Changes in the Quality of Immigrant Flows between the United States and Canada in the 1980s

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Introduction and Background

It is well known that individuals will desire to migrate to the location where returns to their abilities are the highest, other things equal. An important theoretical factor in determining these returns is the relative distribution of earnings in the source and host countries. If the earnings variance in the country of origin is less than that in the country of destination (assuming that the correlation in earnings across the two countries is positive and sizeable), we expect those from the upper tail of the distribution (that is, those with the highest skills) to migrate, since incomes will be maximized by doing so. Conversely, if the source country has a wider distribution of income than the host country, we would expect those from the lower tail of the distribution (that is, those with the lowest skills) to migrate, since they too would have higher earnings in the host country. In the context of a bilateral immigration relationship, this would then imply that the country with the wider distribution of income would have high-quality immigrants wishing to enter, while the country with the more egalitarian distribution would have lower quality immigrant flows.

Borjas (1988, 1993) has addressed immigration between Canada and the United States within the context of such a wealth-maximization model. He found that Canadians tend to perform well in the United States (in terms of earnings) relative to native Americans, while Americans tend to perform relatively poorly in Canada, even when controlling for different human capital characteristics.[1] He hypothesized the less equal distribution of income in the United States relative to Canada could cause this outcome and that individuals seeking to maximize wealth will self-select into the appropriate economy. Thus, insofar as people are free to move between the two countries, Canadians of high ability will choose to migrate to the United States where returns to their abilities are higher, while Americans of more limited ability will seek to enter Canada, since they will earn higher wages than they could in the United States. This result is somewhat puzzling given that Canada has, since the 1960s, explicitly followed an immigration policy that has sought to enhance the quality of

The American Review of Canadian Studies (Winter 1999): 621-649

immigrants (from all source countries) by screening out those who may have a limited ability to assimilate into the Canadian labor market. Borjas, however, argues that the screening process in Canada can only evaluate immigrants on the basis of observable characteristics, not unobservable characteristics, even though the latter are also important in determining earnings.[2] He asserts that: "In the end, regardless of what immigration policy says, only those persons who gain from immigration do so. Governments legislate, but it is people who immigrate" (Borjas 1990, 216).

It has also been well established that there has been an increase in earnings inequality in the United States in the 1980s. The review article by Levy and Murnane (1992) shows that earnings inequality increased in the United States for both males and females in the 1980s. Furthermore, this trend towards increased inequality has favored high-skilled workers. In Canada, earnings inequality also increased over the same period (Morissette, Myles, and Picot 1993; Morissette and Bérubé 1996; Burbridge, Magee, and Robb 1997). In comparing the changes in the earnings distribution in the two countries, however, the literature suggests that this increase in earnings dispersion has been much greater in the United States (Blackburn and Bloom 1993; Gottschalk and Smeeding 1997; Richardson 1997).[3]

This change in the relative earnings distribution should be manifest in the form of varying immigrant quality on either side of the border. In this context, the implication of the wealth-maximization model is that the United States should attract Canadians with greater skills (and hence higher earnings) than natives, and that Americans with lower skills (and hence lower earnings) compared to natives would continue to migrate to Canada. In other words, we would expect that the 1980s cohort of Americans who immigrated to Canada would be qualitatively inferior (in terms of earnings) to those Canadians who immigrated to the United States during that same time period when we control for observable labor market characteristics.

The specific objectives of this paper are: first, to address the economic assimilation of Canadians in the United States and Americans in Canada during the 1980s; second, in doing so to test the robustness of the Borjas assertion that the types of immigrant selection biases are, at least in part, due to the relative earnings dispersion; and, third, to explicitly include female migrants in the analysis. The international migration literature, in general, and the relevant work by Borjas, in particular, focus on the earnings assimilation of prime-age male immigrants, but there is increasing evidence that female immigrants may have different labor market experiences than male immigrants (Baker and Benjamin 1997; Beach and Worswick 1993; Fagnan 1995).

Methodology

To ascertain whether the wealth-maximization model is correct in determining the immigration patterns between the two countries in the 1980s, the now-common methodology of using synthetic panels will be employed. The initial research on the economic adaptation of immigrants (Chiswick 1978) used cross-sectional data from the United States to estimate the earnings disadvantage that immigrants experienced at the time of entry into the U.S. labor market (the entry effect), as well as the rate of growth of earnings that exceeded the growth rate of native earnings (the assimilation effect). The use of only one cross-section, however, does not permit the estimation of qualitative differences between successive entry cohorts (the cohort effect). The reason for this is simple. In a cross-section of data, we view each immigrant at time t. If the immigrant has been in the host country for ten years, his/her year of entry (*YOE*) would be t-*10* and the number of years since migration (*YSM*) would equal ten. Thus, if the observation is drawn from the 1990 U.S. census, for example, $t = 1990$, *YSM* = 10, and *YOE* = 1990 - 10 = 1980. In other words, *YSM* = t - *YOE*. Thus, *YSM* and *YOE* are perfectly correlated and one of these independent variables must be dropped to successfully estimate the model. Since we are normally concerned with the rate of assimilation of immigrants in the population (that is, the coefficient on *YSM*), the coefficients on *YOE* variables are constrained to be equal. Consequently, we cannot determine if there are qualitative differences in immigrants based on their entry cohort. This is a serious limitation, given that changes in U.S. and Canadian immigration policies have been shown to result in qualitative differences between immigrant cohorts.[4]

To circumvent this problem, Borjas (1985) advocated using synthetic-panel data, whereby observations are drawn from two separate cross-sectional data sets. These are not true panel data, but they do allow the estimation of both assimilation and cohort effects.[5] We use a slight variation of this model. While Borjas used separate equations to estimate native and immigrant earnings, we follow the lead of Bloom, Grenier and Gunderson (1995) who used the basic empirical model of Chiswick (1978) augmented to allow for the estimation of cohort-specific effects. The main difference is that the socioeconomic characteristics of natives and immigrants are constrained to be the same in each case. Thus, for example, estimation of a single equation would imply that a year of education has the same effect on earnings for both natives and immigrants. This could be a serious limitation if the group of immigrants was more heterogenous, but given the structural similarities of the two countries, it seems like a reasonable assumption for our purposes.[6] Thus, the following regression model is estimated:

$$\ln earn = X\beta + \alpha IMMIG + \eta YSM \cdot IMMIG + \textstyle\sum_j \delta_j C_j \cdot IMMIG + \lambda CENS \tag{1}$$

where ln*earn* is the natural logarithm of the nominal yearly earnings and X is a vector of individual socioeconomic characteristics for immigrants and natives that contains information on years of education and experience, marital status, and residence in an urban area. We also control for the number of weeks worked and part-time status. Although not a perfect measure, it has been shown that increases in inequality of weekly earnings in Canada have in part been the result of increased inequality in weekly hours worked (Morissette, Myles and Picot 1993; Morissette 1995; Doiron and Barrett 1996). The vector β is the estimate of returns to these skills, which is constrained to be equal for both immigrants and natives; *YSM* is the number of years that the immigrant has resided in the host country; *IMMIG* is a dummy variable equal to one for immigrants; and C_j is a dummy variable for the entry cohort of the immigrant.[7] Finally, the variable *CENS* is a dummy coded to one if the observation was drawn from the 1990 U.S. census or 1991 Canadian census.

