A COMPARATIVE ANALYSIS OF THE LABOR MARKET IMPACT OF INTERNATIONAL MIGRATION: CANADA, MEXICO, AND THE UNITED STATES

Abdurrahman Aydemir and George J. Borjas
Statistics Canada and Harvard University

November 2005
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ABSTRACT

The North American experience with international migration stands in unique contrast to much of the rest of the world. This paper uses microdata drawn from the national censuses of Canada, the United States, and Mexico, and applies the same methodological framework to these data to examine the impact of international migration on the labor market. We find a numerically comparable and statistically significant inverse relation between immigrant-induced shifts in labor supply and wages in each of the three countries: A 10 percent labor supply shift is associated with about a 3 to 4 percent opposite-signed change in wages. Despite the similarity in the wage elasticity, the impact of international migration on the wage structure differs significantly across countries. In Canada, international migration substantially narrowed wage inequality because immigrants in Canada tend to be disproportionately high-skill. In the United States, international migration substantially increased wage inequality because immigrants in the United States tend to be disproportionately low-skill. In Mexico, however, emigration rates are highest in the middle of the skill distribution and lowest at the extremes. As a result, international migration greatly increased relative wages in the middle of the Mexican skill distribution and lowered relative wages at the extremes. Paradoxically, the large-scale migration of workers from Mexico may have reduced slightly the relative wage of the low-skill workers remaining in that country.
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I. Introduction

There has been a resurgence of large-scale international migration throughout much of the world in recent decades. Nearly 3 percent of the world’s population now lives in a country where they were not born (United Nations, 2002). These population flows have rekindled the debate over the social and economic consequences of international migration in many countries, and motivated many economists to develop and estimate models designed to measure this impact.

The North American experience with international migration—particularly in Canada, the United States, and Mexico—stands in unique contrast to much of the rest of the world. First, Canada has long had a policy of importing workers to augment its workforce. As a result, the foreign-born share in the Canadian population is higher than in most other developed countries. Second, the United States receives the largest immigrant influx (in absolute size) of any country in the world. Finally, the emigration of Mexicans, almost exclusively to the United States, has drained the Mexican economy of a large fraction of its workers in a relatively short time.

The textbook model of a competitive labor market has clear and unambiguous implications how wages and employment opportunities in a particular country should adjust to these labor supply shocks, at least in the short run. In particular, immigration should lower the wage of competing workers and increase the wage of complementary workers, of workers whose skills become more valuable because of immigration.

Despite the common-sense intuition behind these theoretical predictions, the economics literature has—at least until recently—found it very difficult to document the inverse relation between immigrant-induced supply shocks and wages. Because immigrants typically cluster into a small number of geographic areas in most receiving countries, much of the literature attempts

* Dr. Aydemir is a Senior Economist at Statistics Canada; Dr. Borjas is a Professor of Economics and Social Policy at the Kennedy School of Government, Harvard University, and a Research Associate at the National Bureau of Economic Research. We are grateful to Daniel Hamermesh, Richard Freeman, Lawrence Katz, and Steve Trejo for helpful comments on a previous draft. This paper represents the views of the authors and does not necessarily reflect the opinion of Statistics Canada.
to estimate the labor market impact of immigration in a receiving country by comparing economic conditions across local labor markets in that country. In particular, many studies estimate the correlation between measures of immigrant penetration in local labor markets and measures of economic outcomes, such as native wages or employment rates (see, for example, Grossman, 1982, Borjas, 1987, Altonji and Card, 1991, and Card, 1991, 2001, for the United States; Roy, 1997, for Canada). The sign of this “spatial correlation” has often been interpreted as indicating the direction in which immigrant supply shocks affect native wages; a negative correlation, for instance, would be interpreted as indicating that immigrant-induced increases in labor supply lower wages in the labor market. Although there is a lot of dispersion across studies, there is a strong central tendency towards finding a near-zero spatial correlation. This weak correlation has been interpreted as indicating that immigrants have little impact on the wage structure of the receiving country (see the surveys by Friedberg and Hunt, 1995, and Smith and Edmonston, 1997).

Borjas, Freeman, and Katz (1997) challenged this conventional wisdom by arguing that the spatial correlation—the correlation between labor market outcomes and immigration across local labor markets—may not truly capture the economic impact of immigration if native workers respond by moving their labor or capital to localities seemingly less affected by the immigrant supply shock. Borjas (2003) used the insight that the labor market impact of immigration may be measurable only at the national level to examine how the wages of U.S. workers in particular skill groups—defined in terms of both educational attainment and years of work experience—were related to the immigrant supply shocks affecting those groups. The national labor market evidence indicated that wage growth was strongly and inversely related to immigrant-induced supply increases. This evidence, based on the study of wage trends in the national labor market over a four-decade period, is consistent with the implications of the textbook model of a competitive market.

1 In Canada, the 2001 Census indicates that 64 percent of immigrant men resided in three cities (Toronto, Vancouver, and Montreal), as compared with only 26.3 percent of their native-born counterparts. Similarly, the 2000 U.S. Census indicates that 34.1 percent of immigrant men lived in three metropolitan areas (Los Angeles, New York, and Chicago), while only 10.4 percent of natives clustered in these three localities.

2 Devoretz and Laryea (1997) also study the labor market impact of immigration in the Canadian labor market but use immigrant concentration by industry, rather than metropolitan area, to measure the effect.
This paper examines if the Borjas (2003) methodological framework provides a useful point of departure for investigating the labor market impact of immigration beyond the U.S. context. Our analysis uses the same methodological framework and sample construction to analyze the impact of international migration in Canada, Mexico, and the United States. The joint application of the framework to these three countries is interesting for a number of reasons. First, although both Canada and the United States admit large numbers of immigrants, they pursue very different immigration policies that create large differences in the skill composition of the immigrant population in each country. In the United States, immigration tends to disproportionately increase the number of low-skill workers. Canadian immigration, in contrast, tends to disproportionately increase the number of high-skill workers. In short, different groups of native workers are likely to be affected by the immigrant supply shocks in Canada and the United States. A comparative analysis of the labor market impact of different immigration policies can increase our understanding of how the pursuit of specific policies alters the distributional and efficiency impacts of international migration.

Second, the comparison of Mexico with the other two countries should provide a mirror-image of the economic impact of international migration flows. Between 1980 and 2000, for example, both the United States and Canada experienced substantial immigrant inflows, increasing the number of working men by an average of 10.4 percent in Canada and 10.9 percent in the United States. In contrast, Mexico experienced a 14.9 percent decrease in the size of its potential workforce. Two recent studies, Mishra (2003) and Hanson (2005), examine the impact of emigration on Mexican wages. Both studies report a significant positive correlation between

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3 After the initial draft of this paper was completed, we learned of Bonin’s (2005) application of the Borjas framework to the German labor market, and of the reappraisal by Bohn and Sanders (2005) of the published Borjas evidence. Bonin finds that supply shocks in Germany lower wages in the national German labor market, but by less than in the Borjas study. However, the German census does not provide unambiguous information on immigration status, so that the data cannot distinguish between foreign citizens born abroad and foreign citizens born in Germany. As a result, it is unclear exactly what type of supply shock is being measured by the immigration variable. The last section of the Bohn and Sanders study reports an attempt to apply the framework to the Canadian context. The Bohn-Sanders analysis of the Canadian data uses some of the much smaller publicly released Public Use Microdata Files, and obtains results that are weaker than those reported below that are based on the entire Census files maintained by Statistics Canada. In a companion paper (Aydemir and Borjas, 2005), we show that the sampling error associated with calculating the immigrant supply shock in relatively small samples leads to substantial attenuation bias in the measurement of the wage impact of immigration.

4 The Mishra study is closely related to the descriptive regression analysis presented in Section V because the Mishra study is itself designed to be an application of the Borjas methodological framework to the Mexican
Mexican wages and emigration. Our analysis confirms the existence of this basic correlation and extends the existing work in two distinct ways. First, we show that the Mexican data—like the corresponding data for Canada and the United States—can be fruitfully analyzed using a structural model of factor demand that leads to roughly similar estimates (across receiving countries) of the relevant elasticities of substitution. Second, we show that despite this similarity, the link between the Mexican wage structure and emigration is not at all what one expects—namely, an increase in the relative wage of the low-skill workforce that is the source of most Mexican immigrants in the United States. Because Mexican emigration rates are relatively lower both for workers at the bottom and the top of the skill distribution, international migration has not increased (and may, in fact, have lowered) the relative wage of low-skill workers.5

Third, because our study analyzes similar data across the three countries (drawn from microdata samples for each country’s national census) and imposes the same theoretical structure on these data, the paper reports the results of a relatively rare methodological experiment. In short, we attempt to determine if the simple theoretical insights implied by the laws of supply and demand lead to relatively similar qualitative (and quantitative) responses in the labor markets, despite the very different nature of the labor supply shocks, institutions, and economic conditions in the three countries.

The key finding of our study is that there is a numerically sizable and statistically significant inverse relation between labor supply shocks and wages in each of the countries under study. Even though the average wage response of international migration in each of the countries is relatively similar—a 10 percent labor supply shift is associated with a 3 to 4 percent opposite-signed change in wages—the impact of international migration plays a drastically different role in the evolution of each country’s wage structure. In Canada, international migration substantially narrowed wage inequality. In the United States, international migration substantially increased wage inequality. In Mexico, however, international migration greatly increased the relative wages of workers in the middle of the skill distribution, but lowered the relative wages of workers at the extreme of the skill distribution. Paradoxically, despite the large-

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5 See Chiquiar and Hanson (2005) and Ibarraran and Lubotsky (2005) for studies of the structure of emigration rates across Mexican skill groups.
scale migration of low-skill workers from Mexico to the United States, the relative wage of low-skill workers remaining in Mexico may have fallen slightly, further increasing the incentives of low-skill Mexicans to continue their migration to the United States.

II. International Differences in Immigration Policies

Before proceeding to the empirical analysis, it is instructive to briefly describe the large differences in immigration policies in Canada and the United States.

The major impetus for the resurgence of large-scale immigration to the United States in recent decades, particularly immigration from less developed countries, came from the 1965 Amendments to the Immigration and Nationality Act. Before 1965, immigration to the United States was guided by the national-origins quota system. This scheme restricted the number of immigrants, used quotas to allocate visas across countries, and partly used skills to allocate visas among applicants from a given country. The number of visas given to each country was based on the ethnic composition of the U.S. population in 1920. As a result, 60 percent of all available visas were awarded to applicants from only two countries, Germany and the United Kingdom.

The 1965 Amendments repealed the national-origins quota system. Along with subsequent minor legislation, the Amendments set a higher worldwide numerical limit for immigration and enshrined a new objective, the reunification of families, for allocating entry visas among the many applicants.

There has also been a substantial increase in illegal immigration since 1965. It is estimated that 10.3 million illegal aliens resided in the United States in March 2005, with 5.9 million (57 percent of the illegal population) being of Mexican origin (Passel, 2005). Further, the size of the illegal population has been growing very rapidly in recent years, by around 700 thousand illegal immigrants annually since 2000.

The increasing importance of family preferences in the awarding of entry visas, combined with the increasing number of unscreened (and low-skill) illegal aliens, resulted in a substantial shift in the skill composition of the foreign-born workforce: A large and growing group of low-skill immigrants dominates the immigrant influx into the United States. The social and economic impacts of this skewing of the immigrant skill distribution are, in fact, at the core of the issues that dominate the American immigration debate.
As in the United States, Canadian immigration policy until the early 1960s was also based on a national-origin preference system that limited immigration of certain national origin groups while facilitating the entry of others. Canada moved away from the national origin allocation scheme in 1962 and replaced it with a system that emphasized the skills of visa applicants in awarding entry visas. In 1967 Canada introduced the point system that aimed explicitly at selecting immigrants with highly sought skills. The point system awards points to visa applicants who have particular socioeconomic characteristics (e.g., more schooling and fluent English or French language skills), and then sets a “passing grade” that determines the subset of visa applicants who qualify for a visa.

