated. I use Card and DiNardo's (2000) method to check for native migration, and the results indicate that natives do not move out of immigration-intensive sub-markets. (The mathematical derivation and empirical results are presented in the Appendix.) Therefore, the bias due to native mobility does not seem to be a crucial problem with Canadian data.

In this section, I first categorize skill type by individuals' educational attainment and define geographic area by CMA of residence. Regression results of the first difference equation and the two-stage approach are presented respectively, and the robustness of the results is checked by other specifications of skill-area groups.

5.1 First Difference Regressions with Unadjusted Mean Wages

The first difference regression is specified by equation (4), in which the dependent variable is the change in (unadjusted) mean logarithm of native weekly wages with educational level j in city k, and the independent variables are the change in immigrant-native ratio, change in unemployment rates and changes in average X_{jk} variables. Since there are four education types in 19 CMAs, the number of skill-area groups is 76 with this specification. The three censuses provide three types of intercensal differences: two five-year intervals 1991-1996 and 1996-2001, and one ten-year interval 1991-2001. I plot the changes in average native wages $\Delta \ln W$ against the changes in immigrant to native ratios $\Delta(M/N)$ for these intervals in Figure 1 and 2, and highlight skill groups in the two largest host cities of immigrants, Toronto and Vancouver.

In Figure 1, most of the plots gather around the origin, and regression lines of periods 1991-1996 and 1996-2001 both have flat slopes. Although Toronto and Vancouver are outliers, exclusion of these two most immigrant-intensive cities does not significantly change the slopes. Figure 2 graphs the decadal differences between 1991 and 2001. The regression line is still nearly horizontal, indicating a low correlation between immigrant inflows and changes in native wages.

Table 2 reports the coefficient estimates of $\Delta(M/N)$ from regressions with the three types of differences. The first column is the baseline model in which the skill- and area-specific effects are both assumed to be time invariant and are therefore cancelled out in the first difference equation. For example, the 0.044 coefficient in the third row indicates a 0.44 percent increase in native's wage growth rate, given a 10 percentage-point rise in the change of immigrant-to-native ratio in an education-area group during 1991-2001.[3] However, as illustrated by the diagrams, the estimates are insignificant and close to zero.

In order to separate out area- and skill-specific effects, I consequently include educational attainment, CMA, and both[4] in the regressions. The inclusion of education dummies alone does not affect the coefficients, as estimates in the second column are

[3] I also pool together the two five-year differences and run the same OLS regression including a time dummy variable to identify the intercensal difference in the intercept. The result is again insignificantly positive, but larger in magnitude. (A Chow test of the consistency of coefficients reveals no structural difference between the two periods.)

[4] Joint tests show that the area dummies are insignificant for the two five-year intervals, but significant for the ten-year period. The education variables are all insignificant over the four different specifications.

similar to those in the first column. However, when I only control for CMA-specific effects, the coefficients of *Δ(M/N)* in the third column become smaller in magnitude or even turn negative, but their significance is not increased. One explanation for the lower estimates is that economic growth in these cities optimally affects all education groups, and increases native wages. The inclusion of CMA dummies takes away the positive area effects on wages and lowers the estimates of immigrant effect. The last column reports regression results when both CMA and education effects are controlled. They are mostly closer to zero than the other columns, indicating a negligible effect on native wage growth.

5.2 Two-Stage Regressions with Adjusted Mean Wages

In the previous model, I assumed that returns to native human capital characteristics were unchanged between any two censuses. In order to allow for time-varying coefficients on these control variables, I use the two-stage approach specified by equation (8). While the first stage includes a set of *X* variables, only changes in *(M/N)* and dummies indicating skill and area groups are used in the second stage.

Before reporting the regression results, I again plot the dependent variable $\Delta\hat{\theta}$, changes in the adjusted mean of the logarithm of native weekly wages, against changes in the immigrant-to-native ratio for the two five-year differences in Figure 3. The regression line is slightly positive during 1991-1996, but nearly horizontal during 1996-2001. Unlike Figure 1, these two regression lines have substantially different intercepts, indicating that natives experienced faster wage growth (controlling for observable characteristics) during 1996-2001 than in the preceding five years. Additionally, Figure 4 plots the 1991-2001 decadal differences in which the regression line is moderately negative.