This equation gives estimates of the entry effect (α), which shows the earnings premium of immigrants at the time of landing; the assimilation effect (η), reflecting the rate at which immigrant earnings converge towards those of natives; and the cohort effect (δ_j), which reflects qualitative differences in the earnings of various entry cohorts. Finally, the period effect (λ) reflects changes in the economy over the intercensal period, which are (necessarily) assumed to be equal for immigrants and natives. Following Bloom, Grenier, and Gunderson (1995), we can also estimate the years to earnings parity with natives from these coefficients by adding the entry and cohort effects and dividing by the assimilation effect; that is, $-(\alpha + \delta_j) / \eta$.

Data

The Canadian data are drawn from the individual files of the 1981 (2 percent) and 1991 (3 percent) Canadian censuses. In each case, all American-born are retained, while a 1/10 subsample of the Canadian-born was randomly chosen. The United States data are obtained by merging the 5 percent (Sample A) housing and individual records of the 1990 U.S. census with the 1/1000 sample (Sample B) of the 1980 U.S. census. All observations were kept from the 1980 sample, while a 1/100 subsample of the American-born was randomly generated from the 1990 sample. All Canadian-born individuals were retained from each sample.[8]

The samples were further limited to include only non-institutionalized individuals between the ages of twenty-five and sixty-four who worked at least one week in the year prior to the census, were not self-employed, did not attend school, and had at least $1000 in (nominal and local currency) earnings in the reference calendar year. The income variable is the natural logarithm of annual earnings. This includes wage and salary income, as well as self-employment income. Although we eliminated the self-employed from the samples in both countries, those who are primarily engaged in paid employment may still have some income from self-employment. To keep the income variables comparable between countries, this was the variable we decided to utilize.

The years of education variable (*YEARSED*) was coded to equal the number of years corresponding to the highest level education completed. For example, high school graduation or GED was coded to equal twelve, some post-secondary education or an associate degree was coded as fourteen years of education, while a bachelor's degree was coded as sixteen, a master's or professional degree as eighteen years, and a doctorate as twenty. Experience (*EXP*) was calculated using the familiar Mincer proxy (*AGE* - *YEARSED* - 6). In both the U.S. and Canadian data the age variable is continuous, but since the variable *YEARSED* was derived from discrete intervals, *EXP* was negative in a few cases and was therefore bottom-coded to zero.

The variable *MARRIED* was coded to one if the respondent said that he or she was married, and zero otherwise. *URBAN* is a dummy variable reflecting residency in an urban area. These definitions differ somewhat between the two countries. *WKSWK* is simply the number of weeks worked by the individual in the year preceding the census. Finally, the variable *PTWORK* is a dummy that was set to equal one if the respondent worked less than thirty hours per week. This variable did exist in the Canadian data and was derived from the U.S. data to be consistent with the Canadian definition.

For immigrants to Canada, the cohort *YPRE61* was coded to one for those Americans who entered Canada before 1961, while Y61 was coded to one for those who entered during the five-year period beginning in 1961 (1961-65), Y66 set to equal one for those who entered between 1966 and 1970, and so forth. The coding of the American data was done in a similar fashion (*YPRE60*, *Y60*, *Y65*, ...), reflecting the different coding in the raw U.S. data. The variable years since migration (*YSM*) was obtained by subtracting the year at the midpoint of the entry cohort from the year of the census. For example, if an immigrant to Canada is from the Y61 cohort and the observation was drawn from the 1991 census, then *YSM* would equal twenty-eight (that is, 1991 less 1963, the midpoint of the 1961-65 entry cohort).[9]

The final Canadian male sample consists of 2,187 American-born and 17,895 Canadian-born individuals. The numbers for American-born and Canadian-born females in the Canadian sample are 2,066 and 14,044, respectively. In the U.S. data, there are a total of 54,647 American-born and 6,965 Canadian-born males. The numbers for females in the U.S. data are 43,677 American-born and 7,404 Canadian-born.

Table 1 presents the summary statistics for males and females in the Canadian sample. American males earned more than their Canadian-born counterparts—about 13 percent more in 1981, and about 15 percent more in 1991. This could be explained by the fact that the average American-born male had more than two more years' additional education in each year. Higher annual earnings, however, cannot be explained by a larger number of weeks worked nor by a higher incidence of full-time work compared to Canadians.

American females in Canada, by contrast, are more comparable to their Canadian-born counterparts. In addition to the higher earnings of American-born females (2 percent in 1981 and 9 percent in 1991), they had about 1.5 years more education compared to Canadian women in each of the two censuses. Both groups in both years are comparable in terms of experience, weeks worked, and other characteristics.

Table 2 shows the summary statistics for males and females in the U.S. data. Canadian males earned about 15 percent more than American males in 1980, increasing this advantage to some 23 percent in 1990. This difference, however, is unlikely the result of higher levels of educa-

TABLE 1: Summary Statistics for U.S.-born and Canadian-born, Canadian Census, 1981 and 1991 (standard deviations are in parentheses)