For the most part, immigrants who entered Canada between 1970 and 2000 were admitted under one of four major categories: 1) family class migrants, covering immediate family members; 2) nominated relatives, covering close relatives; 3) independent migrants, covering various subcategories of skilled workers as well as entrepreneurs and investors; and (4) refugees. Individuals applying under the categories of nominated relatives or independent migrants are subject to the points test.

Another important difference between the United States and Canada is the latter’s explicit tie of the annual level of immigration to the macroeconomic environment, increasing the level of immigration during economic booms and reducing it during recessions. These adjustments often were accomplished by lowering the number of immigrants admitted under the independent migrants class—the Canadian authorities, after all, have considerably less flexibility in adjusting the number of immigrants under the family and refugee classes. However, the tap-on, tap-off policy was abandoned in the early 1990s, and there have been relatively high immigration levels since. Because of these adjustments, the share of immigrants belonging to each of the categories has fluctuated significantly over time. In particular, the share of independent immigrants rose from 21 percent in 1984 to 59 percent in 2000 (Citizenship and Immigration Canada, 2001), and over three quarters of those admitted in the independent migrant category were skilled workers.

The third country in our study, Mexico, is not a major immigrant-receiving country, but is a major sending country. The 2000 Mexican Census asked families to name the location of any relatives who had migrated abroad between 1995 and 2000, and 97 percent of these families reported the United States as the relatives’ destination country (Caponi, 2004). The 2000 U.S. Census enumerated 9.2 million Mexican-born persons (or 29.5 percent of all foreign-born
persons in the United States). In 2000, the Mexican population stood at 100.3 million, suggesting an emigration rate of 8.4 percent. In short, a significant fraction of the Mexican population has left that country for the United States.

There are no legal restrictions on which persons can leave Mexico—except for those imposed by U.S. immigration policy and border patrol enforcement. Mexican immigrants have constituted the largest component of legal immigration in the United States in recent decades. In the 1990s, for instance, the United States admitted 9.1 million legal immigrants, and 24.8 percent of those immigrants originated in Mexico (U.S. Immigration and Naturalization Service, 2003). As noted above, Mexicans also make up a disproportionately large share of the illegal immigrant population in the United States.

Recent studies by Chiquiar and Hanson (2005) and Ibarraran and Lubotsky (2005) examine how the skill composition of Mexican immigrants in the United States compares to that of Mexicans who choose to remain in Mexico. The vast majority of Mexican emigrants are high school dropouts (63 percent of Mexican working men enumerated in the 2000 U.S. Census are high school dropouts). Chiquiar and Hanson (2005) point out, however, that the emigration rate is higher for Mexicans who are high school graduates than for Mexican high school dropouts.6

The changes in immigration policies in Canada and the United States—and the increasingly powerful pull of the U.S. labor market to large segments of the Mexican population—have important implications for the overall trends in international migration in the three countries. Figure 1 shows the trend in the immigrant or emigrant share in the male workforce (i.e., the fraction of the workforce aged 18-64 that is foreign-born in Canada or the United States, or the fraction of the workforce that emigrated from Mexico) for each of the three countries under study.7 The immigrant share in Canada has been relatively stable over the past 30 years, hovering at around 20 percent (except for a slight dip in the 1980s). The immigrant share in the United States declined from 1960 to 1970, but has increased significantly since. In 1970, less than 5 percent of the male workforce was foreign-born; by 2000, the immigrant share had increased to almost 15 percent. Finally, the emigrant share in Mexico increased dramatically in

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6 An important caveat in this conclusion, emphasized by Caponi (2004), is that the rate of emigration is highest for Mexicans who never attended school. This is not a trivial point because 10 percent of the Mexican male workforce in 2000 did not have any formal education.

7 The data used to calculate these statistics is discussed in detail in the next section.
the past three decades. As recently as 1970, the emigration rate of male workers was under 3 percent. By 2000, the emigration rate had risen to 16 percent.

III. Methodological Framework and Data

As noted in the introduction, the potential problems associated with exploiting wage differences across local labor markets to identify the labor market impact of immigration are now well understood. Natives (and pre-existing immigrants) may respond to the adverse wage impact of immigration by moving their labor or capital to other cities (Filer, 1992; Frey, 1995; Card, 2001; Borjas, 2006). These regional flows diffuse the impact of immigration across all regions, suggesting that the labor market impact of immigration cannot be measured by comparing economic conditions across localities and may only be measurable at a national level.8

In a recent study, Borjas (2003) proposed a new methodology that allows direct estimation of the labor market impact of immigration at the national level. We use this framework to estimate and compare the impact of international migration in three national labor markets: Canada, Mexico, and the United States. As in Borjas (2003), we define the skill groups that make up a national labor market in terms of both education and labor market experience. This definition, of course, implicitly assumes that workers with the same level of schooling but with different levels of experience are imperfect substitutes in production (Welch, 1979; Card and Lemieux, 2001). We then use the time-variation in the share of immigrants within each skill group to identify the impact of immigration on the wage structure.

The data used in this analysis comes from microdata Census files from Canada, the United States, and Mexico. Our study of the Canadian labor market uses all available microdata files from the Canadian Census (1971, 1981, 1986, 1991, 1996, and 2001). Each of these files represents a 20 percent sample of the Canadian population (except for the 1971 file, which represents a 33.3 percent sample). In the U.S. context, we use the 1960, 1970, 1980, 1990 and 2000 Integrated Public Use Microdata Sample (IPUMS) of each decennial Census. The 1960 file represents a 1 percent sample of the U.S. population, the 1970 file represents a 3 percent sample, and the 1980 through 2000 files represent a 5 percent sample. Finally, our study of the Mexican

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8 In addition, immigrants may not be randomly distributed across localities. If immigrants cluster in cities with thriving economies (and high wages), there would be a built-in positive correlation between immigration and wages.
labor market uses the 1960, 1970, 1990 and 2000 IPUMS of the Mexican censuses. The 1960 file represents a 1.5 percent sample of the Mexican population, the 1970 file represents a 1 percent sample; the 1990 file represents a 10 percent sample; and the 2000 file represents a 10.6 percent sample. Throughout most of the paper, the empirical analysis is restricted to men aged 18 to 64 who participate in the civilian labor force. The data appendix describes the construction of the sample extracts for each of the three countries in detail.

Our study of the U.S. data uses the convention of defining an immigrant as someone who is either a noncitizen or a naturalized U.S. citizen. In the Canadian context, we define an immigrant as someone who reports being a “landed immigrant” (i.e., a person who has been granted the right to live in Canada permanently by immigration authorities), and is either a noncitizen or a naturalized Canadian citizen. Finally, the Mexican census does not provide a count of the number of emigrants. As noted above, almost all Mexican emigrants had chosen the United States as their destination, suggesting that a valid count of the number of Mexican emigrants can be obtained from the U.S. Census (abstracting from the undercount problem in the U.S. Census). Hence we define the number of emigrants from Mexico at a particular point in time as the number of Mexican-born persons enumerated by the corresponding U.S. Census.

As in Borjas (2003), skill groups are defined in terms of both educational attainment and years of labor market experience. To analyze the Canadian trends, we classify workers into five distinct education groups using the detailed information on the type of degree available in the Canadian census. The five education groups are: (1) workers who are high school dropouts; (2) workers who have either a high school or a vocational degree; (3) workers who have a high school and vocational degree or have a post-secondary certificate or diploma below the bachelor’s degree; (4) workers who have a bachelor’s degree; and (5) workers who have a post-graduate degree.

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9 There is no microdata sample for the 1980 Census because the primary documents were destroyed by an earthquake.

10 Since 1991, the Canadian Censuses include non-permanent residents. This group includes those residing in Canada on an employment authorization, a student authorization, or a Minister’s permit, or who were refugee claimants at the time of Census (and family members living with them). Non-permanent residents accounted for 0.7, 0.4 and 0.5 percent of the samples used in this study in 1991, 1996 and 2001, respectively, and are included in the immigrant counts for those years.

11 Our grouping of education groups in the Canadian context was partly suggested by a Mincerian earnings regression estimated in the pooled 1971-2001 censuses that included controls for detailed type of educational
Our analysis of the U.S. data uses five education groups that roughly correspond to the Canadian categories: (1) high school dropouts (workers who have less than 12 years of completed schooling); (2) high school graduates (workers who have exactly twelve years of schooling); (3) workers who have some college (thirteen to fifteen years of schooling); (4) college graduates (workers who have sixteen years of schooling); and (5) workers with postgraduate education (workers who have more than 16 years of schooling).12

The Mexican schooling system (and educational distribution) differs greatly from that of Canada and the United States. Students get a primary-level degree after 6 years of education, a lower-secondary degree after 9 years, and a secondary level degree after 12 years. There are therefore, sizable spikes in the educational distribution at these termination points. It is worth noting that the bulk (94.9 percent in 1970 and 73.3 percent in 2000) of Mexican working men are high school dropouts (i.e., did not have a secondary level degree) in the period under study. We use five education groups in Mexico to capture these characteristics of the Mexican system: (1) high school dropouts with 0 to 8 years of schooling (i.e., workers whose completed schooling is less than a lower secondary education); (2) high school dropouts with 9 to 11 years of schooling (i.e., workers who have completed a lower secondary education); (3) high school graduates (i.e., workers who have completed a secondary education, either in a general or technical track); (4) workers with some college (i.e., workers who report completing a general track secondary education and also attending some college, or workers who completed a technical track secondary education and obtained a post-secondary technical degree); and (5) college graduates (i.e., workers who report completing a university education).13 Finally, because we rely on the attainment (and a quadratic in work experience). We found that workers with either a high school or a vocational degree have similar earnings, while those who hold both degrees have significantly higher earnings. Further, the earnings of this latter group were similar to the earnings of workers with a post-secondary certificate or a diploma below bachelor’s level. Finally, we found a sizable earnings difference between workers who have a post-graduate degree and workers with only a bachelor’s degree.

12 In Canada “college” typically refers to 2-year post-secondary institutions that grant a certificate or diploma below Bachelor’s level. Throughout the text, however, we use the term “college graduate” to refer to Bachelor’s degree holders in all three countries that we study.

13 In 2000, 45.0 percent of male workers in Mexico were high school dropouts with 0 to 8 years of schooling, 28.4 percent were high school dropouts with 9-11 years, 11.2 percent were high school graduates, 3.9 percent had some college, and 11.6 percent were college graduates.
U.S. Census to obtain a count of Mexican emigrants, we define five corresponding education groups in the U.S. data, based on their years of completed schooling.\(^\text{14}\)

We group workers into a particular years-of-experience cohort by using potential years of experience, roughly defined by Age – Years of Education – 6. The Canadian Census provides detailed information on the number of years that a worker attended grade school, post-secondary education below university, and university. By adding these variables, we can calculate the total years of schooling. The U.S. and Mexican Census data do not always report the number of years of school attended in such detail, so that we use the following approximation. We assume that age of entry into the labor market is 17 for high school dropouts, 19 for high school graduates, 21 for persons with some college, 23 for college graduates, 25 for workers with post-graduate degrees, and then calculate years of experience accordingly.\(^\text{15}\) The analysis is restricted to persons who have between 1 and 40 years of experience. Workers are aggregated into five-year experience groupings (i.e., 1 to 5 years of experience, 5 to 10 years, and so on) to capture the notion that workers who have roughly similar years of experience are more likely to affect each other’s labor market opportunities than workers who differ significantly in their work experience.

The skill cells corresponding to educational attainment \((i)\), experience level \((j)\), and calendar year \((t)\) define a skill group at a point in time for a given national labor market (for expositional convenience, we omit the index indicating the country). We define the immigrant supply shock by:

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(1)\quad p_{ijt} = \frac{M_{ijt}}{(M_{ijt} + N_{ijt})},
\]

where, in the case of Canada and the United States, \(M_{ijt}\) gives the total number of immigrants in the particular skill group; and \(N_{ijt}\) gives the total number of native workers in the national

\(^{14}\) In particular, the education groups used to classify Mexican emigrants are: 0-8 years of schooling, 9-11 years, 12 years, 13-15 years, and at least 16 years.