Table 3 reports the coefficient estimates of Δ(*M/N*) from OLS regressions of equation (8). The coefficients are all insignificant and similar to those in Table Results with the ten-year interval are similar to the sum of the first two panels and are around zero. For example, a 10 percentage point increase in the immigrant to native ratio during 1991-2001 reduces native wage growth rate by roughly 2 percentage points after controlling for both area and education. The last row reports the regression results by pooling the two five-year differences, and the estimates are even smaller in magnitude. However, Chow test results show that native wage growth is subject to structural changes between the two five-year periods, which weakens the reliability of the estimates in the last row.

6 Endogeneity of Immigrant Intensity and IV Regressions

Technically, it would be ideal to have exogenous inflows of immigrants into the sub-labour markets. However, such a condition cannot be easily satisfied and, on the contrary, most countries adopt immigration policies that are counter-cyclical.

Altonji and Card (1991) argue that immigrants might be attracted to cities with a booming economy and with relatively high average wages, causing a positive relationship between immigrant density and natives' earnings. Thus, the ordinary least square (OLS) method would result in positively biased estimates of the impact. They suggest using the stock of existing immigrants as an instrumental variable (IV), as new immigrants tend to reside in cities with a large population of earlier cohorts with the same ethnic background. The stock of immigrants has been widely used in a series of empirical works, such as Borjas (1996) and Addison and Worswick (2002).

In Canada, new entrants also tend to live in immigrant-intensive cities. This is proved by the similarity in recently arrived immigrants and old cohorts' geographic distribution. I then use the immigrant-native ratio in the base year as the IV for changes in the immigrant-to-native ratio of each sub-labour market.[5]

6.1 IV Regression Results

IV estimates from first difference regressions are reported in Table 4.[6] The coefficient estimates are bigger in absolute value than OLS results. For example, the 0.730 estimate of baseline model with 1991-2001 decadal difference implies a 7.3 percent increase in native wage growth rate when there is a 10 percentage point increase in the ratio of immigrants to natives. However, most estimates are still insignificant. The positive relationship between changes in natives' wages and changes in immigrant intensity is again primarily driven by area-specific effects, as all the estimates become insignificant when CMA dummies are added. I run Durbin–Wu–Hausman tests for endogeneity of $\Delta(M/N)$ to justify the use of IV. The results show no evidence for endogeneity in the two five-year intervals, but in regressions with the ten-year difference and pooled data (the last two rows) the null hypothesis of no endogeneity is rejected in the baseline model and the model controlling for education-specific effect (column 1 and 2). Since the IV estimates are all positive in these cells, it is safe to conclude that IV regressions do not indicate negative impact of immigration either.

I then use IV in the second stage of the two-stage regression method. The coefficient estimates of $\Delta(M/N)$ are presented in Table 5. The baseline models show a significantly positive effect of immigration in all regressions. For example, during 1991-1996, the native wage growth rate rises by as high as 8 percentage points when the immigrant ratio increases by 10 percentage points. Inclusion of education attainment does not affect the results, as the coefficient estimates in the second row are similar to those in the first row. However, when only CMA indicators are included in the regression, they control for area-specific effects and make the estimates less positive. The effect of immigration is smallest in magnitude after controlling for education and area effects. Although the existing immigrant-to-native ratio is not a perfect instrument, the IV estimates imply that the seemingly positive relation between native wage growth and changes in immigrant ratios is largely due to area effects. When the area effects are controlled, the estimates become insignificant, indicating that immigration has almost no impact on native wage growth.

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[5] Grant (1998) suggests alternative IVs that are derived from immigrants' intended occupation at entry. However, this piece of information is not available in Canadian Censuses.

[6] Although the existing share of immigrants in a city may be a good predicator for changes in immigrant intensity of the corresponding city, such relationship is not strong across skill-area groups. Regressing $\Delta(M/N)$ over the existing (M/N) yields positive coefficient, but the R squared is no more than 0.30. This weak correlation may be the reason for spurious IV estimates under different model specifications.

6.2 *Sensitivity Test*

I further check the robustness of the two-stage regression model by using different categorizations of sub-labour markets, namely occupation-CMA, education-province and education-occupation.

It has been documented that immigrants' foreign education is often poorly recognized in the host country, and that they are in a disadvantageous position in finding jobs matching their level of education (Sweetman 2003). If a large proportion of immigrants working on positions that mismatch their educational attainment, defining skill type by education might not correctly reflect the labour market competition between immigrants and natives. Therefore, I alternatively categorize skill types by occupations. This alteration is made possible by a common occupation variable defined on the 1991 classification basis in the three censuses. As shown in Table 6, both OLS and IV results are insignificantly negative for all years. Durbin–Wu–Hausman tests indicate that there is endogeneity of $\Delta(M/N)$ in baseline models and models controlling for occupation dummies (column 1 and 2) of the 1996-2001 and 1991-2001 periods, where IV estimates are positive and large in absolute value.

