# A Note on Canadian Migration to the United States During the 1980s

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**Abstract**

Considerable media attention has been recently directed towards the flow of highly talented Canadians to the United States. There are firm theoretical reasons, however, to believe that qualitative differences in migration began in the 1980s, owing to the widening distribution of earnings and increased returns to education in the United States relative to Canada. Either of these could result in qualitative improvements in the migration flow. Furthermore, U.S. immigration policy remained essentially unchanged during the 1980s, thus providing an ideal time period to study this phenomenon. We use a flexible empirical approach to document these changes in immigrant quality using 1980 and 1990 U.S. census data. Our results are mixed regarding the hypothesis of relative improvements in immigrant quality.

Acknowledgements to be added here.

JEL Classification: J61

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## I. Introduction and Background

The 1980s in the United States were subject to two related phenomenon that theoretically could have changed the qualitative composition of Canadian immigration over that decade, before the current debate on the "brain drain" had even started. First, the distribution of earnings widened in the United States by a greater degree than it did in Canada (Blackburn and Bloom 1993; Gottschalk and Smeeding 1997; Richardson 1997). Assuming a high correlation between earnings in the United States and Canada, Canadians of high ability (and thus ability to earn higher incomes in the United States) would be more likely to migrate. Second, and related to the first point, returns to education increased in the United States relative to Canada over this period (Burbridge, Magee, and Robb, 2002). This provided another impetus for the migration of highly educated Canadians.

Evidence on these qualitative differences in migration from Canada to the United States is sparse. Borjas (1988) presented empirical evidence that supported his theory that the wider distribution of income in the United States resulted in highly skilled Canadian males self-selecting into the United States labour market. Mueller (1999) followed essentially the same methodology using updated data from the 1990 U.S. Census and discovered that highly skilled Canadians of either gender continued to migrate to the United States in the 1980s.[1]

In what follows, we have little to say about the quantitative aspects of Canadian migration to the United States. Rather, we explicitly address the shortcomings of these previous qualitative studies. Although both of these qualitative studies used two decennial censuses, which allowed for the identification of both assimilation and cohort effects,[2] the general problem was that the cohort effects were constrained to be equal across censuses, while the assimilation effect was constrained to be the same for all immigrants,

[1] Recent studies by Schoeni (1997) and Cohen, Zach and Chiswick (1997), for example, have also addressed qualitative changes in immigration to the United States over the 1980s. Unfortunately, the Canadian-born in each of their samples is included with other immigrant source countries such as Australia and the United Kingdom.

[2] Qualitative aspects of immigrant cohorts are generally the result of two effects. First, the assimilation effect is a measure the return to earnings for time spent in the host country. Second, there is cohort effect which means that successive groups of immigrants may be qualitatively different as a result of different levels education, etc. The two of these work together to determine the quality of immigrants in an economy.

irrespective of census year or entry cohort. As outlined by Borjas (1985), this can bias both cohort and assimilation effects. This study, along with recent work by LaLonde and Topel (1992) and Baker and Benjamin (1994), has shown that these effects do differ by entry cohort. The model estimated below does not impose these constraints on the estimation. As with these three related studies, we use a general model in which coefficient estimates can vary both between census years and between immigrants and the native-born, thus avoiding these potential estimation biases.[3] In addition, the analysis that follows will explicitly include females. In many studies of immigrant quality, it is only the earnings of male immigrants that are analyzed.

The 1980s provide a particular good decade to study this phenomenon since U.S. immigration policy remained relatively unchanged. Major changes in U.S. policy occurred with the Immigration Act of 1965. The next major policy changes were in 1989 and 1990. The former year is when the Canada-U.S. Free Trade Agreement came into effect. It included provisions for the temporary migration of skilled workers between the two countries.[4] The latter year is when the U.S. Congress enacted the Immigration Act of 1990, which came into effect in 1992. This new act increased the overall number of permanent immigration slots, nearly tripled the number of employment-based admissions, increased the number of visas for temporary workers, and made the transition from temporary to permanent immigrant status much easier (Lowell, 2001). Since it is these policy changes that have been responsible for much of the debate since the late-1990s, especially regarding the interpretation of much of the data,[5] the 1980s provide a good period of analysis for our purposes.

The next section of the paper will discuss the methodology to be used in the empirical part of the paper. Section III discusses the 1980 and 1990 United States Census data that were used in the analysis. Section IV presents the results from the estimation of model. The final section concludes.

[3] A series of Chow tests (not reported here) support rejection of these equality constraints. Testing for differences between census years, and between American-born and Canadian-born within each of these samples, for both males and females, we decidedly rejected the equality constraints.
[4] These provisions have continued since 1994 under the North American Free Trade Agreement.
[5] See Finnie (2001) and related comments for a good review of the current debate.