\(^{15}\) This approximation is probably much more appropriate for workers in the United States than in Mexico, particularly for the least educated workers. We experimented with alternative assumptions (e.g., high school dropouts enter the labor market at age 15) and the results are quite similar to those reported below.
economy. The variable $p_{ijt}$ then gives the immigrant share (i.e., the fraction of the workforce that is foreign-born). In the Mexican context, $M_{ijt}$ gives the number of workers belonging to a particular skill group who left Mexico (as enumerated by the U.S. Census), and $N_{ijt}$ represents the number of Mexicans who chose to remain in Mexico. The variable $p_{ijt}$ then gives the emigrant share (i.e., the fraction of workers in a particular skill group who have left the country).

**IV. Trends in International Migration by Skill**

It is useful to begin by illustrating the trends in immigration and emigration across education groups for each of the three countries. The top panel of Figure 2 reports the education-specific trends in the immigrant/emigrant shares. In Canada, the immigrant share among workers who have at most a high school degree was either relatively constant over the past three decades or declined slightly, and the immigrant share among workers with some college declined substantially. At the same time, however, there was a sizable increase in the immigrant share among workers who have at least a college degree. This shift towards a high-skill immigrant influx was precisely the goal of the point system. The immigrant share among college graduates rose from 21.4 to 26.7 percent between 1986 and 2001, while the immigrant share among workers with a post-graduate degree rose from 32.5 to 38.2 percent. In short, the changes in Canadian immigration policy favorably shifted the skill mix of the immigrant influx and substantially increased the supply of high skill workers.

The U.S. experience stands in striking contrast. Unlike in Canada, the immigrant share rose for all education groups. Although the immigrant share among highly educated workers (particularly among workers with a post-graduate education) increased, the middle panel of Figure 2 shows that the increase was much more rapid among workers who are high school dropouts. As recently as 1980, for instance, only 10.9 percent of the high school dropout workforce was foreign-born. By 2000, 40.9 percent of high school dropouts were foreign-born. In contrast, the immigrant share among college graduates was 7.3 percent in 1980 and rose to 13.4 percent in 2000, while the immigrant share among post-graduate degree holders was 9.0 percent in 1980 and rose to 17.1 percent in 2000.

The bottom panel of Figure 2 shows that the emigration rate for Mexican workers is larger for workers in the middle of the education distribution (i.e., for workers who are either high school graduates or have some college). As noted above, however, Mexico has relatively
few workers with this “middle” level of education. As a result, even though the emigration rate for high school dropouts is relatively low, it leads to a significant numerical outflow of low-skill workers. Moreover, as Figure 2 shows, the emigration rate of Mexican high school dropouts, particularly of the least-educated high school dropouts, rose rapidly during the 1990s, from 8.6 percent in 1990 to 16.0 percent in 2000. It is also worth noting that the most highly educated Mexican workers (those with at least a college degree) have the lowest emigration rate (5.3 percent in 2000).

The merging of Census data from Mexico and the United States to calculate Mexican emigration rates raises one potential problem. We have implicitly assumed that the education of Mexicans now residing in the United States is, in fact, the educational attainment that would have been observed had these workers remained in Mexico. This assumption does not create a problem for Mexicans who migrated as adults because only a very small number of these adults get additional schooling in the United States. However, the problem could be potentially severe for Mexicans who migrated as children. We are effectively assuming that the educational attainment they received in the U.S. education system parallels exactly the educational attainment they would have received in Mexico. This assumption is surely false and, as a result, we may be assigning “too much” education to these workers. The assumption, however, affects only a relatively small fraction of the sample of Mexican emigrants: Only 20.6 percent of the Mexicans in our 2000 U.S. Census sample migrated before age 14. We will show below that our results are not sensitive to alternative assumptions about the educational attainment of this subsample of workers.

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16 More precisely, only 1.3 percent of Mexican “stayers” (i.e., Mexicans enumerated by the Mexican census) in 1970 were high school graduates and only 1.2 percent had some college. Even by 2000, the fraction of workers who were high school graduates or had some college had risen to only 11.2 and 3.9 percent, respectively.

17 For example, only 6.9 percent of Mexicans who migrated between the ages of 15-24 during the 1990s are enrolled in school in 2000.

18 Although the potential misclassification of the “child immigrants” into relatively higher education groups could potentially explain the very large emigration rates of Mexican workers who are high school graduates or have some college, the data suggest that the misclassification does not play an empirically important role. Suppose that we assign all child immigrants from Mexico the educational distribution of comparably aged Mexicans who did not emigrate. In 2000, this counterfactual increases the emigration rate of high school dropouts with 0-8 years of schooling from 16.0 to 17.6 percent, and decreases the emigration rate of high school graduates from 29.7 to 25.7 percent.
Our study uses the variation in supply shocks across skill groups defined in terms of educational attainment and work experience to identify the labor market impact of immigration. It turns out that there are substantial differences in the age structure of international migrants in all three countries, even within a specific education group, and the nature of these differences change over time. Figures 3 and 4 illustrate the trends in the immigrant supply shocks for Canada and the United States, respectively. The data clearly show that the Canadian supply shocks differ from the U.S. supply shocks not only in terms of the Canadian influx being disproportionately more educated, but also in the age structure of the immigrants. Within any education group, the immigrant influx in the United States in recent decades tends to particularly increase the supply of relatively young workers (with less than 20 years of experience). In contrast, the immigrant influx in Canada tends to most increase the supply of older workers (with more than 25 years of experience). For example, over 25 percent of Canadian workers with a high school diploma and 30 to 40 years of experience were foreign-born. In contrast, the immigrant share among the youngest Canadian workers with a high school diploma was only around 12 percent. In the United States, the immigrant share among younger high school graduates hovered around 15 percent, while the immigrant share among their older counterparts was around 10 percent. It seems that the Canadian workers who are, in theory, most likely to be adversely affected by immigration are a mirror image of the American workers who are most vulnerable to immigrant-induced supply shocks. In the United States, the targeted natives seem to be younger, low-skill workers. In Canada, they would seem to be older, high-skill workers.

Figure 5 illustrates the detailed trends in emigration rates for the Mexican labor market. The data reveals a significant change in the age distribution of the low-skilled emigrants. Before 1990, emigration rates for low-skill workers were either relatively constant across experience categories or tended to be higher for older workers. In 2000, however, the emigration rates of high school dropouts begin to exhibit a pronounced inverse-U shape, so that emigration rates tend to be largest for high school dropouts with 15 to 25 years of experience.

The supply shocks illustrated in Figures 3-5 form the key independent variable of our empirical analysis. Our objective is to determine the link between these supply shocks and the evolution of the wage structure in each of the three countries. The earnings data are drawn from

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19 To avoid clutter in the figures illustrating the supply shocks in Canada, we only show the trend lines for the 1971, 1981, 1991, and 2001 Census years.
each of the respective Censuses. Unless otherwise indicated, we restrict our calculation of mean log earnings for each skill group (i.e., each $i, j, t$ cell in each country) to workers who, in addition to the restrictions listed in the previous section, also report information on their earnings and are not enrolled in school. Note that, by construction, the sample includes both salaried and self-employed workers. In the case of the Canada and U.S. censuses, the available earnings data refer to annual earnings in the year prior to the Census. We use these data to construct measures of log annual earnings and log weekly earnings for each skill cell. The Mexican census reports the worker’s monthly earnings. As we will discuss below, it is unclear from the questionnaire whether Mexican workers responded to the census survey in terms of “usual” monthly earnings or last month’s earnings. All earnings are deflated to constant dollars (1999 for the United States and 2000 for Canada and Mexico).

To illustrate the international differences in the trends in relative wages over the sample period, Figure 6 plots the time series of the wage gap between college graduates and high school dropouts for both young (6 to 10 years of experience) and older (31 to 35 years of experience) workers, as well as the corresponding wage gap between college graduates and high school graduates. The relative wage of younger high-educated workers rose in both Canada and the United States after 1980. However, the trends are quite different for older workers: the relative wage of experienced college graduates rose in the United States, but fell substantially in Canada. It is curious that the workforce of older, high-skill workers is precisely the one that was hit the hardest by the Canadian immigrant supply shock. Although the different evolution of the wage structure in the two countries has received a great deal of attention, the factors generating these differences are not well understood (Boudarbat, Lemieux, and Riddell, 2003; Card and Lemieux, 2001; Beaudry and Green, 2000).

Finally, the trends in the relative wage of high skill workers in Mexico show little resemblance to either the Canadian or U.S. trends. The relative wage of high-skill workers fell substantially in Mexico between 1970 and 1990, before beginning to rise slowly in the 1990s. In an important sense, the remainder of this paper attempts to determine if the cross-country differences in the evolution of relative wages can be at least partly understood in terms of the different immigration-induced supply shocks experienced by each country.

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20 To simplify the presentation of the trends in relative wages, Figure 6 pools together all college graduates in Canada and the United States and pools together all high school dropouts in Mexico.
V. Descriptive Results

Let $y_{ijt}$ denote the mean value of a particular labor market outcome for men who have education $i$, experience $j$, and are observed at time $t$. We calculate $y_{ijt}$ using the sample of native-born men in our study of the Canadian and U.S labor markets, and the sample of Mexicans residing in Mexico (i.e., the “stayers”) in our study of the Mexican labor market. The empirical analysis reported in this section stacks these data across skill groups and calendar years and estimates the following regression model separately by country:

\begin{equation}
 y_{ijt} = \theta_{ijt} + S + X + T + (S \times X) + (S \times T) + (X \times T) + \varphi_{ijt},
\end{equation}

where $S$ is a vector of fixed effects indicating the group’s educational attainment; $X$ is a vector of fixed effects indicating the group’s work experience; and $T$ is a vector of fixed effects indicating the time period. The linear fixed effects in equation (2) control for differences in labor market outcomes across schooling groups, experience groups, and over time. The interactions $(S \times T)$ and $(X \times T)$ control for the possibility that the impact of education and experience changed over time, and the interaction $(S \times X)$ controls for the fact that the experience profile for a particular labor market outcome may differ across education groups. Note that the specification in (2) implies that the labor market impact of labor supply shocks is identified using time-variation within education-experience cells. The standard errors are clustered by education-experience cells to adjust for possible serial correlation.

The alternative dependent variables used in our study of labor market outcomes in Canada and the United States include the mean of log earned annual income and log earned weekly income, as well as the fraction of weeks worked during the calendar year prior to the Census (defined as weeks worked divided by 52 in the sample of all persons, including nonworkers). The Mexican census provides limited information on labor market outcomes. We use the log of earned monthly income.\(^{21}\) Because of differences in the coding of work status

\(^{21}\) The income measure available in the 1970 Mexican census is the worker’s total monthly income, which includes both earned and unearned income. The regression model in (2) includes period effects as well as interactions between period effects and both educational attainment and work experience. These fixed effects should adjust for the measurement problem created by the change in the definition of the income variable across censuses. The regression results are very similar when we exclude the 1970 Mexican cross-section from the analysis.
across survey years in the Mexican census, we also use the labor force participation rate of the
skill group as a dependent variable.\textsuperscript{22}

Table 1 reports both weighted and unweighted estimates of the adjustment coefficient $\theta$. The weighted regressions weigh the observations by the sample size used to calculate the
dependent variable $y_{ijt}$.\textsuperscript{23} Row 1 of the top panel presents the basic weighted estimates of the
adjustment coefficient for Canada. Consider initially the results when the dependent variable is
the log weekly earnings of native Canadian workers. The coefficient is -0.507, with a standard
error of 0.202. It is easier to interpret this coefficient by converting it to an elasticity that gives
the percent change in wages associated with a percent change in labor supply. Let $m_{ijt} = M_{ijt}/N_{ijt}$,
or the percentage increase in the labor supply of group $(i, j, t)$ attributable to immigration. We
define the “wage elasticity” as:

$$
\frac{\partial \log w_{ijt}}{\partial m_{ijt}} = \frac{\theta}{(1 + m_{ijt})^2}.
$$

By 2000, immigration had increased the number of workers in the Canadian labor market by
25.8 percent. Equation (3) implies that the wage elasticity—evaluated at the mean value of the
immigrant supply increase—can be obtained by multiplying $\theta$ by approximately .63. The wage
elasticity for weekly earnings is then -0.32 (or -0.507 $\times$ 0.63). Put differently, a 10 percent
immigrant-induced increase in the number of workers in a particular skill group reduces the
wage of that group by 3.2 percent.\textsuperscript{24}

\textsuperscript{22} The 1960 Mexican census does not provide detailed information on a person’s work status. We restrict
the analysis of the labor force participation rate to the 1970-2000 surveys.