Since Borjas (1996) argues that enlarging geographic boundaries may reduce the probability of native cross-area migration and lessen the upward bias on the estimates, I then substitute province for city to define area and run the two-stage regressions by dividing individuals into education-province groups. As reported in Table 7, the OLS results are still insignificant or positive. However, the IV results vary in sign. For example, during 1991-1996, a 10 percentage point rise in the immigrant-native ratio reduces native wage growth rate by about 13 percent when provincial fixed effects are controlled, but the negative effect is greatly reduced when both education and province indicators are included. Regression results based on the ten-year difference also indicate strong area effects, as the inclusion of province dummies turns the estimates negative and on the margin of significance.

Finally, I extend the two-stage regression approach to the national level and substitute education and occupation for *SKILL* and *AREA* terms in equation (8). Now that the bias due to native geographic mobility is eliminated, the estimated effect of immigration is expected to be more negative. However, most OLS results in Table 8 are still close to zero and statistically insignificant, and the IV estimates are even more positive for the two five-year intervals. Only the IV estimates in the last two columns of the 1991-2001 difference are significantly negative. For example, the coefficient –0.324 means that a 10-percentage point increase in the immigrant-native ratio is associated with a 3-percentage point drop in the native wage growth rate. Still, the overall effects of immigration on native wages, even estimated at the national level, are insignificant or moderately negative.

7 Conclusion

There have been debates on the effect of immigration on the labour markets in the host country. Immigrant inflows increase the labour supply, while their consumption of goods and services raises the demand for labour input. Thus, the net impact of immigration on the equilibrium wage is theoretically ambiguous. A number of empirical studies using different approaches and data sources have obtained conflicting estimates of the effect. In Canada, this question is particularly important when policy makers need to evaluate the benefits and costs of immigration and the relevant impact on the local economy. However, little literature analyzes this question using Canadian data despite its large foreign born population.

This paper provides a comprehensive analysis of the impact of immigration on native-born Canadians' wages for the period of 1991-2001 using a first difference regression and a two-stage regression. Cross sectional microdata are aggregated by skill-area groups, and changes in the unadjusted and adjusted mean log weekly wage of natives are regressed on the change in immigrant-to-native ratio of the corresponding group. When sub-labour markets are categorized by education-CMA groups, all the OLS regressions yield small and insignificant coefficient estimates of changes in immigrant intensity. Additionally, I use the immigrant to native ratio in the base year as an instrument for its change, in order to address the bias due to endogenous immigrant residential location. I obtain even more positive estimates. The IV regression results with education-CMA groups indicate that the increasing immigrant inflows are even correlated with a small rise in native wage growth rates.

I apply the two-stage regression approach to a variety of specifications of sub-labour markets, including occupation-CMA, education-province and education-occupation groups, to check robustness of my findings. There is no clear evidence of a negative impact of immigration in either the OLS or IV results with these specifications. Although some estimates from regressions where education-occupation groups are used to divide labour market indicate significantly negative estimates during 1991-2001, the effects are small in magnitude. In summary, both first difference and two-stage regressions on 1991-2001 census data indicate that the substantial immigrant inflows after the policy change in late 1980s did not adversely affect native wage growth rates in the following decade.

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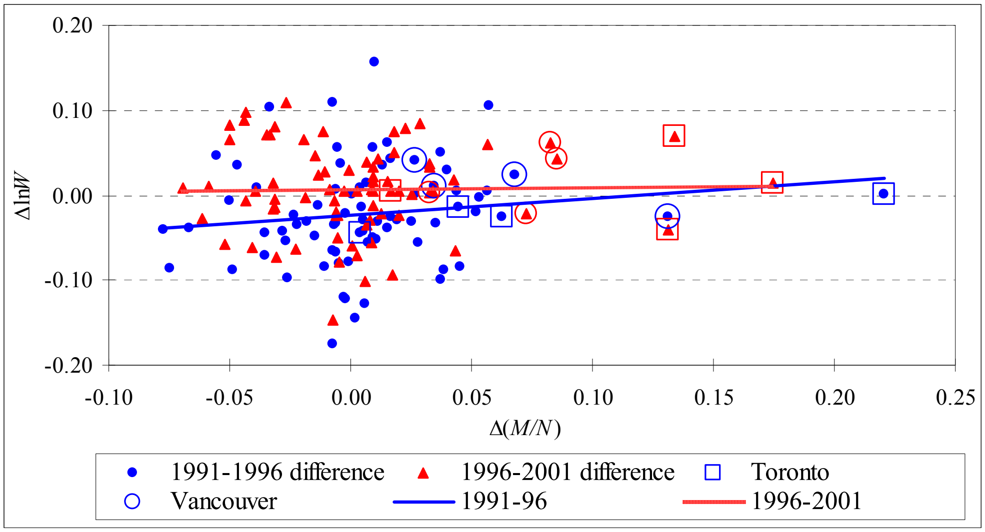
Figure 1