		U.S.-born Males		Canadian-born Males		U.S.-born Females		Canadian-born Females	
Variable	Description	1981	1991	1981	1991	1981	1991	1981	1991
LNEARN	Log of total earnings in previous year	9.859	10.416	9.730	10.261	9.057	9.755	9.036	9.669
		(.737)	(.786)	(.666)	(.733)	(.854)	(.913)	(.793)	(.841)
YEARSED	Years of education	14.24	14.73	11.92	12.64	13.58	14.15	12.05	12.72
		(3.43)	(3.09)	(2.93)	(2.60)	(2.88)	(2.49)	(2.46)	(2.25)
EXP	Years of experience	22.98	21.67	22.54	21.31	21.30	20.79	21.26	20.59
		(12.24)	(10.23)	(12.29)	(10.94)	(12.39)	(10.30)	(11.83)	(10.62)
MARRIED	Married or common law	0.82	0.79	0.81	0.77	0.77	0.79	0.74	0.75
URBAN	Resident of CMA	0.58	0.65	0.47	0.56	0.54	0.64	0.52	0.57
WKSWK	Weeks worked in previous year	46.72	46.82	46.20	45.84	41.60	43.78	42.70	43.84
		(10.44)	(10.67)	(10.89)	(11.62)	(13.49)	(13.13)	(13.64)	(13.15)
PTWORK	Part-time employment in previous year	0.04	0.04	0.04	0.04	0.29	0.26	0.27	0.24
YSM	Years since migration	18.99	21.24	...	...	18.45	20.09	...	...
		(10.92)	(10.73)			(11.13)	(10.98)		
Y86	Immigrated 1986-91	...	0.07	...	...	...	0.08	...	...
Y81	Immigrated 1981-85	...	0.09	...	...	...	0.12	...	...
Y76	Immigrated 1976-80	0.10	0.09	...	...	0.14	0.14	...	...
Y71	Immigrated 1971-75	0.20	0.25	...	...	0.20	0.24	...	...
Y66	Immigrated 1966-70	0.21	0.24	...	...	0.18	0.17	...	...
Y61	Immigrated 1961-65	0.06	0.10	...	...	0.08	0.10	...	...
YPRE61	Immigrated before 1961	0.43	0.15	...	...	0.41	0.14	...	...
Number of observations		932	1255	6331	11564	690	1376	4238	9806

TABLE 2: Summary Statistics for American-born and Canadian-born, United States Census, 1980 and 1990 (standard deviations are in parentheses)

		U.S.-born Males		Canadian-born Males		U.S.-born Females		Canadian-born Females	
Variable	Description	1980	1990	1980	1990	1980	1990	1980	1990
LNEARN	Log of total earnings in previous year	9.631	10.108	9.781	10.337	8.869	9.478	8.955	9.584
		(.680)	(.753)	(.691)	(.782)	(.739)	(.834)	(.697)	(.872)
YEARSED	Years of education	12.67	13.34	12.83	13.90	12.56	13.38	12.53	13.58
		(3.20)	(2.54)	(3.42)	(2.92)	(2.66)	(2.27)	(2.95)	(2.45)
EXP	Years of experience	22.31	21.31	26.37	23.04	22.14	21.23	27.06	24.07
		(12.32)	(11.04)	(12.46)	(11.78)	(12.17)	(11.07)	(12.62)	(11.65)
MARRIED	Married	0.79	0.73	0.85	0.75	0.67	0.66	0.69	0.68
URBAN	Resident of SMSA	0.76	0.79	0.84	0.88	0.76	0.80	0.84	0.87
WKSWK	Weeks worked in previous year	48.34	48.12	49.50	47.99	44.49	45.28	46.77	44.86
		(8.61)	(9.27)	(7.25)	(9.44)	(11.91)	(11.71)	(10.76)	(12.07)
PTWORK	Part-time employment in previous year	0.03	0.03	0.04	0.03	0.16	0.15	0.22	0.18
YSM	Years since migration	...	...	22.11	26.21	...	...	23.76	27.24
				(9.33)	(13.03)			(8.97)	(12.44)
Y85	Immigrated 1985-90	...	...	...	0.10	...	...	...	0.07
Y80	Immigrated 1980-84	...	...	...	0.07	...	...	...	0.06
Y75	Immigrated 1975-79	...	...	0.09	0.07	...	...	0.07	0.07
Y70	Immigrated 1970-74	...	...	0.05	0.06	...	...	0.03	0.07
Y65	Immigrated 1965-69	...	...	0.08	0.12	...	...	0.10	0.13
Y60	Immigrated 1960-64	...	...	0.25	0.19	...	...	0.16	0.19
YPRE60	Immigrated before 1960	...	...	0.53	0.38	...	...	0.64	0.40
Number of observations		34783	19864	136	6829	26032	17645	135	7269

tion amongst the Canadian-born. It is noteworthy that the average Canadian in the 1990 sample had more than one additional year of education compared to the average Canadian in the 1980 sample.[10] Given that the educational earnings differential (the ratio of returns to college education to the returns to high school education) increased more in the United States than in Canada over our period of analysis (Freeman and Needels 1993), this is the pattern we expected to see. Canadian males possess two to four more years of labor market experience than native American males. All other characteristics are similar between the two groups.

The U.S. data largely mirror the Canadian data in that there is more homogeneity in observable characteristics between native and immigrant females compared to males. Canadian-born females earned more than natives in both years—a 9-percent differential in 1980 and an 11-percent differential in 1990. The average Canadian female in the United States does not have any more years of education in either year relative to native females, but she does have about one additional year of education in 1990 compared to 1980. Again, given the increase in the wage differential between college-educated and high school-educated individuals in the United States, we do expect that Canadians who recently entered the U.S. would have more education then their compatriots from previous cohorts. Finally, Canadian women on average have five years more experience in 1980 and three more years in 1990 compared to the average American female. This may prove useful in explaining the difference in earnings.

These summary data for males are comparable to those outlined in Borjas (1988). He analyzed 1970 and 1980 U.S. census data, as well as 1971 and 1981 Canadian census data for working-age males. Thus, when our data overlap with his (the 1980 U.S. census and the 1981 Canadian census), we find no obvious differences in the summary statistics. This is important given the small number of Canadians in our 1980 U.S. census sample. In his sample, Canadians in the United States earned hourly wages about 15 percent higher than natives in 1980, but education levels were essentially the same between the two groups. Borjas also found that Americans residing in Canada earned only about 10 percent more than natives on average, despite having about three additional years of education. Thus, according to Borjas, both countries are attracting immigrants from the other side of the border with high levels of education, but since Canadians in the United States earned significantly higher salaries than comparable Americans, he argued that Canada exports, and the United States imports, individuals with high levels of ability or unobserved skills.

Conversely, Canada imported Americans with high levels of observable skills, but low levels of unobservable skills.

Borjas reasoned that this type of immigration pattern could be the result of the relatively less unequal distribution of income in Canada compared to the United States. As a result, Americans of lesser unobserved ability will choose to migrate to Canada since they will not be penalized as heavily for this lack of ability. Somewhat ironically, the points system in Canada is limited in its ability to screen out qualitatively inferior immigrants, since it focuses almost exclusively on observable characteristics such as education, experience, and language ability. Conversely, Canadians of high unobserved ability will go the United States, since they will be rewarded for these skills. Given that the relative distribution of earnings continued to worsen throughout the 1980s in the United States, we expect to see improvements in cohort quality amongst Canadians migrating south, and qualitative declines in Americans entering Canada throughout the decade. If the underlying differences in the distribution of earnings are indeed responsible for these types of immigrant selection biases within North America, the analysis in the next section should shed some light on this.