\textsuperscript{23} We normalized the sum of weights to equal 1 in each cross-section in the Canadian and U.S. regressions
to prevent the latter censuses from contributing more to the estimation simply because each country’s population
increased over time. We discuss the weighting of observations in the Mexican sample in more detail below.

\textsuperscript{24} The regression model in (2) uses the immigrant share, $p$, rather than the (more natural) relative number
of immigrants, $m$, as the regressor. The main reason for using $p$ as the regressor is that the labor market outcomes
used in this paper tend to be nonlinearly related to $m$, and $p$ is approximately a linear function of log $m$. For example,
the regression of Canadian log weekly earnings on $m$ after controlling for all the fixed effects, yields a coefficient of
-0.21 (.11). The regression of log weekly earnings on $m$ and $m$ squared yields coefficients of -0.63 (.18) and .47 (.17),
respectively. If we evaluate this quadratic expression at the mean of the variable $m$ (0.26), we obtain a wage
elasticity of -0.39, which is closer to that implied by the estimate of $\theta$ from the regression specification that uses the
immigrant share.
The other coefficients reported in the first row of Table 1 indicate that immigration has a slightly more negative impact on the annual earnings of native-born Canadian workers, implying that immigration reduces their labor supply. In fact, the adjustment coefficient in the fraction of weeks worked regression is negative and significant. An immigrant-induced 10 percent increase in labor supply reduces annual earnings by 3.9 percent and the fraction of time worked by 1.5 percentage points.

The next two rows of Table 1 report the corresponding estimates for the United States and Mexico. In the U.S. data, the estimated adjustment coefficient $\theta$ in the log weekly earnings regression is -0.489, with a standard error of 0.223. A test of equality for the adjustment coefficient in the log weekly earnings regression estimated in Canada and the United States would obviously not reject the hypothesis that the two coefficients are equal. By 2000, immigration had increased the number of male workers in the U.S. labor market by 17.2 percent. Equation (3) then implies that the wage elasticity can be estimated by multiplying the coefficient $\theta$ by 0.73. The wage elasticity for weekly earnings in the U.S. is then equal to -0.36, essentially the same numerical response as in Canada. The data also indicate that log annual earnings are more sensitive to immigration in the United States than in Canada, mainly because the labor supply of native workers in the United States is more sensitive to immigration (although, again, a test of equality of adjustment coefficients between the United States and Canada would not reject the hypothesis that the coefficients are equal). An immigrant-induced 10 percent increase in supply reduces the fraction of time worked by 2.5 percentage points, and reduces annual earnings by 6.2 percent.

Finally, row 3 of the table reports the corresponding results for Mexico. Our evidence confirms Mishra’s (2003) finding of a strong positive correlation between the log monthly earnings of Mexican workers in a particular skill group and the emigration rate of that group. The coefficient is +.798 (.443). By 2000, the large-scale emigration of workers from Mexico had reduced the size of the Mexican workforce by 19.1 percent, implying that the wage elasticity is obtained by multiplying the adjustment coefficient times 0.70. The wage elasticity is +0.559, indicating that a 10-percent emigrant-induced reduction in labor supply increases monthly earnings by 5.6 percent. Given the relatively large standard errors of the adjustment coefficients, it is evident that the data cannot reject the hypothesis that the three wage elasticities are the same.
The comparison of the estimated wage elasticities across all three countries is not straightforward because it is difficult to interpret the Mexican monthly earnings data. First, as noted above, the earnings data provided by the 1970 Mexican census (total personal income) is not directly comparable to the data provided by the post-1990 censuses (total earned income). However, it is unlikely that the different earnings definitions across survey years substantially bias the adjustment coefficient. The fourth row of Table 1 re-estimates the Mexican regression model using only data drawn from the 1990 and 2000 cross-sections, and finds that the adjustment coefficient is quite similar (.841 with a standard error of .540).

Further, it is unclear if the adjustment coefficient estimated in the monthly earnings Mexican regression is more comparable to a coefficient estimated in an annual earnings or weekly earnings regression. In particular, there may be a sizable labor supply component in the variation of monthly earnings across Mexican workers. Workers were asked to report their earned income, and the census questionnaire allowed several reporting options. In particular, workers could report weekly, bimonthly, monthly, or annual earnings. The Mexican census bureau then constructed a “monthly earnings” measure, which is the variable available in the public use file. Although it is common for many Mexican workers to be paid by the month, there is also a large seasonal component in employment, particularly in the rural sector. As a result, variation in the monthly earnings measure could be capturing seasonal differences in labor supply across workers (making Mexican monthly earnings conceptually more similar to annual earnings in Canada and the United States). There is strong evidence suggesting that this may be the case. We re-estimated the regression model using the subsample of urban Mexican workers, a subset of workers unlikely to be affected by seasonal fluctuations in agricultural demand. Row 5 of Table 1 clearly shows a drop in the estimate of the Mexican adjustment coefficient to +.652 (.419). The wage elasticity implied by the regression in the urban workers sample is +0.46.

Finally, Table 1 also reports evidence that the labor supply of Mexicans who remained behind is positively affected by the emigration of their compatriots. We estimated equation (2) using the labor force participation rate of the skill group as the dependent variable. The adjustment coefficient is positive (though only marginally significant). In sum, the regression coefficients reported in Table 1 suggest a remarkable similarity in the labor market impact of

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25 Unfortunately, the underlying information used to construct monthly earnings is not available in the IPUMS website, the Mexican census bureau website, or in the public use sample.
supply shocks induced by international migration throughout much of the North American continent: international migration raises wages and increases employment propensities whenever it decreases the size of the workforce, and lowers wages and reduces employment propensities whenever it increases the size of the workforce.

The bottom panel of the table repeats the regression analysis using an unweighted regression model. A comparison of the two panels in Table 1 indicates that the weighted and unweighted regression coefficients are relatively similar for Canada and the United States, but differ strikingly for Mexico when the regressions use the entire 1960-2000 sample period. However, the unweighted regression results for Mexico are unlikely to be very informative. The construction of the Census sample extracts in the various countries suggests that the weights (i.e., the number of observations used to calculate the dependent variable in the specific education-experience cell) are likely to be unimportant in Canada—where the sample extract represents a very large random sample of the population in each cross-section (33.3 percent in 1971 and 20 percent in all other cross-sections), and somewhat less important in the United States, where the sample extract represents a 1 percent random sample in 1960, 3 percent in 1970, and a 5 percent random sample beginning in 1980. The weights play a much greater role in Mexico, where the pre-1990 data represents roughly a 1 percent random sample, but the 1990 and 2000 data represent a 10.0 and 10.6 percent random sample, respectively.26 As a result, the sample size of the pre-1990 samples is much smaller than that of the post-1990 samples, implying that there is likely to be substantial measurement error in the measures of the cell-specific 1960 and 1970 labor market outcomes and emigration shares. Table 1 strongly suggests that this is the case. In particular, note that the weighted and unweighted adjustment coefficients are much closer in value when the regression uses only post-1990 data.

We conducted a variety of sensitivity tests to determine the robustness of our findings to major specification changes in the regression model. Table 2 reports the weighted regression coefficient \( \theta \) obtained from estimating these additional specifications using the log of annual earnings and the log of weekly earnings for Canada and the United States, and the log of monthly earnings for Mexico. For reference purposes, the first row of the table duplicates the baseline coefficients estimated in Table 1.

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26 Moreover, the Mexican population has grown rapidly over time. The 1970 Mexican census has approximately 483,000 observations, while the 2000 Mexican census has 10.1 million observations.
As noted earlier, our study is restricted to a sample of working men in each of the three countries. In particular, the mean labor market outcomes for each skill group are calculated using only the sample of working men, and the immigrant supply shocks (as measured by the immigrant share) only refer to the supply shifts occurring among working men. The second row of Table 2 reports the estimated coefficient $\theta$ when the measure of the immigrant or emigrant share $p_{it}$ uses information on both male and female workers. Despite the likely misclassification of many women into the various experience categories, the estimated adjustment coefficients remain significant, and have roughly the same numerical values as those reported in the baseline row. Similarly, row 3 shows that our findings are unaffected even when one adds working women into the calculation of the mean earnings for each skill group, so that the regression now uses men and women both in the calculation of the dependent variable as well as in the calculation of the immigrant/emigrant share.

Row 4 of the table returns to the baseline sample of male workers but examines instead the labor market impact of immigration using only salaried workers (i.e., it excludes the self-employed when calculating both the cell-specific labor market outcomes and the immigrant/emigrant shares). The exclusion of the self-employed from the regression analysis in Canada and the United States leads to roughly similar adjustment coefficients as those reported in the baseline row. The adjustment coefficient for Mexico, however, falls by about half, and is only now marginally significant. Nevertheless, the use of the salaried sample leads to a remarkable aligning of the adjustment coefficient across all three countries. The shift in the Mexican coefficient when the self-employed are excluded probably occurs because, as noted earlier, the log of monthly earnings in the Mexican census likely contains a significant labor supply component, and this labor supply variation may be substantially greater among the self-employed.

Our analysis has used a 5-year experience aggregation to define the various skill groups. As row 5 of Table 2 shows, the estimated coefficients are unaffected when we exclude workers who have just entered the labor market (i.e., have fewer than 5 years of experience) or workers who are about to retire (have more than 35 years of experience). Similarly, row 6 indicates that the results are robust when we use a 10-year experience aggregation (i.e., 1-10 years of experience, 11-20, and so on) rather than the 5-year aggregation to define the skill groups.
The remaining rows of Table 2 show that the results do not tend to be very sensitive to alternative definitions of the schooling categories or to the wholesale omission of particular schooling groups from the analysis. For instance, we have used a five-way educational classification in all three countries. Row 7 of the table illustrates what happens when we estimate the regressions using only 4 education groups by pooling all college graduates in Canada and the United States, and pooling all high school dropouts in Mexico. As the table shows, a four-way classification of education groups in the three countries actually leads to slightly more negative adjustment coefficients in Canada and the United States, and a much more positive coefficient in Mexico.

It is also of interest to investigate if the regression results are driven by a particular education group. Row 8 illustrates the impact of simply removing the sample of high school dropouts from the analysis in the United States and Canada. The results for Canada are not greatly affected, but the results for the United States become much weaker and are no longer statistically significant. Row 9 reports an equally interesting exercise when we redefine the high school dropout category in the United States to include only workers who have between 9 and 11 years of schooling. In 2000, only 1.7 percent of native-born working men in the United States had fewer than 8 years of schooling, as compared to 22.1 percent of immigrants. It is unclear that these very low-skill immigrants are competing in the same labor market as the native high school dropouts (where 80 percent have 9 to 11 years of schooling). To analyze the sensitivity of our results to this potential mismatch within the high school dropout category, we simply omitted from the calculation of both the cell-specific mean labor market outcomes and the cell-specific supply shocks any workers who have fewer than 8 years of school. The high school dropout category, therefore, now shows the relation between native and immigrant workers who have 9 to 11 years of schooling. Table 2 shows that the estimated adjustment coefficients in the United States are now negative, but still have large standard errors.