## II. Methodology

Following Borjas (1985), LaLonde and Topel (1992), and Baker and Benjamin (1994),[6] the standard earnings function in year $t$ for immigrants who arrived in the host country in period $i$ is:

$$y_t = X_t'\beta_t + \sum_i C_{i,t} + \varepsilon_t, \tag{1}$$

where $y_t$ is a measure of earnings at time $t$, $X_t$ is a vector of observable individual characteristics, $\beta_t$ is the corresponding parameter vector, $C_{i,t}$ are intercepts for $i$ entry cohorts at time $t$. The error term for cohort $i$ at time $t$, $\varepsilon_{i,t}$, can be written as the sum of three unobservable components:

$$\varepsilon_{i,t} = a_{i,t} + b_{i,t} + u_i, \tag{2}$$

where: (1) $a_{i,t,}$ is a cohort-specific assimilation effect and reflects the rate at which the cohort accumulates country-specific human capital; (2) $b_{i,t}$ represents the cohort-specific unobserved time effect and is (usually) considered to be the result of macroeconomic conditions that can differ at the time of entry for each cohort; and, (3) $u_i$ is a cohort-specific fixed effect representing other unobserved factors that affect productivity and hence earnings. This is the result of unobserved talent and is referred to as the "quality" of the immigrant cohort.

The use of a single cross-section allows the estimation of the assimilation effect over $k$ years as $(a_{i,t} - a_{i+k,t})$ or as $(a_{i,t} - a_{i,t-k})$ using two cross sections that are $k$ years apart. Under the assumptions $E[b_{i,t} - b_{i+k,t}] = 0$, and $E[u_i - u_{i+k}] = 0$, the estimates to these returns can be estimated using the $C_{i,t,}$. The first assumption states that the unobserved time effects have the same effect on each cohort (for example, macroeconomic conditions effect all entry cohorts the same). The second assumption states that cohort quality does not vary between immigrant groups. If there are qualitative differences between cohorts, then this could lead to either positive or negative bias in the cross-section estimates of assimilation.

Using two separate cross-sections that are $k$ years apart, the assimilation effect free of this fixed-effect bias can be estimated as $(a_{i,t} - a_{i,t-k})$ by using $(C_{i,t} - C_{i,t-k})$ under the assumptions that the $u_i$ are fixed over time, and that $E[b_{i,t} - b_{i,t-k}] = 0$. This latter effect says that unobserved time effects do not differ between cohorts. This assumption is likely to be unrealistic. Economic conditions are likely to have differential effects on different immigrant cohorts, thus biasing the assimilation estimate. For example,

[6] This section essentially follows the exposition of the model as outlined by Baker and Benjamin (1994).

immigrants entering the United States in the 1970 faced an annual unemployment rate of 4.9 per cent, compared to a rate of 7.1 per cent for those who entered in 1980. If we assume that this macroeconomic environment affects some base group in the same way, the remedy to this problem is relatively straightforward. The solution is to compare immigrant earnings to the growth in those of some base group *n*, which will be native-born Americans in our estimates.[7] This is accomplished by running an additional regression:

$$y_t = X'_{n,t}\beta_{n,t} + C_{n,t} + \varepsilon_{n,t}, \tag{3}$$

where $\varepsilon_{n,t} = b_{n,t} + u_n$. Equation (3) is the analog to equation (2) for the native-born control group *n*, but with no assimilation effect. Thus, using two cross sections the difference $(C_{i,t} - C_{i,t-k})$ - $(C_{n,t} - C_{n,t-k})$ will be equal to $(a_{i,t} - a_{i,t-k})$, as long as $E[b_{i,t} - b_{i,t-k}] - E[b_{n,t} - b_{n,t-k}] = 0$. In other words, as long as the time effects for both natives and immigrants are equal, these effects will not bias our estimate of immigrant assimilation using the two cross sections.

The relationship between the quasi-panel and cross-section approaches can be expressed as follows. The predicted average level of earnings of cohort *i* in period *t* can be expressed as:

$$\hat{y}_{i,t} = \overline{X}'_{i,t}\hat{\beta}_t + \hat{C}_{i,t}, \tag{4}$$

and the predicted average level of earnings for this same cohort in year $t - k$ using the average level of observables of cohort *i* at time *t* is:

$$\hat{y}_{i,t-k} = \overline{X}'_{i,t}\hat{\beta}_{t-k} + \hat{C}_{i,t-k}. \tag{5}$$

Similarly, the predicted earnings for a cohort that has the same number of years in the United States as those of cohort *i* in year $t - k$ (using the average level of observables of cohort *i*) is:

$$\hat{y}_{i+k,t} = \overline{X}'_{i,t}\hat{\beta}_t + \hat{C}_{i+k,t}. \tag{6}$$

[7] LaLonde and Topel (1992) use a number of base groups in their analysis including the native born, earlier immigrant cohorts and, for comparisons with Mexican and Latin American immigrants, American-born Hispanics. They find that their inferences are not sensitive to the choice of base group. Using Canadian census data, Baker and Benjamin (1994) also try a number of different "base groups" arguing that there does not exist a strong argument for a "natural" base group. They, however, investigate immigrants from a number of source countries making the choice of an appropriate base group more complex. Similarly, Grant (1999) finds that her results are not sensitive to the choice of natives as the base group. Given the intentions of this paper, it seems reasonable to assume that the American-born are the natural base group in our analysis given the similarities between the two countries.