Empirical Results

To investigate this matter further and to check for results consistent with those of Borjas, we proceed with ordinary least squares (OLS) estimation of the standard earnings equation. We first conduct this exercise by estimating the earnings equation in the cross-section. As outlined above, it is well known that cohort and assimilation effects cannot be simultaneously estimated in a single cross-section, since there is no way of knowing whether differences in earnings over time are the result of qualitative differences in cohorts, or differences in the rate of assimilation of the immigrant population into the new labor market. Thus, we pool the two cross-sections from either country to form a synthetic panel. In this way, we can estimate both cohort and assimilation effects, although the period effect (which is due to changes in technology, the business cycle, etc.) must be constrained to be equal for both immigrants and natives.

Tables 3 and 4 show the highlights of the regression results for males and females using the natural logarithm of annual earnings as the dependent variable.[11] The cross-section results suggest that the average male immigrant in Canada actually improved his earnings capacity at the time

TABLE 3: Entry, Assimilation, and Cohort Effects, and Corresponding Years to Equality Calculations, Males (absolute values of t-statistics are in parantheses)

	Canada				United States			
	1981	1991	Pooled	Years to Equality	1980	1990	Pooled	Years to Equality
Entry effect	-0.1789	-0.1149	-0.0826		0.2275	0.2314	0.1960	
	(5.123)	(3.160)	(1.442)		(1.882)	(14.242)	(7.551)	
Assimilation effect	0.0076	0.0036	0.0059		-0.0082	-0.0033	0.0101	
	(4.912)	(2.426)	(2.186)		(1.636)	(6.132)	(2.076)	
Cohort effect								
Y86/Y85				14.05				-19.34
Y81/Y80			-0.0973	30.60			-0.0654	-12.89
			(1.263)				(1.598)	
Y76/Y75			-0.1055	31.99			-0.0824	-11.21
			(1.506)				(1.426)	
Y71/Y70			-0.0541	23.25			-0.1809	-1.49
			(0.783)				(2.258)	
Y66/Y65			-0.0876	28.95			-0.2415	4.48
			(1.153)				(2.387)	
Y61/Y60			-0.0656	25.21			-0.3263	12.85
			(0.717)				(2.640)	
YPRE61/YPRE60			-0.0831	28.19			-0.4982	29.81
			(0.789)				(2.754)	
Years to equality	23.64	31.59			27.59	69.82	86.91	
R^2	0.4475	0.4149	0.4889		0.3558	0.4489	0.4676	
N	7263	12819	20082		34919	26693	61612	

TABLE 4: Entry, Assimilation, and Cohort Effects, and Corresponding Years to Equality Calculations, Females (absolute values of t-statistics are in parantheses)

	Canada				United States			
	1981	1991	Pooled	Years to Equality	1980	1990	Pooled	Years to Equality
Entry effect	-0.1178	-0.0947	-0.0401		0.1878	0.1271	0.0995	
	(2.774)	(2.696)	(0.701)		(1.486)	(7.553)	(3.600)	
Assimilation effect	0.0028	0.0022	0.0060		-0.0067	-0.0010	0.0066	
	(1.478)	(1.484)	(1.908)		(1.345)	(1.758)	(1.389)	
Cohort effect								
Y86/Y85				6.68				-15.01
Y81/Y80			-0.1749	35.83			0.0205	-18.10
			(2.383)				(0.487)	
Y76/Y75			-0.1153	25.89			-0.0734	-3.94
			(1.655)				(1.274)	
Y71/Y70			-0.0860	21.01			-0.1428	6.53
			(1.170)				(1.815)	
Y66/Y65			-0.1087	24.79			-0.1258	3.97
			(1.304)				(1.269)	
Y61/Y60			-0.1629	33.83			-0.1620	9.43
			(1.640)				(1.331)	
YPRE61/YPRE60			-0.2027	40.46			-0.2856	28.08
			(1.684)				(1.613)	
Years to equality	41.50	42.43			28.05	130.88		
R^2	0.5396	0.4968	0.5629		0.5106	0.5546	0.5951	
N	4928	11182	16110		26167	24914	51081	

of entry. Compared to Canadian males, his annual earnings were about 11 percent less in 1991, up from an earnings disadvantage of almost 18 percent in 1981. Certainly this is contrary to our expectations. The assimilation effect estimates also differ between the two cross sections. In both instances they are small, but the estimates imply that immigrants in 1981 assimilate twice as quickly as those in 1991 (.76 versus .36 percent per year). With this information, it is easy to determine the number of years it will take for the average American age-earnings profile to intersect with that of native Canadians. Following Bloom, Grenier, and Gunderson (1995) this is obtained by dividing the entry effect by the assimilation effect.[12] Thus, the average American male in Canada as of 1981 would take almost twenty-four years to equal the earnings of comparable natives, while the same process would take more than thirty-one years for those in the 1991 sample. By these measures, it appears that there has been a decline in the ability of Americans to assimilate into the Canadian labor market. Although more recent immigrants do start off with less of an earnings disadvantage, they assimilate at a slower rate.[13]

The results from the pooled data reflect those obtained from a comparison of the cross-sectional results. Namely, there appears to be an improvement in the quality of immigrants who entered Canada after 1981. The entry effect coefficient suggests that the most recent wave of American males entered Canada with an 8-percent earnings disadvantage relative to comparable Canadians. Since the coefficient is statistically insignificant, however, we cannot say with certainty that their earnings actually differed from Canadian earnings. Nor can we be certain that earlier cohorts differ from the most recent cohort, since the coefficient estimates are not significant, either individually or jointly.[14] What we can say is that the earlier cohorts did differ from Canadians in their entry wages. This is because the entry effect plus the cohort effect is significantly different from zero at the one-percent level for Y66 through Y81, and at the 10-percent level for YPRE61 and Y61.[15]

We can also estimate the number of years each cohort will take to equal the earnings of demographically comparable Canadians. These estimates reflect those above in not supporting the hypothesis that there has been a qualitative decline in immigrants' earnings in the 1980s. Quite to the contrary. A simple calculation implies that the most recent cohort would require approximately fourteen years to catch up to the earnings of the average Canadian male, although this is not significantly different than zero.[16] For all previous cohorts, however, the number of years to equality ranges from twenty-three to thirty-two years, and these are all significant at at least the 5-percent level.[17]

Thus, when we compare the most recent entry cohort to earlier cohorts, we cannot ascertain any significant change in cohort quality. By comparing the earnings of all cohorts relative to those of native Canadians, all entry cohorts from Y66 to Y81 have significantly lower earnings than otherwise comparable Canadians. This suggests that the most recent cohort of Americans in Canada is not at the same earnings disadvantage as earlier American immigrant waves. Indeed, the most recent wave has no discernible difference in relative earnings.