Row 10 provides a mirror image result for the Canadian estimates. As we showed earlier, immigration into Canada has particularly increased the supply of highly educated workers. If we omit all college graduates from the analysis, the U.S. results are basically unchanged, but the Canadian adjustment coefficients are near zero. If we only exclude the most highly educated

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27 It would be close to nonsensical to remove this sample from the study of Mexican labor market outcomes because high school dropouts make up the preponderance of that workforce.
workers from the data (workers with post-graduate education), row 11 shows that the adjustment coefficients again becomes negative and close in value to that presented in the baseline row. The lesson, therefore, seems to be that the observed negative labor market impact of immigration remains visible in the data as long as we do not remove from the data almost all of the variation in the immigrant supply shock.

Finally, row 12 shows that the results are robust when we adjust the Mexican emigration rates for the possibility that “child immigrants” obtained their education in the United States and would have obtained far less education had they remained in Mexico. In particular, we assign each person who was 14 or younger at the time of their emigration the educational distribution of comparably aged Mexicans who remained in Mexico. This assignment tends to slightly raise the emigration rate of high school dropouts and to slightly lower the emigration rate of workers in the middle of the education distribution. Table 2 shows that this counterfactual measure of the Mexican emigration rates barely changes the estimated wage impact of international migration on Mexican wages.

In sum, a systematic sensitivity analysis of the regression model in equation (2) across three countries greatly affected by international migration reveals that migration-induced labor supply shifts generate opposite-signed wage shifts, as predicted by the simplest textbook model of a competitive labor market. Remarkably, these wage responses seem to be of almost the same magnitude across countries. A 10 percent shift in supply induced by international migration generates a 3 or 4 percent opposite-signed shift in wages.

VI. Structural Estimates

A. Theory and Evidence

The regression results presented in the previous section are best interpreted as the findings from a descriptive exercise that reveals an important inverse relation, roughly similar across the three countries under study, between supply shifts and wages. The exercise did not impose any theoretical structure to estimate the response elasticities. A fuller understanding of the underlying correlation (and of its implications) requires imposing more structure on the data by specifying the technology of the aggregate production function. This structural approach would make it possible to estimate not only the effect of a particular supply shift on the wage of competing workers, but also the cross-effects on the wage of other workers. Borjas (2003)
proposed an empirically tractable approach by assuming that the aggregate production function can be represented in terms of a three-level CES technology: Similarly educated workers with different levels of work experience are aggregated to form the effective supply of an education group; and workers across education groups are then aggregated to form the national workforce.28

Suppose the aggregate technology for a national economy at time $t$ is given by the linear homogeneous production function:

$$(4) \quad Q_t = \left[ \lambda_K K_t^v + \lambda_L L_t^v \right]^{1/v},$$

where $Q$ is output, $K$ is capital, $L$ denotes the aggregate labor input; and $v = 1 - 1/\sigma_{KL}$, with $\sigma_{KL}$ being the elasticity of substitution between capital and labor ($-\infty < v \leq 1$). The vector $\lambda$ gives time-variant technology parameters that shift the production frontier, with $\lambda_K + \lambda_L = 1$. The aggregate $L_t$ incorporates the contributions of workers who differ in both education and experience. Let:

$$(5) \quad L_t = \left[ \sum_i \theta_{it} L_{ii}^\rho \right]^{1/\rho},$$

where $L_{ii}$ gives the number of workers with education $i$ at time $t$, and $\rho = 1 - 1/\sigma_E$, with $\sigma_E$ being the elasticity of substitution across these education aggregates ($-\infty < \rho \leq 1$). The $\theta_{it}$ give time-variant technology parameters that shift the relative productivity of education groups, with $\sum \theta_{it} = 1$. Finally, the supply of workers in each education group is itself given by an aggregation of the contribution of similarly educated workers with different experience. In particular:

$$(6) \quad L_{ii} = \left[ \sum_j \alpha_{ij} L_{ijit}^\eta \right]^{1/\eta},$$

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28 The three-level CES technology slightly generalizes the two-level approach used in the labor demand context by Bowles (1970) and Card and Lemieux (2001).
where \( L_{ijt} \) gives the number of workers in education group \( i \) and experience group \( j \) at time \( t \); and \( \eta = 1 - 1/\sigma_X \), with \( \sigma_X \) being the elasticity of substitution across experience classes within an education group \((-\infty < \eta \leq 1)\). Equation (6) incorporates an important identifying assumption: the technology coefficients \( \alpha_{ij} \) are constant over time, with \( \sum_j \alpha_{ij} = 1 \).

Borjas (2003) shows that the elasticities of substitution \( \sigma_X \) and \( \sigma_E \) can be estimated in a two-equation framework that uses the data on mean wages and immigrant supply shocks for specific skill groups summarized in the previous sections. In particular, the marginal productivity condition implies that the wage for skill group \((i, j, t)\) can be written as:

\[
\log w_{ijt} = \delta_t + \delta_u + \delta_y - \frac{1}{\sigma_X} \log L_{ijt},
\]

where \( \delta_t = \log \lambda_{Lt} + (1 - \nu) \log Q_t + (\nu - \rho) \log L_t \), and is absorbed by period fixed effects; \( \delta_u = \log \theta_{it} + (\rho - \eta) \log L_{it} \), and is absorbed by interactions between the education fixed effects and the period fixed effects; and \( \delta_y = \log \alpha_{ij} \), and is absorbed by interactions between education fixed effects and experience fixed effects. The regression model in (7), therefore, identifies the elasticity of substitution across experience groups.

Moreover, the coefficients of the education-experience interactions in (7) identify the parameters \( \log \alpha_{ij} \). As indicated by equation (6), the estimates of \( \alpha_{ij} \) and \( \sigma_X \) permit the calculation of \( L_{it} \), the CES-weighted labor aggregate for education group \( i \).\(^{29}\) We can then move up one level in the CES technology, and recover the elasticity of substitution across education groups. Let \( \log w_{it} \) be the mean log wage paid to the average worker in education group \( i \) at time \( t \). The marginal productivity condition determining the wage for this group is:

\[
\log w_{it} = \delta_i + \log \theta_{it} - \frac{1}{\sigma_E} \log L_{it}.
\]

\(^{29}\) If \( \log \hat{\alpha}_y \) is an estimated fixed effect coefficient, then \( \hat{\alpha}_y = \exp(\log \hat{\alpha}_y) / \sum_j \exp(\log \hat{\alpha}_y) \), which imposes the restriction that the sum of the \( \alpha \)'s is 1.
This equation is closely related to the model estimated by Katz and Murphy [1992, p. 69] that examines how the wage differential between college and high school graduates in the United States varies with relative supplies. Note that \( \sigma_E \) cannot be identified if the regression included interactions of education-period fixed effects to absorb the term \( \log \theta_{it} \). There would be as many fixed effects as there are observations. We initially solve the identification problem by adopting the Katz-Murphy assumption that the technology shifters can be roughly approximated by a linear trend that varies across education groups.\(^{30}\)

It is important to note that ordinary least squares regressions of equations (7) and (8) may lead to biased estimates of \( \sigma_X \) and \( \sigma_E \) because the supply of workers to the various education groups is likely to be endogenous. We assume that the immigrant or emigrant influx in a particular country provides the supply shifter required to identify the labor demand function.\(^{31}\)

Although the three-level CES framework offers the crucial advantage of summarizing the aggregate technology in terms of three elasticities of substitution, this simplification comes at a cost. The CES specification greatly restricts the types of substitution that can exist among the various factors. For example, the elasticity of substitution across experience groups takes on the same value even though the different cohorts of workers vary widely in their experience, and the elasticity of substitution across education groups takes on the same value even though the skill groups vary widely in their education.

Finally, note that the empirical implementation of the three-level CES technology described above does not use any data on the aggregate capital stock, making it difficult to separately identify the value of \( \sigma_{KL} \). We will discuss below plausible assumptions that can be made about this parameter to simulate the impact of immigration on the labor market.

The first row of Table 3 reports the estimates of \( \sigma_X \) for Canada, Mexico, and the United States. For purposes of comparison with the more ambiguous earnings data available in the

\(^{30}\) Note, however, that the Katz-Murphy approximation is based on the analysis of U.S. wages trends and may not hold in other countries. We address this point in more detail below.

\(^{31}\) This instrument would be valid if the influx of international migrants into or out of particular skill groups were independent of the relative wages offered to the various skill categories. It is likely, however, that the number of international migrants in a skill group responds to shifts in the wage structure. Income-maximizing behavior on the part of potential immigrants (emigrants) would generate larger (smaller) supplies in those skill cells that had relatively high (low) wages. This behavioral response would tend to build in a positive correlation between the size of the influx of international migrants and wages in a skill group. The regression coefficients, therefore, understate the negative wage impact of a relative supply increase.
Mexican census, we show the estimates of the structural model using both the log of weekly earnings and annual earnings as the dependent variable (although we limit our discussion to the coefficients estimated from the weekly earnings regressions). In the United States, we obtain an estimate of $1/\sigma_X$ equal to 0.321, with a standard error of 0.120. This implies an elasticity of substitution across experience groups of around 3. The same regression estimated in Canada yields the much lower estimate for $1/\sigma_X$ of 0.030, with a standard error of 0.015, while the same regression estimated in Mexico implies that $1/\sigma_X$ equals 0.286, with a standard error of 0.136. There is, therefore, a sizable range of dispersion in this elasticity across the three countries (with Canada being the outlier).

As noted above, the 3-level CES framework imposed an important specification restriction in the equation that estimates $\sigma_X$. In particular, note that equation (7) does not include any fixed effects to absorb experience-specific changes in wages over time (the $X \times T$ fixed effects). This restriction follows directly from the assumption that the coefficients $\alpha_{ij}$ in equation (6) were time-invariant. One simple way of relaxing this assumption is to follow Card and Lemieux’s (2001, Table 4), and assume that there are linear trends, specific to each experience group, that affect the evolution of wages. Row 2 of Table 3 includes these experience-specific trends. The estimated $1/\sigma_X$ now falls within a narrower range for all three countries: 0.135 (.037) for Canada; 0.122 (.038) for the United States; and 0.234 (.072) for Mexico. These coefficients imply an elasticity of substitution across experience groups of between 4 and 8. Note that these estimated coefficients are only a bit lower than those estimated (using a very different conceptual experiment) by Card and Lemieux (2001, Table 4), where the estimate of $1/\sigma_X$ lies between 0.17 and 0.20 in Canada and the United States.

We use the elasticities of substitution estimated in row 2 to calculate the CES-weighted aggregates of the workforce for each education group. As noted above, we need to make a crucial assumption about the pattern of demand shifts for the various education groups to estimate the second-stage regression model in equation (8). We begin with the simplest assumption: the assumption of a linear education-specific trend. Row 3 of Table 3 reports the estimates obtained for this specification. The IV estimate of the coefficient of the log of the number of workers in a skill group (which estimates the parameter $-1/\sigma_E$) is negative in the
United States (-0.273, with a standard error of .153), essentially zero in Canada, and very large and negative in Mexico (a coefficient of -2.02, with a standard error of 1.59).³²

Other studies have found that the Canadian coefficient estimating the relation between relative wages of different education groups and relative supplies is nearly zero when one assumes a linear trend in this type of CES framework (e.g., Card and Lemieux, 2001, Table 4). Beaudry and Green (2000) report conducting detailed specification tests which confirm that the Katz-Murphy linear trend assumption does not fit the Canadian data and may not provide a useful simplification for understanding how relative supplies affect the wage gap across different education groups in Canada.

In fact, Autor, Katz and Kearny (2004) have recently argued that the linear trend assumption also fails to capture some of what happened to the U.S. wage structure after 1992, the publication date of the Katz-Murphy paper. They document that the growth rate in the relative demand for skilled workers slowed in the 1990s. In particular, they find a 20 percent decline in the secular growth rate of demand for skilled workers during the 1990s as compared to the growth rate prior to the 1990s. To capture this break in the linear trend, we included education-specific splines in equation (8) that allow for a change in the intercept and a change in the slope of the trend in the 1990s.³³ Row 4 of Table 3 shows that the estimate of \(-1/\sigma_E\) with this nonlinear trend is -0.327, although it is estimated imprecisely. This point estimate implies that the elasticity of substitution across education groups is 3.1, roughly twice the size of the Katz-Murphy estimate of 1.4.