Thus, in year *t*, cohort *i* + *k* has the same number of years since migration as cohort *i* does is year *t* + *k*.

Finally, the the predicted earnings for the native-born base group *n* in year *t*, using the average level of observables for immigrant cohort *i*, is:

$$\hat{y}_{n,t} = \overline{X}'_{i,t}\hat{\beta}_{n,t} + \hat{C}_{n,t}. \tag{7}$$

By using equations (4) and (6), the cross-section estimate of assimilation ($a_{i,t} - a_{i+k,t}$) is equal to $(\hat{y}_{i,t} - \hat{y}_{i+k,t})$ and this may be expressed as the sum of two components:

$$\hat{y}_{i,t} - \hat{y}_{i+k,t} = \left[(\hat{y}_{i,t} - \hat{y}_{i,t-k}) - (\hat{y}_{n,t} - \hat{y}_{n,t-k})\right] + \left[(\hat{y}_{i,t-k} - \hat{y}_{i+k,t}) - (\hat{y}_{n,t-k} - \hat{y}_{n,t})\right]. \tag{8}$$

The first term on the right-hand side of this equation measures the "within cohort" growth of the earnings of cohort *i* across the two data sets. It follows the same cohort across the two census periods and is the quasi-panel measure of assimilation. If we substitute from the equations above, we get:

$$\begin{aligned}(\hat{y}_{i,t} - \hat{y}_{i,t-k}) - (\hat{y}_{n,t} - \hat{y}_{n,t-k}) = \overline{X}_{i,t}\left[(\hat{\beta}_t - \hat{\beta}_{t-k}) - (\hat{C}_t - \hat{C}_{t-k})\right] \\ + (a_{i,t} - a_{i,t-k}) + (b_{i,t} - b_{i,t-k}) - (b_{n,t} - b_{n,t-k}).\end{aligned} \tag{9}$$

If we have common net time effects on immigrant and native-born group (as we have assumed), this measure of assimilation contains: (1) a component which captures the net change in the "prices" of observables across immigrants and the base group between periods $t - k$ and $t$; and, (2) a component capturing the change in intercept due to assimilation ($a_{i,t} - a_{i,t-k}$). Thus, this within cohort growth is the measure of assimilation of cohort *i* over the ten-year period between censuses.

The second term in equation (8) measures the change in earnings for a cohort with a fixed number of years in the United States across the two data sets. If the quality of cohorts has been declining over time, then this term will be positive. Substituting from the above, this equals:

$$\begin{aligned}(\hat{y}_{i,t-k} - \hat{y}_{i+k,t}) - (\hat{y}_{n,t-k} - \hat{y}_{n,t}) = \overline{X}_{i,t}\left[(\hat{\beta}_{t-k} - \hat{\beta}_t) - (\hat{C}_{n,t-k} - \hat{C}_t)\right] + (a_{i,t-k} - a_{i+k,t}) \\ + (b_{i,t-k} - b_{i+k,t}) - (b_{n,t-k} - b_{n,t}) + (\mu_i - \mu_{i+k}).\end{aligned} \tag{10}$$

Given the assumptions made about the unobserved time effects, equation (10) has three parts: (1) the net effect of changes in the prices of observables; (2) ($a_{i,t-k} - a_{i+k,t}$), the difference across time in the labour market outcomes of two cohorts at similar stages of assimilation; and, (3) ($\mu_i - \mu_{i+k}$), the difference in the fixed effect across cohorts. Again, if there has been a decline in immigrant quality over time, this term will be positive. By contrast, if there has been an increase in immigrant quality, the term will be negative.

In sum, this methodology allows us to differentiate within cohort effects (which reflect the earnings assimilation of a single entry cohort) from across cohort effects (which reflect qualitative changes in different entry cohorts). This is done with a flexible estimation procedure which does not constrain coefficient estimates for immigrants and natives to be equal, nor are these constrained to be equal within each of these groups across census years. For example, given the increased returns to education in the United States during the 1980s, constraining coefficients to be equal across census years could introduce biases into coefficient estimates and lead us to erroneous inferences.[8]

## III. Data

The data are obtained by merging the 5 per cent individual records of the 1980 and 1990 U.S. Censuses. A 1/100 subsample of the American-born was randomly generated for both years while all Canadian-born individuals were retained. Our sample was further limited to include only non-institutionalized individuals between the ages of 25 and 64 who worked at least 40 weeks in the year prior to the census, were not self-employed, did not attend school, and had at least $1000 in real earnings (in 1989 dollars) in the reference calendar year. The income variable is the natural logarithm of real annual earnings. This includes wage and salary income, as well as self-employment and farm income, in the year preceding the respective census year (i.e., 1979 and 1989). Although we eliminated the self-employed from the sample, those who are primarily engaged in paid employment may still have some income from these other two sources.[9]

The years of education variable was coded to equal the number of years corresponding to the highest level of education completed (in the 1990 data) or number of years of schooling completed (in the 1980 data). Since the education variables were somewhat different (especially for post-secondary education) in the 1980 and 1990 data, the variable was constructed to be as comparable as possible across censuses. For example, in the 1990 data, some post-secondary education or an associate degree was coded as 14 years of education, while a bachelor's degree was coded as 16, a master's or professional degree as

[8] As mentioned above, Chow tests allowed us to reject the restrictions on coefficients. See footnote 3.
[9] In an earlier version of this paper (which did not include occupation and industry controls), the analysis was also conducted using the log of the real annual wage and the log of real weekly earnings. Also, although we included only those with 40 plus weeks worked, similar to some of the recent immigration literature (Baker and Benjamin 1994; Grant 1999), we also did estimates relaxing this restriction. In any of these alternative estimations, there was no substantial change in the results.