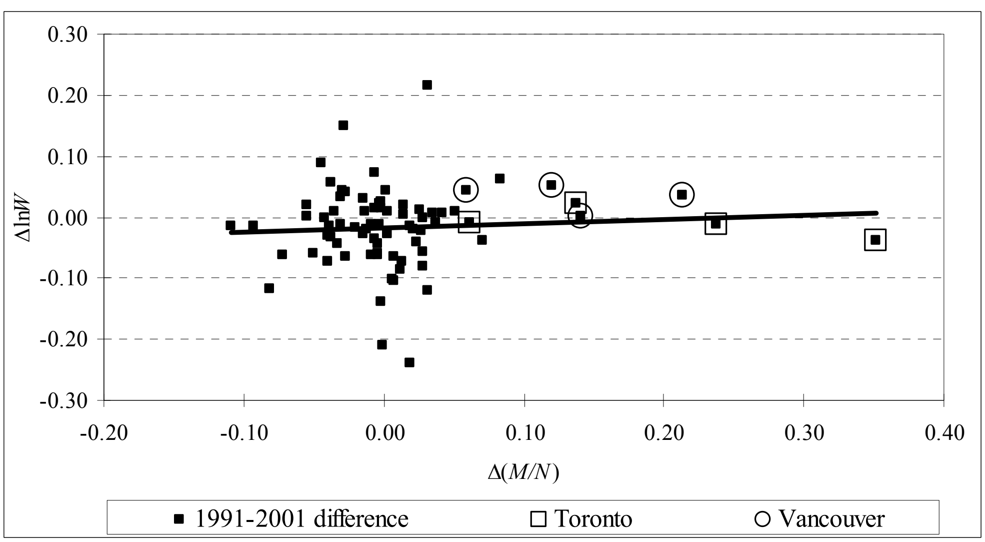
Native ΔlnW (Unadjusted Mean of Log Weekly Wages) against Δ(M/N) by Education-CMA Groups for 1991-1996 and 1996-2001 Intervals



NOTE: Samples include men and women aged 16-64, who have worked at full-time positions for a full year.

Figure 2

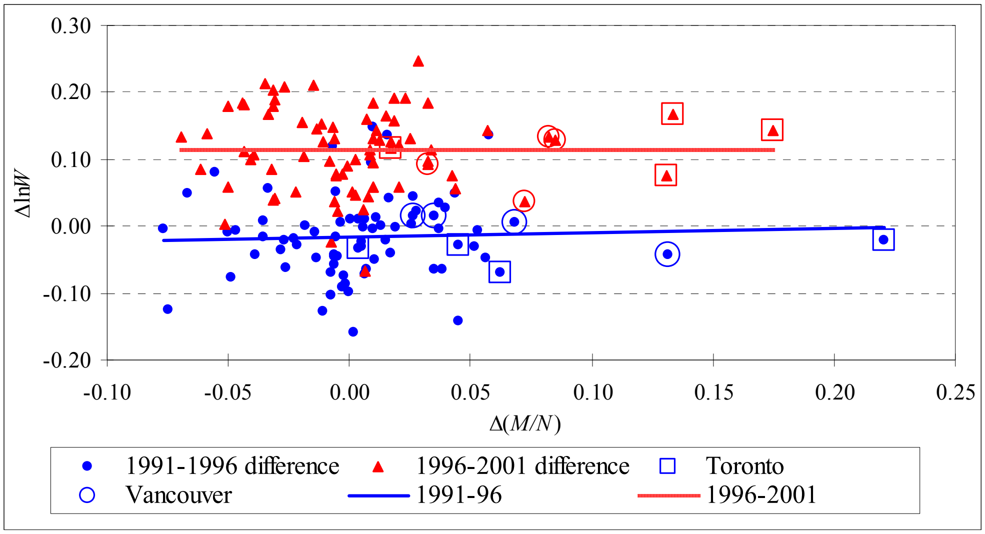
Native ΔlnW (Unadjusted Mean of Log Weekly Wages) against Δ(M/N) by Education-CMA Groups for 1991-2001 Interval



NOTE: Same as Figure 1.