The nature of the nonlinearity in the Canadian trends in education-specific relative demand is not well understood. We use a non-parametric approach to allow for the fact that the Katz-Murphy linearity assumption fails to identify the “true” slope of the relative demand curve.

³² To conserve degrees of freedom in the Mexican regression, we only allow for differences in the linear trend between high school dropouts (who make up most workers in Mexico) and all other workers, rather than allow linear trends for each education group. This restriction is discussed in more detail below.

³³ Autor, Katz, and Kearny (2004) show that the trend \((t)\) in the United States is best specified as \(\beta_1 t + \beta_2 dt\), where \(d\) is a dummy variable indicating an observation after 1992. Rewrite the trend as \(\beta_1 (1 + dr) t\), where \(r = \beta_2 / \beta_1\). Autor, Katz, and Kearny estimate \(r = 0.2\), so that we can build in the Autor-Katz-Murphy results into an alternative specification for the regression model by redefining the trend variable as linear up to the 1990 census, and as .8t for the 2000 observation. Specifically, for each education group, we could define a trend variable that increases at the rate of one per year between 1960 and 1990. The trend variable then increases at a rate of 0.8 per year between 1990 and 2000. The estimate of \(1/\sigma_E\) obtained by this method is 0.347, with a standard error of 0.253.
In particular, we again introduce splines in the linear trend to allow for nonlinearities. The splines allow for different slopes in the trend coefficient for each decade (i.e., the 1970s, the 1980s, and the 1990s). In addition, we allow the trend to have a different intercept in the 1990s, a decade where the Canadian wage structure is perceived to have changed substantially. Row 4 of Table 3 clearly shows that the introduction of this nonlinearity in the education-specific demand shifts leads to an estimate of $-1/\sigma_E$ that is much more similar to that found in the United States. The estimate is -0.416, with a standard error of 0.137, implying an elasticity of substitution across education groups of about 2.4.

The estimated elasticity of substitution across education groups in the Mexican data is also sensitive to the nature of the assumption made about the trend in relative demand shifts of education groups. The analysis of the Mexican data, however, requires additional consideration simply because there are so few observations available. The second-stage regression in the Mexican analysis has only 20 observations (4 cross-sections and 5 education groups). The introduction of education-specific generalized nonlinear trends would almost exhaust all available degrees of freedom. To permit identification, therefore, we assume that the structure of relative demand shifts differs mainly between the workers who are high school dropouts (the bulk of the workforce) and all other workers. In effect, we save degrees of freedom by restricting the education-specific trends to be the same for all workers with at least a high school diploma. We then define education-specific linear trends, education-specific splines (starting in 1990), and education-specific shifters in the trend (again beginning in 1990). As row 4 of Table 3 shows, the coefficient of the regression is -0.355 (0.099), implying an elasticity of substitution across education groups of 2.8.

Although it is reassuring that the results tend to be more similar across countries once we address the identification issues regarding relative shifts in demand, it is important to emphasize that the estimates from the second-stage models should be interpreted with caution. Both the original linear trend assumption in the Katz-Murphy model and the nonlinear trend assumptions used in Table 3 play an important role in the identification of the elasticity of substitution, and different assumptions about the trend lead to very different estimates of the elasticity of substitution. Put bluntly, the assumption that must be made about trends in relative demand shifts in order to estimate $\sigma_E$ “may not be innocuous” (Card and Lemieux, 2004, p. 713). Our own experience with the data from all three countries suggests that the estimate of the coefficient in
the second stage tends to be more negative once we introduce some nonlinearity in the demand shocks. The results summarized in Table 3, therefore, are representative of a variety of different specifications that lead to roughly similar results.

Finally, the last row of Table 3 shows that the choice of the elasticity of substitution across experience groups chosen to calculate the CES-weighted aggregate of the size of the workforce in an education group does not play an important role. In particular, we estimated equation (8) using the actual log size of the workforce (rather than the CES-weighted aggregate). The estimates of the elasticity of substitution across education groups reported in rows 4 and 5 of the table are quite similar.

B. Simulating the Wage Effects of Immigration

The factor price elasticity giving the impact on the wage of factor $y$ of an increase in the supply of factor $z$ is defined as:\(^{34}\)

\[
\varepsilon_{yz} = \frac{d \log w_y}{d \log L_z} = s_z \frac{Q_y}{Q_y} \frac{Q_z}{Q_z}.
\]

where $s_z$ is the share of income accruing to factor $z$; and $Q_y = \partial Q/\partial L_y$, $Q_z = \partial Q/\partial L_z$, and $Q_{yz} = \partial^2 Q/\partial L_y \partial L_z$.

Borjas (2003, pp. 1365) uses this definition to derive the factor price elasticities implied by the three-level CES technology.\(^{35}\) These elasticities depend on the various elasticities of substitution as well as factor shares. Our simulation assumes that labor’s share of income is 0.62

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\(^{34}\) The factor price elasticity holds marginal cost and the quantities of other factors constant.

\(^{35}\) There are three factor price elasticities that are of interest in the three-level CES framework: the own factor price elasticity, giving the impact of an increase in the supply of workers with a particular education and experience level on their own wage; the within-education-branch factor price elasticity, giving the impact of an increase in the number of workers with a particular education and experience on the wage of workers with the same education but different experience; and the across-education-branch factor price elasticity, giving the impact of an increase in the number of workers with a particular education and experience on the wage of workers with different education.
in Canada, 0.7 in the United States, and 0.67 in Mexico.\textsuperscript{36} We use the latest cross-section available for each of the three countries to apportion the assumed aggregate labor share among the various education-experience groups.\textsuperscript{37}

For each country, the simulations use the estimated elasticity of substitution across age groups obtained in the log weekly earnings generalized specification (row 2 in Table 3), and the estimated elasticity of substitution across education groups estimated from the log weekly earnings nonlinear trend specification (row 4 in Table 3). The simulations also require an assumption about $\sigma_{KL}$. We conduct the calculations using three alternative assumptions: 0.5, 1.0, and 1.5. In the U.S. context, Hamermesh (1993, p. 92) concludes that the aggregate U.S. economy can be reasonably described by a Cobb-Douglas production function, suggesting that $\sigma_{KL}$ equals one. In the Canadian context, the available evidence suggests that the elasticity is closer to 0.5 than it is to 1.0 (Perrier, 2005). Finally, Shah (1992) estimates that $\sigma_{KL}$ is around 0.9 in the Mexican economy.

We use the calculated elasticities of factor price to simulate the impact of the flow of international migrants entering (or leaving) a particular country in the 1980-2000 period. The marginal productivity condition for the typical worker in education group $s$ and experience group $x$ can be written as $w_{sx} = D(K, L)$, where $L$ is a vector giving the number of workers in each education-experience group. Suppose initially that the capital stock is constant—so that capital does not adjust at all to the immigrant influx. The short-run impact of immigration on the log wage of group $(s, x)$ is then given by:

\begin{equation}
\Delta \log w_{sx} = \sum_i \sum_j e_{sx,ij} m_{ij}, \tag{10}
\end{equation}

where $m_{ij}$ gives the migrant-induced percentage change in labor supply in skill cell $(i, j)$. Because we are interested in simulating the impact of a particular cohort of international migrants, the

\textsuperscript{36} These assumptions are based on the estimates of Autor, Katz, and Krueger (1998) for the United States, and International Monetary Fund (2004) for Mexico. We used data on total labor compensation and total value added of the business sector for 1961-2001 to compute the Canadian labor share.

\textsuperscript{37} We use annual earnings (monthly earnings in the case of Mexico) for men and women in the latest census cross-section available in each country to calculate the fraction of all reported earnings accruing to each education-experience cell.
cohort that migrated between 1980 and 2000, our simulation of the Canadian and U.S. labor markets defines the supply shock as:

\[
m_{ij} = \frac{M_{ij,2000} - M_{ij,1980}}{0.5(N_{ij,1980} + N_{ij,2000}) + M_{ij,1980}},
\]

so that the baseline population used to calculate the percent increase in labor supply averages out the (changing) size of the native workforce during the 1980-2000 period and treats the pre-existing immigrant population as part of the “native” stock.

We do not have Mexican census data for 1980. Nevertheless, we can calculate the impact of the 1980-2000 emigrant flow because our emigration data is drawn from the U.S. Census. We define the supply shock used in the Mexican simulation as:

\[
m_{ij} = \frac{M_{ij,2000} - M_{ij,1980}}{0.5(N_{ij,1990} + N_{ij,2000}) + (M_{ij,2000} - M_{ij,1980})},
\]

where we use the average of the “stayer” workforce (i.e., Mexicans who remained in Mexico) over the available period, and we ignore the emigrants who had left Mexico prior to 1980 in calculating the percent change in labor supply attributable to emigration.

Table 4 summarizes the results of the simulation. Consider initially the top panel, which presents the results for Canada. The first row of the panel reports the measure of the supply shock attributable to the 1980-2000 immigrant influx, by education group. On average, this immigrant influx increased labor supply by 10.4 percent. The next three rows report the results of the simulation implied by equation (10), where the average percent wage change calculated for each education group is obtained by weighting the predicted impact for each education-experience cell by the income share accruing to that cell.

Regardless of the assumed value of the elasticity of substitution between labor and capital, the Canadian simulation reveals that, in the short run, low-skill workers either lost slightly or gained from the immigrant influx, while high-skill workers lost substantially. Because the available evidence suggests that \(\sigma_{KL}\) in Canada is close in value to 0.5, the short-run simulation indicates that the wage of high school dropouts was basically unaffected by
immigration, the wage of high school graduates fell by 4.2 percent, and the wage of workers with a post-graduate degree fell by almost 15 percent.

Of course, over time the capital stock will adjust as investors take advantage of the immigrant-induced higher-than-normal rental rate of capital. If the capital stock adjusted completely to the immigrant influx, the rental rate of capital would return to its pre-existing level. It can be shown that this alternative counterfactual implies that the long-run impact of immigration on the log wage of group \((s, x)\) is given by:

\[
\Delta \log w_{sx} = \frac{s_K}{\sigma_{KL}} \bar{K} + \sum_i \sum_j \varepsilon_{sxij} m_{ij},
\]

where \(s_K\) is capital’s share of income; and \(\bar{K}\) is the percent change in the capital stock induced by immigration. It is easy to show that the optimal change in the capital stock, \(\bar{K}\), can be represented as a weighted average of the immigrant supply shocks in the various skill groups, where the weights are the shares of income accruing to the various education-experience cells.\(^{38}\)

Note that equation (12) differs from equation (10) only by a positive constant, \((s_K / \sigma_{KL}) \bar{K}\). Full capital adjustment mutes the adverse wage impact of immigration but leaves the relative wage effects unchanged. In fact, because the first-level production function in equation (4) has constant returns to scale, the immigrant supply shock induces a corresponding shift in the capital stock that leaves the capital/labor ratio constant and the average wage in the economy unchanged (where the average wage is the weighted average of the group-specific wage changes in (12), weighted by the group’s share of income). Moreover, the long-run effects are the same regardless of the assumed value of the elasticity of substitution between labor and capital.