18 years, and a doctorate as 20. The 1980 data, by contrast, contains an education variable for number of years of schooling completed. Thus, four years of college completed was assumed to be equivalent to a bachelor's degree, while 6 years of completed college education were coded as equivalent to a master's or professional degree, and 8 years or more was assumed to be equivalent to a doctorate. Experience was calculated using the familiar Mincer proxy (i.e., age – years of education - 6).[10]

A marriage variable was coded to one if the respondent said that he or she was married, and zero otherwise. Similarly, a dummy variable was coded to one if the individual worked at least 40 weeks in the year prior to the census. A number of variables were also coded to equal one related to the number of hours per week the individual normally worked. For Canadian immigrants, entry cohorts are five-year periods beginning with 1960-64, a ten-year cohort for those who entered between 1950 and 1959, and a single cohort for those entering before 1950.

The final sample contains 12,271 Canadian-born and 34,076 American-born males. The numbers for females are 10,784 Canadian-born and 24,694 American-born individuals. Appendix Table A-1 provides a detailed disaggregation of sample sizes by place of birth, gender and entry cohort.

Table 1 contains the summary statistics for American-born and Canadian-born males and females in the 1980 and 1990 censuses. In either census year and for both genders, Canadians have higher real annual earnings compared to their American counterparts and this earnings gap has increased between census years. In 1980, Canadian males had earnings that were about 15 per cent higher while 10 years later the gap increased to about 23 per cent. For females, the earnings advantage increased from about 3.5 per cent to 13 per cent over the same period.

Of course these are unadjusted earnings differentials and the introduction of control variables may eliminate any earnings difference between the two groups. Indeed, there are important differences in these characteristics. The average Canadian-born male had only slightly more education than his average American-born counterpart in 1980. Canadian-born females had slightly less than the American-born in that year. By 1990, Canadian males had about one-half year more education, while the average Canadian female had about one-quarter year more. Given rates of return to education in 1990 of over 10 per cent per

[10] In the original data, the age variable is continuous, but since the years of education variable was not coded as a continuous variable, experience was marginally negative in a few cases and was therefore bottom-coded to zero.

year for the U.S.-born (see Appendix Table A-2), these educational attainment differences are expected, but alone could not account for the entire earnings differential between Canadians and Americans.

We now turn our attention to the multivariate estimates.

## IV. Multivariate Estimation Results

Table 2 presents the estimates of equations (1), (9), and (10) above.[11] The upper panel is for males, and the lower panel is for females. The first column contains the 1990 cross-section estimates of immigrant assimilation for various immigrant cohorts. These estimates are constructed from the full regression results (Appendix Table A-2) by subtracting the intercept coefficient for one entry cohort from the intercept coefficient for the cohort with ten fewer years in the United States. This gives the cross-section estimate of ten years of assimilation. The estimate for the 1975-79 cohort, for example, is obtained by subtracting the intercept for the 1985-90 cohort (the most recent immigrant cohort in the 1990 census) from the intercept for the 1975-80 cohort (i.e., 8.6761 – 8.6699 = 0.0062). Positive coefficients suggest that the immigrant cohort has positive economic assimilation over the 10-year period while negative coefficients are suggestive of immigrants experiencing declining relative incomes over the same time period.

In these cross-sectional estimates for males, only the 1960-64 entry cohort displays negative economic assimilation as evidenced by the coefficient which differs significantly from zero at the five per cent level. For females, the cross-sectional estimates suggest that three of the four entry cohorts did not experience significant (at the five per cent level) positive or negative assimilation over the decade. Only the 1970-74 cohort had significantly negative assimilation. Thus, the cross-sectional results are mixed for different entry cohorts with either significantly negative coefficients or coefficients statistically indistinguishable from zero. But, these are precisely the types of biases we expect to see in the cross-section if there has been an improvement in immigrant quality over time. In other words, these negative assimilation effects in the cross-section could be due to a general improvement in the earnings capacities of new Canadian immigrants in the U.S. economy relative to earlier immigrant cohorts.

As discussed above, these cross-sectional cohort estimates of assimilation can be broken down into components within and across cohorts, both with and without native comparisons (the first and second

terms in equation (8) above). The cross-section estimates can be suggestive of economic assimilation, but can be biased upwards as a result of declining immigrant quality over time. The within cohort estimate will reflect the true assimilation of the immigrant cohort, whereas the across cohort component will reflect any qualitative changes in the average individual within a given cohort. Columns 2 and 3 of Table 2 present these estimates without any comparisons to the earnings experiences of the American-born over the same period. The fourth and fifth columns include these comparisons. Again, the rationale for this latter exercise is that it allows for the control of changing macroeconomic conditions that (by assumption) equally affect both Canadians and Americans in the U.S. economy.