Figure 3

Native $\Delta\hat{\theta}$ (Adjusted Mean of Log Weekly Wages) against $\Delta(M/N)$ by Education-CMA Groups for 1991-1996 and 1996-2001 Intervals

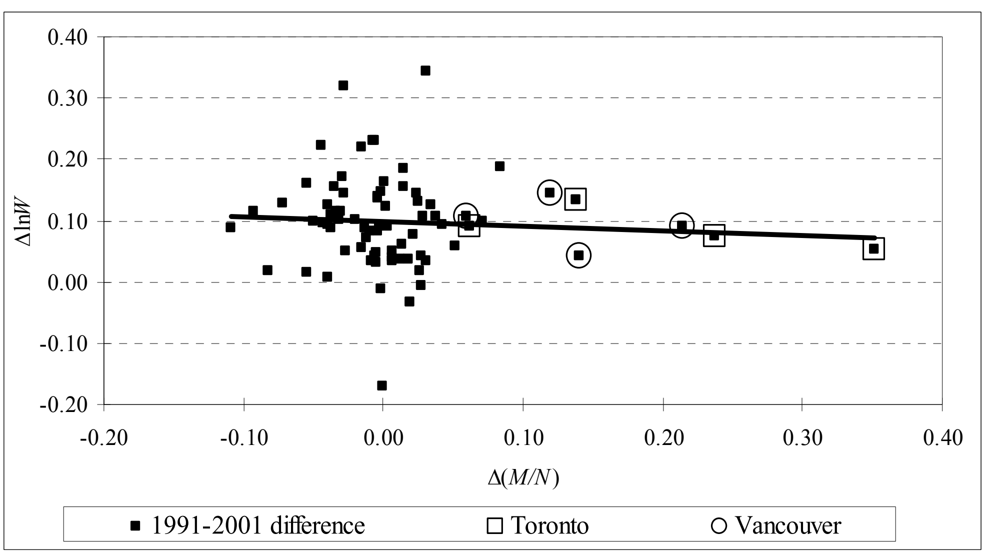


NOTE: Samples include men and women aged 16-64, who have worked at full-time positions for a full year.

The $\hat{\theta}$s are coefficient estimates of skill-area dummies from the first stage regression.

Figure 4

Native $\Delta\hat{\theta}$ (Adjusted Mean of Log Weekly Wages) against $\Delta(M/N)$ by Education-CMA Groups for 1991-2001 Interval



NOTE: Same as Figure 3.

Table 1
Sample Means of Natives and Immigrants: Census 1991, 1996 and 2001

	1991		1996		2001	
	Natives	Immigrants	Natives	Immigrants	Natives	Immigrants
Weekly wages ($)	573.5	573.9	573.0	547.2	579.4	552.4
Age	37.8	42.2	39.2	42.7	39.8	43.2
Male	0.588	0.594	0.578	0.583	0.568	0.563
Visible minority	0.013	0.372	0.012	0.434	0.018	0.513
Married	0.708	0.774	0.713	0.764	0.686	0.752
Educational attainment:						
Less than high school	0.242	0.266	0.196	0.221	0.178	0.197
High school diploma	0.327	0.268	0.308	0.251	0.301	0.250
Certificate	0.266	0.257	0.301	0.280	0.318	0.276
University	0.164	0.209	0.196	0.248	0.203	0.277
Occupation:						
Senior manager	0.012	0.012	0.012	0.012	0.017	0.015
Middle manager	0.113	0.108	0.104	0.092	0.116	0.108
Professional	0.155	0.163	0.172	0.182	0.176	0.192
Semi-professional and technician	0.063	0.058	0.061	0.056	0.077	0.072
Supervisor	0.020	0.018	0.020	0.017	0.018	0.016
Supervisor of crafts and trades	0.028	0.025	0.027	0.025	0.027	0.021
Administrative and senior clerical	0.078	0.059	0.070	0.054	0.065	0.052
Skilled sales and service	0.049	0.053	0.050	0.052	0.043	0.044
Skilled crafts and trades	0.082	0.086	0.078	0.075	0.088	0.083
Clerical personnel	0.126	0.109	0.129	0.114	0.109	0.100
Intermediate sales & service	0.091	0.076	0.092	0.083	0.085	0.075
Semi-skilled manual worker	0.104	0.126	0.105	0.134	0.104	0.131
Other sales and service	0.052	0.067	0.053	0.073	0.048	0.058
Other manual worker	0.027	0.039	0.025	0.031	0.026	0.033
CMA:						
Montreal	0.125	0.102	0.121	0.101	0.122	0.101
Toronto	0.120	0.401	0.115	0.396	0.114	0.424
Vancouver	0.052	0.101	0.053	0.114	0.051	0.119
Other CMAs	0.322	0.263	0.322	0.261	0.328	0.253
Non-CMA	0.381	0.132	0.388	0.127	0.384	0.102

NOTE: Samples include men and women aged 16-64, who have worked at full-time positions for a full year.