The possibility that the capital stock is not constant as immigrants enter the country may be particularly important in the Canadian context, even in the short run. Canadian immigration policy awards visas to independent migrants under the investor category. These investor-immigrants are required to make investments in Canada. In particular, an investor makes a prescribed investment of $400,000 payable to the Receiver General of Canada. These funds are

\(^{38}\) To simplify notation, let \(\ell\) be the subscript indicating the education-experience skill group. It is easy to show that the immigrant-induced change in the capital stock \(\bar{K} = \sum_{\ell} s_{L} m_{\ell} / s_{L}\), where \(s_{L}\) is labor’s share of income.
allocated to participating provinces and territories in Canada, which typically use the funds for job creation and economic development. The full amount of the investment (without interest) is repaid to the investor after approximately five years.39

The long-run simulation reported in Table 4 for Canada indicates that, once full capital adjustment takes place, the complementarity among the various education groups implies that low-skill workers in Canada have gained substantially from the country’s pursuit of an immigration policy that encourages the admission of high-skill workers, while high-skill workers have lost. In particular, the wage of high school dropouts increased by 6.0 percent, while the wage of workers with at least a college education fell by 6 to 8 percent. Put differently, immigration has increased the wage of high school dropouts relative to the wage of workers with at least a college diploma by at least 12 percent.

As we noted earlier, the estimates of the various elasticities of substitution in the three countries are extremely sensitive to the assumptions made regarding the trends in the technological parameters. One simple way of assessing the robustness of the simulation results to these assumptions is to simply use the own wage elasticity estimated in the descriptive regressions presented in Section V. The regression coefficients reported in Table 1 (from a generic regression of log weekly wages on the size of the immigrant supply shock and various vectors of fixed effects) indicate that the own wage elasticity in Canada is -0.32. The “non-structural elasticity” row in Table 4 uses this estimate of the own elasticity and sets all cross-elasticities equal to zero to simulate the impact of immigration on the Canadian wage structure.40 The results clearly show that the implications of the non-structural evidence for the impact of immigration on the wage structure are very similar to those obtained from the structural model when one assumes that the elasticity of substitution between labor and capital is between 0.5 and 1.0. In short, although the structural approach provides a much more complete picture of how the

39 The funds collected by this program totaled about $400 million per year between 1986 and 2002 (Business Immigration Program Statistics 2002, Citizenship and Immigration Canada).

40 Because cross-elasticities of factor price estimated in structural models of labor demand are typically very small, the assumption that they are zero does not impose a gross distortion on the data. In the Canadian context, for instance, the three-level CES framework implies that the cross-elasticity of log weekly wages with respect to shifts in the number of workers in other skill groups is smaller than 0.05 (in absolute value). This evidence is consistent with the full set of cross-elasticities reported in Borjas (2003, p. 1367) for the United States.
supply shocks work themselves through the labor market, the implications of the structural evidence are remarkably similar to those of the simpler correlations estimated in Section V.

It is of interest to compare these predicted wage shifts with what actually happened to the Canadian wage distribution. The bottom row of the top panel reports the change in the (real) log weekly wage of each education group between 1980 and 2000. All education groups in Canada suffered a decline in the real wage between 1980 and 2000. It is evident that immigration into Canada cannot explain the differences in relative wage shifts across education groups. There are clearly many other important factors that affect the Canadian wage structure and that have been essentially absorbed by the fixed effects included in the regression models. Note, however, that immigration may well be a crucial factor in accounting for the disproportionately large drop in real wages suffered by workers with post-graduate degrees in Canada. Regardless of the assumption made about the elasticity of substitution, and regardless of whether the simulation is conducted in the short or long run, the immigrant-induced relative wage drop for this group is larger than the relative wage drop that actually occurred. Immigration, in effect, prevented the real wage of the most educated workers in Canada from increasing dramatically, as it did in the United States.

Finally, note that our long-run simulation assumes that the rental rate of return to capital adjusts back to what it was in the pre-immigration regime. It is likely that there exist other flows of goods and capital (and increasingly so after the launching of NAFTA in 1994) that will tend to alter the rate of return to capital in all three countries. Our simulation ignores the existence and wage impact of these flows. Note, however, that the three-level CES framework implies that the relative wage effects of immigration on the various education groups would be precisely those revealed by the simulation in Table 4 even after accounting for these other factors.

In contrast to the significant immigrant-induced narrowing of the wage distribution that occurred in Canada, immigration into the United States has significantly widened the wage distribution. The 1980-2000 influx increased the average size of the workforce by about 11.1 percent, but the supply shock was much greater for low-skill workers. If we assume that the elasticity of substitution between labor and capital in the United States is 1.0, the short-run simulation reveals that immigration lowered the wage of high school dropouts by 7.4 percent, the wage of workers in the middle of the education distribution by about 2.5 percent, and the wage of workers with at least a college education by just under 4 percent. In the long run, after the
rental rate of return to capital adjusts back to its level in the pre-immigration regime, the simulation indicates a sizable drop of 4.1 percent in the relative wage of high school dropouts, a 1 percent gain in the relative wage of workers in the middle of the education distribution, and a slight reduction in the wage of highly educated workers.

The United States actually experienced a 19.8 percent decline in the real wage of high school dropouts, an 8.2 percent increase in the real wage of college graduates, and a 20.0 percent increase in the wage of workers with a post-graduate education. Hence the simulation reveals that it is unlikely that immigration can explain much of the trend in real wages for the high-skill workforce, but strongly suggests that immigration played an important role in lowering the real wage of low-skill workers. The short-run simulation implies that immigration accounts for about 40 percent of the real wage drop, while the long-run simulation puts the impact of immigration at about 20 percent.

Finally, the Mexican simulations lead to a somewhat unexpected conclusion. Even though the 1980-2000 emigrant flow from Mexico reduced its workforce by 14.6 percent, and even though most Mexican immigrants in the United States are high school dropouts, the high school dropouts remaining in Mexico did not gain much from this population flow and could conceivably lose slightly in the long run! As emphasized earlier, the emigration rates in Mexico are higher for workers in the middle of the education distribution. The simulation shows that these workers, in fact, are the ones who gain the most from emigration. If the elasticity of substitution between labor and capital is 1.0, as implied by available estimates, large-scale emigration increased the wages of workers with high school diplomas or some college by over 10 percent in the short run, and increased the wages of high school dropouts by only 4 percent. If the capital stock were to fully adjust, Mexican emigration would increase the wage of workers in the middle of the education distribution by 5 to 8 percent, but slightly reduce the wages of workers at the bottom of the education distribution by 1 percent. Paradoxically, even though most Mexican immigrants in the United States are high school dropouts, the structure of the population flows out of Mexico may end up actually lowering the relative wage of high school dropouts remaining in Mexico. The explanation of the paradox lies in the fact that the higher emigration rates of workers with slightly higher education levels makes the skills provided by the high school dropouts remaining in Mexico relatively more abundant.
It is of interest to compare these predicted wage changes with what actually happened in the Mexican labor market between 1990 and 2000 (the 1980 wage data is not available). Note that the largest real wage increases occurred for workers with the highest education levels, particularly for workers with some college, and that the smallest wage increase (particularly in the urban sample that more finely controls for labor supply differences) occurs among high school dropouts. It is difficult to assess the importance of emigration from Mexico in determining the evolution of the wage structure since many other factors are clearly at work. Nevertheless, it is clear that, as predicted by the simulation, the emigration of large numbers of high school dropouts does not seem to have substantially increased the real wage of the low-skill Mexican workforce.

VII. Summary

One of the central questions in the economics of immigration concerns the impact of immigrants on the labor markets of sending and receiving countries. Economic theory suggests that, at least in the short run, immigrant-induced shifts in labor supply should lead to opposite-signed changes in the wage of competing workers. This wage response is a crucial parameter not only in the study of the efficiency and distributional impact of international migration, but also in the policy debate over how to best regulate the population flows.

This paper uses microdata drawn from the national censuses of Canada, the United States, and Mexico, and applies the same methodological framework to these data to examine the impact of international migration on the labor market. The comparative analysis generated a number of findings. Perhaps the most important is that there is a numerically sizable and statistically significant inverse relation between immigrant-induced shifts in labor supply and wages in each of the three countries.

Even though the evidence suggests a roughly similar numerical wage response to labor supply shocks in each of the countries—a 10 percent labor supply shift is associated with about a 3 or 4 percent decrease in wages—the impact of international migration on the wage structure

41 The calculation of the actual change in the real Mexican wage between 1990 and 2000 takes into account the 1000-to-1 Mexican peso devaluation that took place in 1993, as well as the 477.81 percent inflation rate during the decade reported by the Bank of Mexico. Given the large numbers involved in these adjustments, the numerical value of the percent wage change in the real wage should be interpreted with caution. The change in relative wages across education groups, however, is invariant to the deflators used.
differs significantly across countries. In Canada, international migration substantially narrowed wage inequality because immigrants in Canada tend to be disproportionately high-skill. In the United States, international migration substantially increased wage inequality because immigrants in the United States tend to be disproportionately low-skill. In Mexico, however, emigration rates are highest in the middle of the skill distribution and lowest at the extremes. As a result, international migration has greatly increased relative wages in the middle of the Mexican skill distribution and lowered the relative wages at the extremes. Paradoxically, the large-scale migration of workers from Mexico may have reduced slightly the relative wage of the low-skill workers remaining in that country. These findings may have significant implications for our understanding of the link between immigration policy and the economic impact of immigration. The differential policy choices made by the three largest countries in North America have led to substantially different impacts of international migration on their wage structures.

Our study focused exclusively on the cross-country differences in the distributional impact of international migration flows. It would be interesting to determine if these international differences in distributional impacts have been accompanied by sizable differences in the efficiency gains. Further, our analysis explicitly ignored the many factors that tie together the labor markets of the three countries. In theory, the flows of international migrants across the three countries—there is, after all a sizable population flow not only from Mexico to the United States, but also from Canada to the United States—should accelerate the process of income convergence in the North American continent, which, in turn, should reduce the incentives to migrate. Moreover, the launching of the North American Free Trade Agreement (NAFTA) in January 1994 will have inevitable repercussions on the rate of income convergence and on the size and skill composition of the population flows across the countries. The addition of the key variables that link together these labor markets to a study of the economic impact of international migration should greatly increase our understanding of perhaps the central question in the economics of immigration.
DATA APPENDIX: CONSTRUCTION OF CENSUS EXTRACTS
AND VARIABLE DEFINITIONS

Canada:
The data are drawn from the 1971, 1981, 1986, 1991, 1996 and 2001 Canadian Census microdata files maintained by Statistics Canada. Each of these confidential data files represents a 20 percent sample of the Canadian population, except for the 1971 file which represents a 33.3 percent sample. Statistics Canada also provides Public Use Microdata Files (PUMFs) to Canadian post-secondary institutions and to other researchers. The public use samples represent a much smaller proportion of the Canadian population (e.g., a 2.7 percent sample in 2001). The analysis is restricted to men aged 18-64. A person is classified as an immigrant if he reports being a landed immigrant in the Canadian census, and is either a noncitizen or a naturalized Canadian citizen; all other persons are classified as natives. Unless otherwise noted, sampling weights are used in all calculations.

Definitions of education and experience: We use the Census variables $dgreer$ indicating “highest degree, certificate and diploma” and $trnucr$ indicating “trade or non-university certificate” for the 1981 to 2001 Censuses to classify workers into five education groups: high school dropouts; workers with either a high school diploma or a vocational degree; workers with both a high school and vocational degree or a post-secondary certificate or diploma below Bachelor’s degree; Bachelor’s degree holders; and post-graduate degree holders. The coding of the relevant variables changes across Censuses. For the 2001 Census these five groups are identified by i) $dgreer=1$ or 11; ii) $dgreer=2$ or ($dgreer=3$ and $trnucr ≠5$ and $trnucr ≠7$); iii) $dgreer=4$ or $dgreer=5$ or ($dgreer=3$ and $trnucr=5$ or 7); iv) $dgreer=6$; and v) $dgreer=7, 8, 9$ or 10. The highest degree variable in the 1971 Census only identifies university degree, certificate and diploma holders (and aggregates all others as “not applicable”). We rely on years of grade school ($highgrad$), vocational training ($training$), and years of post-secondary education below university ($otheredu$) to make the 1971 classifications comparable to later Census years. Our construction of the education categories in 1971 assumes that if a worker does not have a Bachelor’s degree but has 2 or more years of post-secondary education below university level, that worker possesses a post-secondary certificate or diploma. We also assume that Canadians who have eleven or more years of grade school and were born in Newfoundland or Quebec Provinces are high school graduates. All other Canadian-born and all immigrant men need 12 or more years of grade school to be considered high school graduates. This assumption recognizes the existence of different schooling systems across provinces and assumes that a Canadian-born worker’s entire grade school education is completed in the province where they were born. Canadian censuses also provide detailed information on the number of years an individual attended grade school (the variable $hgradr$ in the 2000 census), post secondary education below university ($ps_otr$), and university ($ps_uvr$). We calculate the total years of schooling by adding these variables and define work experience as Age - Years of Education - 6. We restrict the analysis to persons who have between 1 and 40 years of experience. Workers are classified into one of 8 experience groups. The experience groups are defined in five-year intervals (1-5 years of experience, 6-10, 11-15, 16-20, 21-25, 26-30, 31-35, and 36-40).