The within cohort estimates without native comparisons are generally small and insignificantly different from zero for males. The only exception is the 1970-74 entry cohort which has a significantly positive assimilation effect. Similarly, for females, only the 1960-64 and 1970-74 cohorts have significantly positive assimilation effects. With native comparisons, these results hold, although the magnitude of the male assimilation effect is much larger; about 11.8 per cent in the estimate with native comparisons compared to only about 9.3 per cent without.

Since we are mainly concerned with the "quality" of newer Canadians in the United States, the across cohort estimates is what we will concentrate on. Without native comparisons, there is evidence of qualitative cohort changes amongst two of the four Canadian male entry cohorts. With native comparisons, in the cases the three earliest male entry cohorts we do find statistically significant qualitative differences across immigrant cohorts. The negative coefficients suggest that the entry wages of these cohorts were some 8-11 per cent *less* than those of the corresponding entry cohort 10 years into the future. Thus, for example, the 1970-74 male entry cohort had entry earnings about 8.1 per cent less than those who entered 10 years later. In other words, these estimates do show an increase in the entry earnings of males in the early-1980s, but not in the latter half of the decade. Furthermore, these changes in immigrant quality began at least as early as the 1970s as evidenced by the significant negative coefficients on the two 1960s cohorts. This result is consistent with the earnings distribution in the United States becoming less equal before the 1980s (Juhn 1999; Juhn, Murphy and Pierce 1993).

[11] An alternative specification of the model including only nonminorities was conducted in an earlier version of this paper which used slightly different data. There were only minor changes in the results.

The results for females are less definitive. The change in cohort quality for females is limited to the first entry cohort of the 1980s. The 1970-74 female entry cohort had entry earnings about 15 per cent less than the average of those who entered between 1980 and 1984. This may be the result of earlier female entry cohorts being more likely to be tied movers than those who emigrated in the 1980s, and thus may have been less likely than more recent female migrants to make decisions based on their own labour market considerations.

Thus, in terms of the entry earnings of the 1980s entry cohorts, the implication of these results is that the 1980-84 entry cohort of males or females had higher entry wages compared to their counterparts that arrived ten years before. We cannot say with certainty, however, that the second entry cohort of the 1980s (1985-90) performed significantly better upon entry. The result that the early-1980s entrants performed relatively well in the United States upon entry, could be due to the longer and deeper recession experienced by Canada at the beginning of the 1980s. Thus, this may reveal a larger-than-normal pool of well-qualified Canadians available to migrate to the United States during this period.

In sum, these results are mixed regarding the qualitative changes in the flow of Canadians to the United States during the 1980s. Immigrants of either gender over the first half of the decade do show significantly larger relative entry earnings compared to those who entered 10 years earlier. Those who entered in the latter half of the decade, however, do not have significantly different entry earnings. In other words, in terms of earnings, there does appear to have been a qualitative improvement in the average Canadian male or female entering the United States in the first, but not the second half, of the decade.

## V. Conclusions

Our results suggest that the much-discussed brain drain from Canada to the United States of the 1990s may have in fact begun much earlier, at least qualitatively. The widening distribution of income and higher returns in to education in the 1980s in the United States relative to Canada provided the motives necessary for an improvement in the quality of this migration. After controlling for individual income generating characteristics and macroeconomic conditions in the United States, we find that the 1980-84 entry cohort of Canadians of either gender displayed significant qualitative improvements as measured by earnings relative to those who entered 10 years earlier. Furthermore, for males, these qualitative

improvements also apply to the two 1970s entry cohorts. Thus, we provide mixed evidence of qualitative differences Canadian migration in the 1980s.

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**Table 1: Summary Statistics for Canadian-born and American-born, Males and Females, U.S. Census, 1980 and 1990**

(standard deviations are in parentheses)