These results are contrary to our expectations and also somewhat inconsistent with the more recent immigration literature. Bloom, Grenier, and Gunderson (1995) group immigrants from the United States and Europe into one category and find some evidence of declining cohort quality, especially for the 1981-85 cohort (the most recent data set used by these authors was 1986). They argue that this is likely the result of the inability of the labor market to absorb immigrant labor during recessionary periods. However, since our pooled data include 1991 data, we see that the 1981-85 entry cohort did not perform much differently in the Canadian economy relative to earlier cohorts, although they did perform somewhat (albeit insignificantly) worse than the most recent entry cohort. Our results could be due to a general increase in the quality of Americans choosing to enter Canada, but the recession hindered the earnings potential of entrants between 1981 and 1985.[18]

The American data for males yield quite different results. Both cross-section estimates show that the entry effect for Canadian males is positive, meaning that Canadian males enter the U.S. labor market with higher earnings than comparable American males. A small negative assimilation effect implies that it would take a number of years for the age-earnings profile of American males to catch up to the earnings of the Canadian-born (or that the earnings of Canadian males equaled those of Americans at some time before migration).

The pooled data paint a more realistic picture of the age-earnings profile of immigrants. The most recent entry cohorts begin with a large earnings advantage, and then increase this advantage over native males by about one percent annually.[19] For example, those who entered between 1985 and 1990 did so with about 20 percent higher earnings than comparable Americans. Those who entered between 1970 and 1975 did so with only a small earnings advantage, and those who entered before 1970 had earnings below those of comparable American males. Only for those who entered after 1980, however, are these differences significantly different than zero, meaning that these individuals did have higher earnings than similar American males.[20] These cohort differences are also reflected in the years-to-equality calculations. These results are supportive of the hypothesis that the relative widening of the distribution of earnings in the United States resulted in an improvement in immigrant quality. In sum, the post-1980 wave of Canadian males performs significantly better in the U.S. labor market relative to both the American-born and also relative to previous waves of Canadians.

These results are not entirely consistent with those of Borjas (1988). In the U.S. data, he discovered that earlier cohorts of Canadian immigrants had significantly lower wages compared with more recent cohorts. In other words, the quality of Canadians (based on unobservable characteristics) entering the U.S. had been steadily improving. As expected, we find that this qualitative improvement has continued into the 1980s, although the most recent wave of immigrants does not enjoy earnings that are significantly greater than the earnings of those who have entered since 1975. Conversely, Borjas found that the quality of Americans entering Canada had been steadily deteriorating through successive cohorts. We find no significant evidence of a general deterioration of immigrant quality into the 1980s. To the contrary, the trend is towards a modest improvement, with the most recent cohort having earnings statistically indistinguishable from native earnings.

The female data in Table 4 exhibit much the same pattern as the data for males, although both assimilation and entry effects tend to be smaller than comparable estimates in the male data.[21] For Americans in Canada, the estimates of the entry and assimilation effect are not suggestive of a decline in immigrant quality, at least in the cross-section. Similarly, the pooled estimates show no distinct pattern of cohort quality deterioration. The coefficients on each entry cohort are essentially zero, both individually and jointly.[22] The exception is the 1981-85 entry cohort which did have earnings about 17.5 percent lower than the most recent cohort. This result could be due to the sensitivity of immigrant earnings to the recession in Canada in the early-1980s, along with the fact that female immigrants from the United States are more likely than males to be non-independent immigrants and therefore were not admitted on the basis of their labor market credentials. This result is consistent with De Silva (1997), who compared entry cohort differences within various immigration classes. He found that immigrants who entered as independent immigrants suffered a less dramatic decline in entry wages upon landing during the recession of the early 1980s compared to those who did not enter as independents.[23] Compared to Canadians, however, we can say with some certainty that all but the most recent wave of immigrants have performed worse in the Canadian labor market.[24]

For Canadian females in the United States, the cross-section estimates show little evidence of relative earnings differences between the two censuses. As with Canadian males in the United States, Canadian females have a positive entry effect and a small (though insignificant) assimilation effect in the pooled data. The entry cohort coefficients, however, are individually and jointly insignificant, indicating that there has been no change in cohort quality.[25] Relative to the earnings of American females, however, Canadian women who entered in the 1980s do have significantly higher earnings.[26]

In sum, although there is no evidence of changes in cohort quality (with the exception of the Y81 cohort), American women, like American males, appear to be narrowing the gap between themselves and the Canadian-born. For Canadian women in the United States, we cannot say with any degree of certainty that Canadians have improved their earnings relative to earlier immigrant waves, but relative to the native-born their position did improve during the 1980s.

Discussion

We do find some evidence of changes in the cohort quality of immigrants between the United States and Canada over the 1980s. In the U.S. data, our estimates generally give us the expected results and are supportive of the hypothesis that the relative widening of the earnings distribution in the United States is responsible for these types of immigrant selection biases. In Canada, however, the smaller increase in the dispersion of earnings has not resulted in a decline in the earnings capacity of the most recent American immigrants. To the contrary, there appears to be a modest, though statistically insignificant, change in the earnings of successive waves of immigrants of either gender. This gradual improvement has resulted in the most recent cohort of Americans (both males and females) having earnings that are not statistically different from those of native Canadians on average.