Table 2
OLS Estimates of Δ(*M/N*), First Difference Regressions with Education-CMA Sub-Markets

	(1)	(2)	(3)	(4)
Census 1991-1996	0.170	0.206	-0.042	0.036
	(0.178)	*(0.190)*	*(0.230)*	*(0.275)*
Census 1996-2001	-0.024	-0.074	-0.001	-0.061
	(0.203)	*(0.212)*	*(0.280)*	*(0.301)*
Census 1991-2001	0.044	-0.003	-0.065	-0.035
	(0.149)	*(0.148)*	*(0.208)*	*(0.185)*
Census 1991-1996 and	0.119	0.123	-0.030	-0.042
1996-2002 pooled	*(0.118)*	*(0.120)*	*(0.146)*	*(0.153)*
Control for skill or area	Neither	Education only	CMA only	Both

NOTES: Robust standard errors are in parentheses. Significance levels are indicated by * for 10%, ** for 5%, and *** for 1 %. Each observation is a sub-labour market categorized by skill and area. The dependent variable is the unadjusted mean of the logarithm of native weekly wage. Regressions control for changes in the proportion of males, age groups, marital status, visible minority and occupations, and changes in unemployment rate. The sample is restricted to male and female natives, aged 15–64, who have worked at full-time positions for a full year.

Table 3
OLS Estimates of Δ(*M/N*), Two-Stage Regressions with Education-CMA Sub-Markets

	(1)	(2)	(3)	(4)
Census 1991-1996	0.207	0.293	0.007	0.129
	(0.160)	*(0.166)*	*(0.191)*	*(0.204)*
Census 1996-2001	-0.075	-0.119	-0.182	-0.327
	(0.143)	*(0.150)*	*(0.197)*	*(0.218)*
Census 1991-2001	0.098	0.099	-0.158	-0.199
	(0.124)	*(0.125)*	*(0.183)*	*(0.187)*
Census 1991-1996 and	0.071	0.071	-0.046	-0.053
1996-2002 pooled	*(0.106)*	*(0.107)*	*(0.137)*	*(0.139)*
Control for skill or area	Neither	Education only	CMA only	Both

NOTES: Robust standard errors are in parentheses. Significance levels are indicated by * for 10%, ** for 5%, and *** for 1 %. Each observation is a sub-labour market categorized by skill and area. The dependent variable is the adjusted mean of the logarithm of native weekly wage. Regressions in the second stage control for changes in unemployment rate. The sample is restricted to male and female natives, aged 15–64, who have worked at full-time positions for a full year.

Table 4
IV Estimates of Δ(*M/N*), First Difference Regressions with Education-CMA Sub-Markets

	(1)	(2)	(3)	(4)
Census 1991-1996	0.522	0.463	0.622	0.569
	(0.380)	*(0.374)*	*(0.810)*	*(0.926)*
Census 1996-2001	0.162	0.158	-0.537	-0.816
	(0.481)	*(0.447)*	*(1.322)*	*(1.429)*
Census 1991-2001	0.730#	0.780#	0.737	1.256
	(0.372)	*(0.411)*	*(1.031)*	*(1.522)*
Census 1991-1996 and	0.832*#	0.826*#	0.510	0.762
1996-2002 pooled	*(0.298)*	*(0.303)*	*(1.486)*	*(4.173)*
Control for skill or area	Neither	Education only	CMA only	Both

NOTES: Same as Table 2.

Durbin-Wu-Hausman statistic is significant at .05 level, indicating the variable Δ(*M/N*) may be endogenous and justifying the use of IV estimates

Table 5
IV Estimates of Δ(*M/N*), Two-Stage Regressions with Education-CMA Sub-Markets

	(1)	(2)	(3)	(4)
Census 1991-1996	0.804*#	0.798*	0.304	0.100
	(0.343)	*(0.335)*	*(0.679)*	*(0.700)*
Census 1996-2001	0.541*#	0.543#	0.190	0.264
	(0.339)	*(0.348)*	*(1.479)*	*(0.943)*
Census 1991-2001	0.707*#	0.700*#	0.555	0.061
	(0.232)	*(0.233)*	*(0.993)*	*(1.759)*
Census 1991-1996 and	0.660*	0.667*	0.144	1.830
1996-2002 pooled	*(0.236)*	*(0.241)*	*(2.275)*	*(11.967)*
Control for skill or area	Neither	Education only	CMA only	Both

NOTES: Same as Table 2.