| | Males | | | | | | | | Females | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | 1980 Census | | | | 1990 Census | | | | 1980 Cenus | | | | 1990 Census | | | |
| | Canadian-born | | American-born | | Canadian-born | | American-born | | Canadian-born | | American-born | | Canadian-born | | American-born | |
| *Log of real annual earnings* | 10.411 | (0.615) | 10.259 | (0.594) | 10.454 | (0.656) | 10.226 | (0.623) | 9.646 | (0.638) | 9.611 | (0.629) | 9.813 | (0.685) | 9.682 | (0.665) |
| *Years of education* | 13.007 | (3.253) | 12.883 | (2.891) | 13.997 | (2.897) | 13.436 | (2.511) | 12.492 | (2.409) | 12.608 | (2.426) | 13.633 | (2.408) | 13.402 | (2.213) |
| *Years of experience* | 26.451 | (12.575) | 22.397 | (12.076) | 22.722 | (11.572) | 21.079 | (10.823) | 27.278 | (12.054) | 22.817 | (12.240) | 23.998 | (11.550) | 21.200 | (10.986) |
| *Married* | 0.827 | (0.378) | 0.804 | (0.397) | 0.763 | (0.426) | 0.749 | (0.434) | 0.668 | (0.471) | 0.639 | (0.480) | 0.665 | (0.472) | 0.643 | (0.479) |
| *Weekly hours worked* | | | | | | | | | | | | | | | | |
| 1-19 | 0.013 | (0.113) | 0.015 | (0.122) | 0.006 | (0.077) | 0.007 | (0.085) | 0.066 | (0.248) | 0.053 | (0.223) | 0.048 | (0.213) | 0.042 | (0.201) |
| 20-29 | 0.007 | (0.081) | 0.008 | (0.087) | 0.012 | (0.108) | 0.012 | (0.107) | 0.094 | (0.292) | 0.077 | (0.267) | 0.096 | (0.295) | 0.074 | (0.262) |
| 30-34 | 0.012 | (0.107) | 0.010 | (0.102) | 0.015 | (0.121) | 0.015 | (0.121) | 0.071 | (0.257) | 0.055 | (0.227) | 0.067 | (0.250) | 0.066 | (0.248) |
| 35-39 | 0.044 | (0.206) | 0.035 | (0.184) | 0.028 | (0.166) | 0.029 | (0.168) | 0.124 | (0.329) | 0.127 | (0.333) | 0.100 | (0.300) | 0.104 | (0.305) |
| 40-44 | 0.597 | (0.490) | 0.632 | (0.482) | 0.519 | (0.500) | 0.568 | (0.495) | 0.566 | (0.496) | 0.602 | (0.489) | 0.534 | (0.499) | 0.570 | (0.495) |
| 45-49 | 0.121 | (0.327) | 0.115 | (0.319) | 0.126 | (0.332) | 0.119 | (0.324) | 0.041 | (0.199) | 0.043 | (0.203) | 0.063 | (0.244) | 0.061 | (0.239) |
| 50+ | 0.206 | (0.404) | 0.185 | (0.389) | 0.294 | (0.455) | 0.250 | (0.433) | 0.038 | (0.191) | 0.043 | (0.204) | 0.091 | (0.288) | 0.083 | (0.276) |
| *Occupation* variables | | | | | | | | | | | | | | | | |
| Managerial and professional | 0.352 | (0.478) | 0.280 | (0.449) | 0.388 | (0.487) | 0.257 | (0.437) | 0.273 | (0.446) | 0.235 | (0.424) | 0.379 | (0.485) | 0.297 | (0.457) |
| Technical, sales and administration | 0.183 | (0.387) | 0.191 | (0.393) | 0.224 | (0.417) | 0.214 | (0.410) | 0.476 | (0.499) | 0.470 | (0.499) | 0.434 | (0.496) | 0.448 | (0.497) |
| Service occupations | 0.060 | (0.238) | 0.070 | (0.255) | 0.051 | (0.221) | 0.085 | (0.279) | 0.125 | (0.331) | 0.138 | (0.345) | 0.100 | (0.300) | 0.133 | (0.339) |
| Agriculture, foresting, fishing, etc. | 0.010 | (0.101) | 0.015 | (0.122) | 0.012 | (0.107) | 0.019 | (0.138) | 0.002 | (0.042) | 0.003 | (0.058) | 0.005 | (0.068) | 0.004 | (0.061) |
| Precision production, craft and repair | 0.230 | (0.421) | 0.215 | (0.411) | 0.189 | (0.392) | 0.209 | (0.406) | 0.020 | (0.141) | 0.025 | (0.156) | 0.019 | (0.138) | 0.025 | (0.156) |
| Operators, fabricators and labourers | 0.164 | (0.370) | 0.229 | (0.420) | 0.137 | (0.344) | 0.216 | (0.412) | 0.104 | (0.306) | 0.129 | (0.335) | 0.063 | (0.242) | 0.094 | (0.292) |
| *Industry variables* | | | | | | | | | | | | | | | | |
| Agriculture, forestry, mining, construction | 0.126 | (0.332) | 0.129 | (0.335) | 0.127 | (0.333) | 0.128 | (0.334) | 0.019 | (0.138) | 0.026 | (0.159) | 0.028 | (0.164) | 0.025 | (0.156) |
| Manufacturing | 0.345 | (0.475) | 0.315 | (0.465) | 0.250 | (0.433) | 0.271 | (0.445) | 0.182 | (0.386) | 0.201 | (0.401) | 0.128 | (0.334) | 0.151 | (0.358) |
| Transportation, communications, etc. | 0.092 | (0.290) | 0.124 | (0.330) | 0.094 | (0.292) | 0.116 | (0.320) | 0.048 | (0.215) | 0.053 | (0.225) | 0.047 | (0.212) | 0.054 | (0.227) |
| Wholesale trade and retail trade | 0.141 | (0.348) | 0.146 | (0.353) | 0.159 | (0.366) | 0.165 | (0.371) | 0.196 | (0.397) | 0.181 | (0.385) | 0.184 | (0.387) | 0.185 | (0.388) |
| Services | 0.247 | (0.431) | 0.211 | (0.408) | 0.314 | (0.464) | 0.235 | (0.424) | 0.518 | (0.500) | 0.475 | (0.499) | 0.575 | (0.494) | 0.523 | (0.499) |
| Public administration and active military | 0.048 | (0.215) | 0.074 | (0.262) | 0.056 | (0.229) | 0.085 | (0.279) | 0.036 | (0.187) | 0.064 | (0.245) | 0.039 | (0.193) | 0.062 | (0.242) |
| *Entry Cohort* | | | | | | | | | | | | | | | | |
| 1985-90 | | | | | 0.099 | (0.299) | | | | | | | 0.067 | (0.250) | | |
| 1980-84 | | | | | 0.075 | (0.263) | | | | | | | 0.064 | (0.246) | | |
| 1975-80 | 0.084 | (0.277) | | | 0.075 | (0.263) | | | 0.052 | (0.223) | | | 0.072 | (0.258) | | |
| 1970-74 | 0.051 | (0.219) | | | 0.063 | (0.244) | | | 0.045 | (0.207) | | | 0.068 | (0.252) | | |
| 1965-69 | 0.100 | (0.300) | | | 0.124 | (0.329) | | | 0.101 | (0.302) | | | 0.135 | (0.342) | | |
| 1960-64 | 0.154 | (0.361) | | | 0.194 | (0.395) | | | 0.148 | (0.355) | | | 0.196 | (0.397) | | |
| 1950-59 | 0.288 | (0.453) | | | 0.241 | (0.428) | | | 0.324 | (0.468) | | | 0.279 | (0.449) | | |
| Before 1950 | 0.323 | (0.468) | | | 0.130 | (0.336) | | | 0.329 | (0.470) | | | 0.118 | (0.323) | | |
| *N* | 6,191 | | 16,378 | | 6,080 | | 17,678 | | 5,011 | | 10,546 | | 5,773 | | 14,148 | |