Since the implications of the wealth-maximization have not been satisfied in this case, we must provide some alternative explanation for these results. It may be worthwhile to take a look at changes in Canadian immigration policies during this period. A significant policy change was made in April 1978 as the Immigration Act of 1976 was implemented.[27] This change, in essence, increased the importance of family reunification in Canadian immigration policy, while limiting the size of the group entering with explicit labor market skills. In 1982, in the midst of the recession, the Canadian government announced plans to cut overall immigration levels in the following year and also called for an immediate change, which required non-sponsored applicants to have prearranged employment before being allowed to enter Canada. This change substantially shifted the flow towards non-assessed immigrants, since they did not have to clear this new hurdle. In 1985, the emphasis was shifted from education qualifications to explicit employment criteria. Arranged employment was required for admission in the independent class. Also, job certification requirements were circumvented by increased admissions for the purpose of family reunification. Assisted relatives were assessed on the basis of points, but were given a five-point bonus for family affiliation. In 1986, the arranged-employment restriction was relaxed for all independent applicants, but certain occupations were ineligible for recruitment outside Canada. Also in this year, the business immigration program was established to promote the immigration of entrepreneurs, self-employed persons, and investors.[28] Finally, in 1988 the government expanded the family class to include previously ineligible relatives. This increased the number of unscreened applicants directly, and also indirectly, as this group had processing priority over screened immigrants. Thus, these policy changes had the effect of shifting the inflow substantially toward non-assessed immigrants, since those who desired entry based on the points system were, in essence, relegated to the residual group of entrants.[29]

The effect of these policies on the entry-class composition of Americans is shown in the Canadian data. In 1982, some 60.1 percent of all American males entered Canada as independents, with an additional 37.6 percent entering as family members. By 1991, these positions had changed, as only 39.7 percent entered as independents while family class immigrants increased to 47.4 percent. For females entering Canada, this change is less dramatic but still evident. In 1982, 43.4 percent were admitted as independents, while an additional 54.1 percent were family class. By 1991, these numbers had changed to 29.6 percent and 60.6 percent respectively.[30]

Given this change in the composition of immigrants from the United States, we might expect this group to have similar declines in cohort quality in accord with other recent studies that have addressed all immigrant groups. Following our estimates, however, we do not see a pattern of earnings degradation amongst the most recent waves of American immigrants. To the contrary, the data for both males and females are indicative of an improvement in the earnings ability of Americans in Canada. With the exception of a dip in the earnings of American women who entered Canada during the recession of the early 1980s, we find no evidence of a decline in earnings over this period. This is contrary to our expectations, both in terms of the theory and immigration policy changes, which have unambiguously pointed towards an immigration flow that should be negatively selected from the American-born population.

It was also thought that the Canada-U.S. Free Trade Agreement (FTA) might have had some effect on immigrant quality—in particular, the curious result that American males in the 1986-90 cohort showed a qualitative improvement relative to earlier cohorts. The FTA took effect on 1 January 1989 and made migration between the two countries easier for various well-defined groups of workers. We tested for this by disaggregating the 1986-90 cohort into pre- and post-FTA cohorts and tested for differences in earnings between the two. The result was a positive (and significant) coefficient on the pre-FTA cohort when the post-FTA cohort was used as the control group. Thus, the improvement in the 1986-90 cohort is the result of an improvement in those who entered before the FTA came into effect. Further, employment authorizations granted under the FTA are non-immigrant or non-permanent resident visas.[31] No persons born in the United States claimed to have this immigration status in our subsample of the 1991 Canadian census. Thus, the FTA appears not to have had any impact on any improvement in the most recent male immigrant cohort, at least in these data.

Conclusions

This paper has addressed changes in the earnings abilities of North American immigrants in the Canadian and U.S. labor markets over the decade of the 1980s. Given the relative widening of the earnings distribution in the United States, we expected Canadians to be positively selected into the United States labor market, while Americans of lesser ability would select into the Canadian market. Theoretically, those at the upper end of the Canadian earnings distribution can improve their earnings by migrating to the United States. Conversely, Americans of more limited ability may enter Canada, since their labor market deficiencies will not be penalized as heavily north of the border.

For Canadians in the United States our results are generally supportive of the theory. The earnings of Americans in Canada do not support our expectations. American males entering Canada continue to be at an earnings disadvantage, but this disadvantage is narrowing over successive cohorts relative to Canadian males. For female immigrants in Canada, there is no evidence of changes in cohort quality over time.

Appealing to changes in Canadian immigration policy over this period offers no assistance in explaining this earnings improvement of Americans in Canada. Borjas (1993) says that the points system in Canada plays a much weaker role than would be presumed. The skills filters built into the Canadian system, along with the absence of such filters in the United States, imply that Americans in the Canadian labor market should be quite successful, whereas Canadian immigrants in the United States would perform relatively less well. The results presented above show that there has been an improvement in the earnings ability of Americans migrating to Canada. Why this occurred only for entrants who entered in the latter half of the decade is not clear.

The migration of highly skilled Canadians to the United States has historically been a matter of concern in Canada. These results support this concern. They also suggest that this trend is likely to be intensified before it is reversed. Cutbacks in government expenditures on medicare, education, and research and development are blamed for the exodus of physicians, educators, and scientists. Furthermore, changes in U.S. immigration policy as a result of the FTA and the recent increase in the number of available professional visas will simply act to hasten the departure of Canada's best and brightest.

NOTES

This is a revised version of a paper presented at the 14th biennial meeting of the Association for Canadian Studies in the United States, Minneapolis, 19-23 November 1997. Dan Hamermesh, Gary Hunt, and two anonymous referees provided useful comments. Financial support from the Canadian-American Center, University of Maine, is gratefully acknowledged.

1. Chiswick and Miller (1988) also find a decline in the quality of immigrants to Canada from the United States in the 1970s.

2. "Observable" characteristics are the attributes of individuals that can be quantified. These include variables such as education, experience, and marital status, which commonly appear in data sets. "Unobservable" characteristics are those that cannot be quantified and are a result of qualitative differences between individuals. These characteristics are often referred to as individual talent or, in the present context, as the "quality" of immigrants. Both observable and unobservable characteristics add to the earnings potential of an individual.

3. Richardson notes that this gap between the two countries narrowed between 1989 and 1992. This, however, will not affect the main results of this paper.

4. See Borjas (1994) for a recent review of this literature.

5. Bloom and Gunderson (1991) note the limitations of these data. First, they cannot be used to estimate individual-specific effects, since individuals from either cross-section are not linked. Second, a particular entry cohort observed in the first census year may not be representative of the corresponding cohort in the following decennial census. This could be due to non-random patterns of out-migration or mortality, for example. Third, two census cross-sections only allow for estimation of assimilation and cohort effects, but not period effects. This final point is discussed below.

6. We did estimate separate models for both immigrants and natives and found that coefficient estimates on demographic variables were not dissimilar. Furthermore, substantive results did not vary markedly when we did not constrain these coefficients to be equal for immigrants and natives.

7. We realize that constraining the *YSM* coefficient to be equal for each entry cohort is a limiting assumption, as it measures the average assimilation rates of different entry cohorts. Baker and Benjamin (1994) find that the assimilation effects can vary between immigrant cohorts. They look, however, at immigrants to Canada from all source regions over time. Thus, assimilation rates are more likely to differ in their sample than in our sample, since we address only one source country. Furthermore, given the small number of Canadians in our 1980 U.S. census sample, it was decided that the efficiency gains justified this restriction.