Durbin-Wu-Hausman statistic is significant at .05 level.

Table 6
OLS and IV Estimates of Δ(*M/N*), Two-Stage Regressions with Occupation-CMA Sub-Markets

	(1)	(2)	(3)	(4)
OLS Regressions				
Census 1991-1996	-0.016	-0.016	-0.066	-0.074
	(0.072)	*(0.077)*	*(0.073)*	*(0.078)*
Census 1996-2001	-0.132	-0.166	-0.134	-0.175
	(0.085)	*(0.085)*	*(0.091)*	*(0.091)*
Census 1991-2001	0.025	0.014	-0.021	-0.046
	(0.069)	*(0.072)*	*(0.078)*	*(0.082)*
IV Regressions				
Census 1991-1996	-1.428	-2.619	0.464	0.211
	(4.337)	*(5.879)*	*(0.314)*	*(0.245)*
Census 1996-2001	3.449#	1.996#	-0.354	-0.576
	(4.214)	*(1.488)*	*(0.308)*	*(0.365)*
Census 1991-2001	1.872#	1.726*	0.224	-0.051
	(1.217)	*(0.855)*	*(0.292)*	*(0.310)*
Control for skill or area	Neither	Occupation only	CMA only	Both

NOTES: Robust standard errors are in parentheses. Significance levels are indicated by * for 10%, ** for 5%, and *** for 1 %. # Durbin-Wu-Hausman statistic is significant at .05 level. The dependent variable is the adjusted mean of the logarithm of native weekly wage. Regressions in the second stage control for changes in unemployment rate. The sample is restricted to male and female natives, aged 15–64, who have worked at full-time positions for a full year.

Table 7
OLS and IV Estimates of Δ(*M/N*), Two-Stage Regressions with Education-Province Sub-Markets

	(1)	(2)	(3)	(4)
OLS Regressions				
Census 1991-1996	-0.068	0.196	-0.417	-0.159
	(0.301)	*(0.317)*	*(0.235)*	*(0.243)*
Census 1996-2001	0.220	0.186	-0.031	-0.162
	(0.294)	*(0.333)*	*(0.249)*	*(0.284)*
Census 1991-2001	0.489	0.633*	-0.492	-0.368
	(0.269)	*(0.252)*	*(0.267)*	*(0.215)*
IV Regressions				
Census 1991-1996	24.778#	6.146#	-1.332*#	-0.710
	(84.611)	*(5.801)*	*(0.558)*	*(0.648)*
Census 1996-2001	0.954	0.773	-17.497	-0.361
	(1.063)	*(1.259)*	*(132.229)*	*(0.768)*
Census 1991-2001	3.683*#	2.662*	-3.712	-1.319
	(1.773)	*(1.091)*	*(1.937)*	*(0.645)*
Control for skill or area	Neither	Education only	Province only	Both

NOTES: Same as Table 6.

Table 8
OLS and IV Estimates of Δ(*M/N*), Two-Stage Regressions with Education-Occupation Sub-Markets

	(1)	(2)	(3)	(4)
OLS Regressions				
Census 1991-1996	0.016	0.009	0.165	0.159
	(0.247)	*(0.254)*	*(0.258)*	*(0.267)*
Census 1996-2001	0.094	0.015	0.133	0.036
	(0.072)	*(0.079)*	*(0.075)*	*(0.085)*
Census 1991-2001	0.005	-0.150	0.025	-0.176*
	(0.083)	*(0.085)*	*(0.081)*	*(0.077)*
IV Regressions				
Census 1991-1996	30.603#	15.624#	10.444#	7.969#
	(185.522)	*(30.270)*	*(32.829)*	*(8.784)*
Census 1996-2001	0.221*	0.087	0.267*	0.117
	(0.107)	*(0.135)*	*(0.104)*	*(0.134)*
Census 1991-2001	-0.024	-0.413*#	0.125	-0.324*
	(0.112)	*(0.142)*	*(0.108)*	*(0.125)*
Control for skill	Neither	Education only	Occupation only	Both

NOTES: Same as Table 6.