**Table 2: Estimates of the Effects of 10 Years in the United States for Canadians, by Immigration Cohort**
(standard errors are in parentheses)

| | Males | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | | | w/o Native Comparisons | | | | w/ Native Comparisons | | | |
| | Cross-Section | | Within Cohort | | Across Cohort | | Within Cohort | | Across Cohort | |
| 1975-1980 | 0.0062 | (.032) | 0.0146 | (.033) | -0.0084 | (.031) | 0.0178 | (.034) | -0.0116 | (.032) |
| 1970-1974 | 0.0368 | (.036) | 0.0934 | (.039) ** | -0.0566 | (.038) | 0.1179 | (.040) *** | -0.0811 | (.038) ** |
| 1965-1969 | -0.0585 | (.031) * | 0.0056 | (.028) | -0.0641 | (.032) ** | 0.0407 | (.029) | -0.0992 | (.033) *** |
| 1960-1964 | -0.0650 | (.030) ** | -0.0030 | (.022) | -0.0620 | (.031) ** | 0.0416 | (.023) * | -0.1066 | (.032) *** |
| | Females | | | | | | | | | |
| | | | w/o Native Comparisons | | | | w/ Native Comparisons | | | |
| | Cross-Section | | Within Cohort | | Across Cohort | | Within Cohort | | Across Cohort | |
| 1975-1980 | -0.0023 | (.035) | 0.0468 | (.039) | -0.0492 | (.040) | 0.0144 | (.040) | -0.0167 | (.041) |
| 1970-1974 | -0.0772 | (.035) ** | 0.0944 | (.041) ** | -0.1716 | (.042) *** | 0.0727 | (.042) * | -0.1499 | (.042) *** |
| 1965-1969 | 0.0168 | (.030) | 0.0317 | (.028) | -0.0149 | (.033) | 0.0156 | (.029) | 0.0012 | (.033) |
| 1960-1964 | 0.0494 | (.029) * | 0.0812 | (.023) *** | -0.0318 | (.031) | 0.0687 | (.024) * | -0.0193 | (.032) |

Note: Significance levels of 1, 5 and 10 per cent are denoted by ***, **, and *, respectively.

**Table A-1: Number of Canadian-Born in the U.S. Data, by Census Year, Gender and Entry Cohort**

| | 1990 Census | | 1980 Census | |
|---|---|---|---|---|
| | Males | Females | Males | Females |
| 1985-1990 | 603 | 385 | | |
| 1980-1984 | 454 | 372 | | |
| 1975-1979 | 454 | 413 | 517 | 262 |
| 1970-1974 | 385 | 394 | 314 | 224 |
| 1965-1969 | 751 | 782 | 621 | 507 |
| 1960-1964 | 1178 | 1133 | 956 | 742 |
| 1950-1959 | 1466 | 1613 | 1782 | 1626 |
| Before 1950 | 789 | 681 | 2001 | 1650 |
| Total Canadians | 6,080 | 5,773 | 6,191 | 5,011 |
| Total Americans | 17,698 | 14,148 | 16,378 | 10,546 |
| Total | 23,778 | 19,921 | 22,569 | 15,557 |