8. In each case, a subsample of natives is chosen to increase computational efficiency. Since these are random subsamples, however, it is highly unlikely that the results will be affected.

9. Individuals in the YPRE60/YPRE61 cohort were arbitrarily assigned an entry date of 1950.

10. Cohen, Zach, and Chiswick (1997) find similar increases in educational attainment amongst immigrants to the United States. Their data, however, include Canadians as part of a group that includes migrants from Europe, Australia, and New Zealand.

11. The Y86/Y85 cohort is eliminated from the regressions. Full results can be found in Appendix Tables A1 and A2. It should be noted that these results are robust to the

use of various definitions of the dependent earnings variable. In separate estimates we used a number of different log specifications for the earnings variable, including annual wages, weekly earnings, and hourly wages (available only in the U.S. data) and the entry, assimilation, and cohort effects did not change markedly. Similar results were also obtained when we limited the sample to include only those who worked for the full year (more than thirty-nine weeks). Adding industry and occupation controls did not significantly alter the results (these are presented in Appendix Tables A1 and A2).

12. In terms of the notation in equation (1), this is simply equal to $-\alpha/\eta$.

13. It should be noted that the years-to-equality calculations may be somewhat uncertain for the most recent cohorts, since the calculations are based on only a limited number of years of experience in the foreign labor market. See Bloom, Grenier, and Gunderson (1995) for discussion.

14. An F-test on the joint significance of the entry cohort coefficients yielded a value of 1.31.

15. Here we test the null hypothesis that $\alpha + \delta_j = 0$.

16. For each cohort, this number is calculated by adding the entry effect to the cohort effect, and then dividing by the assimilation effect (again, in terms of equation (1) this is equal to $-(\alpha + \delta_j) / \eta$). Since the most recent entry cohort (Y86) is eliminated from the estimated equation, years to equality is equal to $-\alpha / \eta$.

17. In particular, we test the null hypothesis that $-(\alpha + \delta_j) / \eta = 0$.

18. McDonald and Worswick (1998) show that the unemployment rate is an important determinant of the rate of earnings assimilation experienced by different Canadian immigrant cohorts.

19. An F-test for joint significance of these coefficients yielded a value of 6.23, significant at 5 percent.

20. Again, we test the null hypothesis that the entry effect and the coefficient effect are jointly equal to zero ($\alpha + \delta_j = 0$). For the Y85 cohort, this collapses to $\alpha = 0$. Only in the cases of the Y80 and Y85 cohorts are these differences significantly different from zero at one percent. All others are insignificant at 5 percent.

21. Both Beach and Worswick (1993) and Bloom, Grenier, and Gunderson (1995) generally find similar results when comparing female to male immigrants in their Canadian data. The latter study notes that this suggests that migration selectivity is stronger for men than for women, not surprising since women are more likely than men to be tied movers.

22. An F-test for the joint significance of the cohort coefficients has a value of 3.40, significant only at the 10-percent level.

23. De Silva compared independent immigrants (those admitted based on labor market criteria) with assisted relatives (more distant relatives of permanent Canadian residents) and refugees. Unfortunately, his data did not permit him to analyze the family class immigrants (immediate family members). In 1982, the depth of the recession in Canada, 54.1 percent of all American women entered as family class immigrants, compared to 43.4 percent who entered as independents. These numbers are discussed below.

24. In all instances the null hypothesis that the entry effect and the cohort effect are jointly equal to zero ($\alpha + \delta_j = 0$) can be rejected at at least the 5-percent level of significance. We can also reject the hypothesis that the years-to-equality estimates are zero at the 5-percent level for all cohorts with the exception of the most recent.

25. The F-statistic in this case had a value of 1.95.

26. Only in the cases of entrants in the 1980s can we reject the null that entry wages are equal to those of comparable American-born women.

27. See Fagnan (1995) and Green and Green (1995) for details of the changes in Canadian immigration policy.

28. These immigrants were classified as independent immigrants.

29. Coulson and DeVoretz (1993), Green and Green (1995), and Bloom, Grenier, and Gunderson (1995), among others, discuss declining immigrant skills quality as a result of the switch in the emphasis of immigration policy towards family reunification.

30. Based on author's calculations using data from Immigration Canada (1992) and Employment and Immigration Canada (1984). It should be noted that these are based on country of last permanent residence, not country of birth, since these are the only data available. The correlation between the two, however, is high in the case of those entering from the United States.

31. Recent work by DeVoretz and Laryea (1998), for example, notes that the temporary emigration under the employment visas granted under the FTA has become a back door to permanent emigration into the United States. They also note the large increase in net flows from Canada to the United States that has taken place beginning in 1990. Obviously, our data would fail to capture these individuals, and a similar analysis of recent Canadian emigration to the United States awaits the 2000 United States census.

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APPENDIX

TABLE A1: OLS Estimates of Log Earnings Equations, Males (absolute values of t-statistics are in parantheses)

	Canada				United States			
	w/o controls			w/controls	w/o controls			w/controls
	1981	1991	pooled	pooled	1981	1991	pooled	pooled
YEARSED	0.0589	0.0684	0.0647	0.0531	0.0666	0.0956	0.0768	0.0613
	(25.790)	(32.792)	(41.595)	(30.332)	(64.287)	(67.279)	(91.302)	(60.770)
EXP	0.0312	0.0301	0.0310	0.0274	0.0344	0.0335	0.0346	0.0315
	(14.066)	(15.100)	(20.719)	(18.875)	(32.731)	(24.377)	(41.177)	(38.920)
EXPSQ	-0.0005	-0.0004	-0.0005	-0.0004	-0.0005	-0.0005	-0.0005	-0.0005
	(12.009)	(11.054)	(16.233)	(14.752)	(24.913)	(17.121)	(30.405)	(29.062)
MARRIED	0.2012	0.2073	0.2064	0.1652	0.2191	0.2223	0.2231	0.1847
	(12.904)	(16.833)	(21.296)	(17.632)	(29.468)	(26.971)	(40.238)	(34.480)
URBAN	0.0685	0.0935	0.0843	0.0932	0.1311	0.1841	0.1521	0.1396
	(5.648)	(8.993)	(10.592)	(11.866)	(18.910)	(20.120)	(27.317)	(25.788)
WKSWK	0.0300	0.0283	0.0289	0.0288	0.0312	0.0341	0.0327	0.0314
	(51.907)	(61.314)	(79.847)	(79.680)	(89.298)	(88.332)	(125.786)	(123.939)
PTWORK	-0.4582	-0.5646	-0.5239	-0.4680	-0.3879	-0.7859	-0.5548	-0.4980
	(14.936)	(20.980)	(25.674)	(23.732)	(21.807)	(35.981)	(39.963)	(37.253)
IMMIG	-0.1789	-0.1149	-0.0826	-0.0429	0.2275	0.2314	0.1960	0.1672
	(5.123)	(3.160)	(1.442)	(0.778)	(1.882)	(14.242)	(7.551)	(6.703)
YSM	0.0076	0.0036	0.0059	0.0046	-0.0082	-0.0033	0.0101	0.0070
	(4.912)	(2.426)	(2.186)	(1.752)	(1.636)	(6.132)	(2.076)	(1.505)
Y81/Y80			-0.0973	-0.0739			-0.0654	-0.0559
			(1.263)	(0.999)			(1.598)	(1.424)