Appendix
Test for Native Mobility

Although the descriptive statistics in Section 4 show that there is almost no change in native geographic distribution over the decade, shifts in natives' skill distribution across cities is not clear. Theoretically, when native geographic migration is not affected by immigration, immigrant inflow into a particular skill group in a city will increase the labour supply of this group; otherwise, if native out-migration offsets the immigrant inflow, there will be little change in labour supply. In order to test for the presence of native migration across skill-area groups, I employ the method developed by Card and DiNardo (2000).

First, I define P to be total population of a sub-labour market, and it is the sum of immigrants and natives, or $P = M + N$. The following equation then holds:

A(1) $P_{jk} / P_k = (M_{jk} + N_{jk}) / (M_k + N_k)$

Take the logarithm of both sides:

A(2) $\ln (P_{jk} / P_k) = \ln (M_{jk} + N_{jk}) - \ln (M_k + N_k)$

The percentage change in the share of total population in a skill-area group is then approximately:

A(3) $\Delta\ln (P_{jk} / P_k) = (\Delta M_{jk} + \Delta N_{jk}) / (M_{jk} + N_{jk}) - (\Delta M_k + \Delta N_k) / (M_k + N_k)$

$\Delta\ln (P_{jk} / P_k) = (\Delta M_{jk} + \Delta N_{jk}) / P_{jk} - (\Delta M_k + \Delta N_k) / P_k$

$\Delta\ln (P_{jk} / P_k) = (\Delta M_{jk} / P_{jk} + \Delta N_{jk} / P_{jk}) - (\Delta M_k / P_k + \Delta N_k / P_k)$

Re-write the above equation into the sum of relative growth rate of immigrants and of natives:

A(4) $\Delta\ln (P_{jk} / P_k) = (\Delta M_{jk} / P_{jk} - \Delta M_k / P_k) - (\Delta N_{jk} / P_{jk} - \Delta N_k / P_k)$

Next, I assume that natives' reaction linearly depends on the immigrant inflow:

A(5) $(\Delta N_{jk} / P_{jk} - \Delta N_k / P_k) = a + b\ (\Delta M_{jk} / P_{jk} - \Delta M_k / P_k) + \xi_{jk}$

Substitute it into equation A(4):

A(6) $\Delta\ln (P_{jk} / P_k) = a + (1 + b)(\Delta M_{jk} / P_{jk} - \Delta M_k / P_k) + \xi_{jk}$

Thus, the coefficient $(1 + b)$ shows the relation between immigrant inflow and relative labour supply of skill group j across areas. When this coefficient is close to zero, that is b close to –1, it means that native mobility offsets the immigration-induced impact on labour supply. However, when the coefficient estimate is around 1, or b close to 0, native mobility across areas is then not correlated with immigrant inflow and immigration will increase the relative supply of labour.

In accordance with my skill-area approaches, I run the regression of equation A(6) on the four specifications of sub-markets: education-CMA, occupation-CMA, education-province and education-occupation. The estimates of $(1 + b)$ are reported in the following table, where each cell stands for a separate regression of an intercensal interval. All the estimates of $(1 + b)$ are significantly greater than zero, which implies that natives have not moved away from immigrant intensive skill-area groups to offset the impact of immigration on labour supply. I also test the hypothesis that (1 + *b*) = 1, and most estimates are significantly greater than 1. The results indicate that natives do not move out of a skill-area group in response to immigrant inflow. In fact, native migration is even positively correlated with an increase in immigrant intensity during 1991-1996.

Estimates of (1 + *b*) in Equation A(6) with Various Specifications of Skill-Area Groups

	(1)	(2)	(3)	(4)
Census Intervals	1991-1996	1996-2001	1991-2001	Pooled 1991-1996 and 1996-2001
Education-CMA	3.506#	1.532	2.763#	2.778#
	(0.343)	*(0.301)*	*(0.283)*	*(0.175)*
Occupation-CMA	1.517#	2.066#	2.216#	1.953#
	(0.142)	*(0.161)*	*(0.181)*	*(0.094)*
Education-Province	2.759#	1.119	2.674#	2.445#
	(0.880)	*(0.416)*	*(0.591)*	*(0.371)*
Education-Occupation	2.588#	0.741	1.445#	1.504#
	(0.214)	*(0.106)*	*(0.159)*	*(0.102)*

NOTES: Standard errors in parentheses, all coefficients are significantly different from zero.

Estimates are significantly greater than 1 (or $b>0$) at 5% significance level; otherwise, not different from 1 (or $b = 0$).