**Table A-2: Cross-Section OLS Estimates of Log Real Earnings Equations (from joint estimation), U.S. Census, 1980 and 1990**
(absolute values of t-statistics are in parentheses)

| | Males | | | | | | | | Females | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | 1980 Census | | | | 1990 Census | | | | 1980 Cenus | | | | 1990 Census | | | |
| | Immigrants | | Natives | | Immigrants | | Natives | | Immigrants | | Natives | | Immigrants | | Natives | |
| Years of education | 0.0580 | (20.66) | 0.0682 | (36.13) | 0.0595 | (19.14) | 0.0832 | (40.70) | 0.0413 | (11.14) | 0.0485 | (18.92) | 0.0634 | (18.47) | 0.0778 | (33.59) |
| Experience | 0.0450 | (17.99) | 0.0377 | (25.03) | 0.0412 | (15.98) | 0.0314 | (20.20) | 0.0115 | (4.14) | 0.0110 | (6.23) | 0.0119 | (4.66) | 0.0121 | (7.65) |
| Experience squared | -0.0007 | (15.03) | -0.0006 | (19.59) | -0.0006 | (11.39) | -0.0005 | (14.26) | -0.0001 | (2.56) | -0.0001 | (4.03) | -0.0002 | (3.84) | -0.0002 | (5.34) |
| Married | 0.1702 | (9.63) | 0.1603 | (15.41) | 0.1610 | (10.01) | 0.1731 | (18.84) | -0.0742 | (4.93) | -0.0601 | (5.94) | -0.0699 | (5.02) | -0.0323 | (3.71) |
| Weekly hours worked: 1-19 | -0.4829 | (8.32) | -0.2666 | (8.10) | -0.8549 | (9.92) | -0.9992 | (22.05) | -0.9347 | (32.45) | -0.9746 | (44.65) | -1.2420 | (40.07) | -1.2196 | (58.22) |
| 20-29 | -0.8054 | (9.96) | -0.6323 | (13.73) | -0.6964 | (11.36) | -0.7762 | (21.39) | -0.6616 | (26.65) | -0.6966 | (37.53) | -0.6664 | (29.05) | -0.6798 | (41.68) |
| 30-34 | -0.2315 | (3.79) | -0.3336 | (8.43) | -0.2970 | (5.44) | -0.5159 | (16.14) | -0.3814 | (13.66) | -0.3887 | (18.11) | -0.4414 | (16.54) | -0.3984 | (23.27) |
| 35-39 | 0.0070 | (0.22) | -0.0037 | (0.17) | -0.0870 | (2.17) | -0.0547 | (2.35) | -0.0344 | (1.58) | -0.0811 | (5.48) | -0.0884 | (3.97) | -0.0723 | (5.20) |
| 45-49 | 0.0654 | (3.18) | 0.1221 | (9.46) | 0.0897 | (4.31) | 0.1224 | (9.91) | 0.1614 | (4.58) | 0.0853 | (3.60) | 0.1566 | (5.78) | 0.1762 | (10.02) |
| 50+ | 0.1487 | (8.73) | 0.1469 | (13.67) | 0.2186 | (14.05) | 0.1932 | (20.47) | 0.1843 | (5.00) | 0.1300 | (5.45) | 0.2111 | (9.06) | 0.2282 | (14.79) |
| Cohort: 1985-90 | | | | | 8.6699 | (136.61) | | | | | | | 8.7741 | (122.51) | | |
| 1980-84 | | | | | 8.6234 | (132.37) | | | | | | | 8.8352 | (120.17) | | |
| 1975-80 | 8.8102 | (140.91) | | | 8.6761 | (131.79) | | | 9.0400 | (116.78) | | | 8.7717 | (120.15) | | |
| 1970-74 | 8.6973 | (132.22) | | | 8.6603 | (131.15) | | | 8.9640 | (112.90) | | | 8.7579 | (118.95) | | |
| 1965-69 | 8.7373 | (139.69) | | | 8.6176 | (137.18) | | | 9.0584 | (118.10) | | | 8.7886 | (124.43) | | |
| 1960-64 | 8.7191 | (141.85) | | | 8.5953 | (139.12) | | | 9.0155 | (118.79) | | | 8.8074 | (124.33) | | |
| 1950-59 | 8.7338 | (145.32) | | | 8.5292 | (134.21) | | | 9.0137 | (118.50) | | | 8.7959 | (122.11) | | |
| Before 1950 | 8.6718 | (138.69) | | | 8.5187 | (127.69) | | | 8.9867 | (115.20) | | | 8.7546 | (117.62) | | |
| CONSTANT (natives only) | | | 8.6856 | (243.36) | | | 8.3501 | (237.19) | | | 8.9758 | (190.60) | | | 8.5288 | (202.79) |
| *N* | 6,191 | | 16,378 | | 6,080 | | 17,698 | | 5,011 | | 10,546 | | 5,773 | | 14,148 | |

Note: Results are from joint estimation (by gender) and include occupation and industry dummies, but no common constant term. $R^2$ values have been surpressed because of the estimation with no constants.