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TABLE A1: continued

	Canada				United States			
	w/o controls			w/controls	w/o controls			w/controls
	1981	1991	pooled	pooled	1981	1991	pooled	pooled
Y76/Y75			-0.1055	-0.1054			-0.0824	-0.0531
			(1.506)	(1.560)			(1.426)	(0.959)
Y71/Y70			-0.0541	-0.0544			-0.1809	-0.1196
			(0.783)	(0.812)			(2.258)	(1.556)
Y66/Y65			-0.0876	-0.0853			-0.2415	-0.1710
			(1.153)	(1.155)			(2.387)	(1.762)
Y61/Y60			-0.0656	-0.0602			-0.3263	-0.2286
			(0.717)	(0.677)			(2.640)	(1.928)
YPRE61/YPRE60			-0.0831	-0.0758			-0.4982	-0.3584
			(0.789)	(0.737)			(2.754)	(2.065)
CENS91			0.4930	0.4195			0.4369	0.4584
			(57.081)	(4.303)			(86.906)	(94.275)
CONSTANT	7.0988	7.5219	7.0552	7.1205	6.5858	6.4618	6.3584	6.6305
	(159.053)	(197.882)	(241.425)	(71.846)	(273.782)	(218.506)	(340.452)	(248.603)
R^2	0.4475	0.4149	0.4889	0.5315	0.3558	0.4489	0.4676	0.5107
N	7263	12819	20082	20082	34919	26693	61612	61612

Note: Control variables include occupation and industry.

Table A2: OLS Estimates of Log Earnings Equations, Females
(absolute values of t-statistics are in parantheses)

	Canada				United States			
	w/o controls			w/controls	w/o controls			w/controls
	1981	1991	pooled	pooled	1981	1991	pooled	pooled
YEARSED	0.0849	0.0975	0.0932	0.0570	0.0766	0.1133	0.0927	0.0606
	(24.114)	(35.401)	(42.660)	(22.968)	(57.099)	(68.244)	(87.740)	(49.242)
EXP	0.0105	0.0176	0.0152	0.0105	0.0063	0.0100	0.0088	0.0084
	(3.589)	(7.671)	(8.332)	(6.009)	(5.415)	(7.214)	(9.800)	(9.767)
EXPSQ	-0.0001	-0.0003	-0.0002	-0.0001	-0.0001	-0.0001	-0.0001	-0.0001
	(2.370)	(5.813)	(6.093)	(4.062)	(2.222)	(4.334)	(5.451)	(5.576)
MARRIED	-0.0256	-0.0198	-0.0220	-0.0329	-0.0675	-0.0422	-0.0542	-0.0633
	(1.391)	(1.456)	(2.007)	(3.141)	(9.696)	(5.448)	(10.378)	(12.662)
URBAN	0.0925	0.1458	0.1289	0.1114	0.1348	0.2041	0.1654	0.1565
	(5.884)	(12.377)	(13.599)	(12.099)	(17.794)	(21.553)	(27.555)	(27.105)
WKSWK	0.0303	0.0303	0.0304	0.0294	0.0312	0.0344	0.0328	0.0320
	(51.129)	(67.220)	(84.040)	(83.975)	(113.387)	(110.501)	(157.769)	(159.564)
PTWORK	-0.5654	-0.6022	-0.5903	-0.5475	-0.6577	-0.8054	-0.7325	-0.6667
	(30.986)	(43.507)	(53.201)	(50.335)	(74.357)	(79.494)	(108.765)	(101.769)
IMMIG	-0.1178	-0.0947	-0.0401	-0.0174	0.1878	0.1271	0.0995	0.0793
	(2.774)	(2.696)	(0.701)	(0.319)	(1.486)	(7.553)	(3.600)	(3.010)
YSM	0.0028	0.0022	0.0060	0.0046	-0.0067	-0.0010	0.0066	0.0053
	(1.478)	(1.484)	(1.908)	(1.536)	(1.345)	(1.758)	(1.389)	(1.170)
Y81/Y80			-0.1749	-0.1578			0.0205	0.0299
			(2.383)	(2.259)			(0.487)	(0.748)

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Table A2: continued

	Canada				United States			
	w/o controls			w/controls	w/o controls			w/controls
	1981	1991	pooled	pooled	1981	1991	pooled	pooled
Y76/Y75			-0.1153	-0.0994			-0.0734	-0.0699
			(1.655)	(1.498)			(1.274)	(1.272)
Y71/Y70			-0.0860	-0.0623			-0.1428	-0.1130
			(1.170)	(0.887)			(1.815)	(1.507)
Y66/Y65			-0.1087	-0.0847			-0.1258	-0.0865
			(1.304)	(1.064)			(1.269)	(0.915)
Y61/Y60			-0.1629	-0.1087			-0.1620	-0.1221
			(1.640)	(1.146)			(1.331)	(1.053)
YPRE61/YPRE60			-0.2027	-0.1702			-0.2856	-0.2219
			(1.684)	(1.477)			(1.613)	(1.315)
CENS91			0.5083	0.4599			0.4911	0.5034
			(46.600)	(5.817)			(91.186)	(97.614)
CONSTANT	6.7024	6.9607	6.5378	7.0292	6.4637	6.2485	6.1419	6.7322
	(106.916)	(141.886)	(168.785)	(78.284)	(243.180)	(193.572)	(298.716)	(171.535)
R^2	0.5396	0.4968	0.5629	0.6065	0.5106	0.5546	0.5951	0.6327
N	4928	11182	16110	16110	26167	24914	51081	51081

Note: Control variables include occupation and industry.