Counts of persons in education-experience groups: The counts are calculated in the sample of men who do not reside in collective households, worked at some point in the past year (i.e., have a positive value for weeks worked in the previous calendar year), are not enrolled in
school, and are not in the armed forces during the reference week. The 1986 census does not provide school attendance information so that the construction of the 1986 sample ignores the school enrollment restriction. Our results are not sensitive to the exclusion of this cross-section from the analysis.

Annual and weekly earnings: We use the sample of men who do not reside in collective households, reported positive weeks worked and hours worked (during the reference week), are not in the armed forces in the reference week, and report positive earnings (sum of wages, farmi, and selfi variables, using the variable names corresponding to the 2001 Census). The 1971 census reports weeks worked in the calendar year prior to the survey as a categorical variable. We impute weeks worked for each worker as follows: 7 weeks for 1 to 13 weeks, 20 for 14-26 weeks, 33 for 27-39 weeks, 44 for 40-48 weeks and 50.5 for 49-52 weeks. The average log annual earnings or average log weekly earnings for a particular education-experience cell is defined as the mean of log annual earnings or log weekly earnings over all workers in the relevant population.

Fraction of time worked: This variable is calculated in the sample of men who do not reside in collective households and are not enrolled in school (with the exception of the 1986 Census where school attendance information is not available). The fraction of time worked for each person is defined as the ratio of weeks worked (including zeros) to 52.

United States:

The data are drawn from the 1960, 1970, 1980, 1990, and 2000 Integrated Public Use Microdata Samples (IPUMS) of the U.S. Census. In the 1960 Census, the data extract forms a 1 percent sample of the population. In the 1970 Census, the extract forms a 3 percent sample (obtained by pooling the state, metropolitan area, and neighborhood files). In 1980, 1990, and 2000, the data extracts form a 5 percent sample. The analysis is restricted to men aged 18-64. A person is classified as an immigrant if he was born abroad and is either a non-citizen or a naturalized citizen; all other persons are classified as natives. Unless otherwise noted, sampling weights are used in all calculations.

Definition of education and experience: We use the IPUMS variables educrec to first classify workers into four education groups: high school dropouts (educrec <= 6), high school graduates (educrec = 7), persons with some college (educrec = 8), college graduates (educrec = 9). The college graduate sample is split into workers with 16 years of schooling or with postgraduate degrees using the variables higrade (in 1960-1980) and educ99 (1990-2000). We assume that high school dropouts enter the labor market at age 17, high school graduates at age 19, persons with some college at age 21, college graduates at age 23, and workers with postgraduate degrees at age 25, and define work experience as the worker’s age at the time of the survey minus the assumed age of entry into the labor market. We restrict the analysis to persons who have between 1 and 40 years of experience. Workers are classified into one of 8 experience groups, defined in five-year intervals.

Counts of persons in education-experience groups: The counts are calculated in the sample of men who do not reside in group quarters, worked at some point in the past year (i.e., have a positive value for weeks worked in the period calendar year), are not enrolled in school, and are not in the military during the survey week.

Annual and weekly earnings: We use the sample of men who do not reside in group quarters, reported positive weeks worked and hours worked (last week’s hours in 1960 and 1970; usual hours in 1980 through 2000), are not in the military in the reference week, and report
positive earnings. Our measure of earnings is the sum of the IPUMS variables \textit{incwage} and \textit{incbusfm} in 1960, the sum of \textit{incearn}, \textit{incbus}, and \textit{incfarm} in 1970 and 1980, and is defined by \textit{incearn} in 1990-2000. In the 1960, 1970, and 1980 Censuses, the top coded annual salary is multiplied by 1.5. In the 1960 and 1970 Censuses, weeks worked in the calendar year prior to the survey are reported as a categorical variable. We imputed weeks worked for each worker as follows: 6.5 weeks for 13 weeks or less, 20 for 14-26 weeks, 33 for 27-39 weeks, 43.5 for 40-47 weeks, 48.5 for 48-49 weeks, and 51 for 50-52 weeks. The average log annual earnings or average log weekly earnings for a particular education-experience cell is defined as the mean of log annual earnings or log weekly earnings over all workers in the relevant population.

\textit{Fraction of time worked:} This variable is calculated in the sample of men who do not reside in group quarters and are not enrolled in school. The fraction of time worked for each person is defined as the ratio of weeks worked (including zeros) to 52.

\textit{Mexico:}

The data are drawn from the 1970, 1990, and 2000 IPUMS of the Mexican Census. The data extracts from a 1.5 percent sample in the 1960 census, a 1 percent sample in 1970, a 10 percent sample in 1990, and a 10.6 percent sample in 2000. The analysis is restricted to men aged 18-64. Unless otherwise noted, sampling weights are used in all calculations.

\textit{Definition of education and experience:} We use the IPUMS variable \textit{edattain} to classify workers into the five education groups: high school dropouts with 0-8 years of schooling (\textit{educrec} \geq 110 and \leq 212), high school dropouts with 9-11 years of schooling (\textit{educrec} = 221 or 222), high school graduates (\textit{edattain} = 311 or 321), persons with some college (\textit{edattain} = 312 or 322), and college graduates (\textit{edattain} = 400). Our definition implies that the sample of high school dropouts with 0-8 years of schooling contains workers who have, at best, completed primary education. The sample of high school dropouts with 9-11 years of schooling contains workers who completed primary school but have less than a secondary education. The sample of high school graduates contains those who completed only their secondary general track education or those who obtained a secondary technical degree. The sample of workers with some college includes workers who completed their secondary general track and report having some college or university, or who completed their secondary technical track and have a post-secondary technical degree. The sample of college graduates includes workers who report having completed a university education. Workers who do not report a valid value for the educational attainment variable are omitted from the analysis. There are some inconsistencies in our coding of education classification for workers who report completing a primary education, but less than a secondary education in a technical track. We recoded these workers as being high school dropouts with 0-8 years of school if they did not attend school for at least 9 years. There are also some inconsistencies for workers who report completing a secondary technical degree. Some of these workers report attending school for much fewer than 12 years, while others report attending school for much more than 12 years. We reclassified workers who report completing a secondary technical degree as high school dropouts if they attended school for less than 11 years, and recategorized them as “some college” if they attended school for more than 13 years. We assume that high school dropouts enter the labor market at age 17, high school graduates at age 19, persons with some college at age 21, and college graduates at age 23, and define work experience as the worker’s age at the time of the survey minus the assumed age of entry into the labor market. We restrict the analysis to persons who have between 1 and 40 years of experience. Workers are classified into one of 8 experience groups, defined in five-year intervals.
Counts of persons in education-experience groups: The counts of workers who remained in Mexico are calculated in the sample of men who are at work (the IPUMS employment status variable empstat = 1100), report a positive value for monthly income, and are not enrolled in school. The counts of emigrants are obtained from the U.S. Census are use the same restrictions listed above in our description of the calculation of counts in the U.S. data, with the additional restrictions that the emigrants must be persons who are immigrants and born in Mexico. The 1960 census does not report the variable empstat nor does it report a person’s school enrollment status. The sample of workers in 1960 is then composed of persons who report a positive value for monthly income.

Monthly earnings: The variable gives the total monthly earnings (both salary and self-employment income) in 1960, 1990, and 2000, and the total monthly income in 1970. We omit from the analysis persons who are not “at work” and do not report a valid, positive value for monthly earnings. In each census, the top-coded monthly earnings are multiplied by 1.5.

Labor force participation rate: This variable is calculated in the sample of men who provide a valid answer for their employment status variable (empstat). We define a labor force participant as someone who is either working or unemployed (i.e., someone who is not “inactive”). This variable is only available in the 1970, 1990 and 2000 cross-sections.
References


Figure 1. Trends in the immigrant/emigrant share for male workers, by country

Notes: The trend lines for Canada and the United States give the fraction of the workforce that is foreign-born; the trend line for Mexico gives the fraction of the Mexican workforce that emigrated to the United States.
Figure 2. Trends in the immigrant/emigrant share for male workers, by education and country

Canada:

United States:

Mexico:

Notes: The trend lines for Canada and the United States give the fraction of the workforce that is foreign-born; the trend line for Mexico gives the fraction of the Mexican workforce that emigrated to the United States.
Figure 3 – The immigrant supply shock in Canada, 1971-2001

Note: The immigrant share gives the fraction of the workforce that is foreign-born in a particular education-experience group.
Figure 4. The immigrant supply shock in the United States, 1960-2000

Note: The immigrant share gives the fraction of the workforce that is foreign-born in a particular education-experience group.
Figure 5. The emigrant supply shock in Mexico, 1960-2000

Notes: The emigrant share gives the fraction of the Mexican workforce in a particular education-experience group that emigrated to the United States.
Figure 6. Trends in the relative wage of college graduates, by years of experience

Canada:

United States:

Mexico:

Notes: The figures illustrate the log weekly wage gap between college graduates and the respective education groups. The “young” group of workers has 6 to 10 years of experience; the older group has 31 to 35 years. In this figure, the college graduate sample in the Canadian and U.S. data includes all persons who have at least a college degree, and the high school dropout sample in Mexico includes all persons with less than 12 years of schooling.
Table 1. Relation between the immigrant/emigrant share and labor market outcomes.

<table>
<thead>
<tr>
<th></th>
<th>Earnings outcomes</th>
<th>Employment outcomes</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Log annual earnings</td>
<td>Log weekly earnings</td>
</tr>
<tr>
<td>Weighted Regressions</td>
<td></td>
<td></td>
</tr>
<tr>
<td>1. Canada</td>
<td>-.617 (.246)</td>
<td>-.507 (.202)</td>
</tr>
<tr>
<td>2. United States</td>
<td>-.845 (.472)</td>
<td>-.489 (.223)</td>
</tr>
<tr>
<td>Mexico</td>
<td></td>
<td></td>
</tr>
<tr>
<td>3. All workers</td>
<td>---</td>
<td>---</td>
</tr>
<tr>
<td>4. All workers, 1990-2000</td>
<td>---</td>
<td>---</td>
</tr>
<tr>
<td>5. Urban workers</td>
<td>---</td>
<td>---</td>
</tr>
<tr>
<td>Unweighted Regressions</td>
<td></td>
<td></td>
</tr>
<tr>
<td>1. Canada</td>
<td>-.786 (.200)</td>
<td>-.651 (.148)</td>
</tr>
<tr>
<td>2. United States</td>
<td>-.781 (.706)</td>
<td>-.435 (.358)</td>
</tr>
<tr>
<td>Mexico</td>
<td></td>
<td></td>
</tr>
<tr>
<td>3. All workers</td>
<td>---</td>
<td>---</td>
</tr>
<tr>
<td>4. All workers, 1990-2000</td>
<td>---</td>
<td>---</td>
</tr>
<tr>
<td>5. Urban workers</td>
<td>---</td>
<td>---</td>
</tr>
</tbody>
</table>

Notes: Standard errors are reported in parentheses and are adjusted for clustering within education-experience cells. For Canada and the United States, the table reports the coefficient of the immigrant share variable from regressions where the dependent variable is the mean labor market outcome native-born persons in an education-experience group at a particular point in time. For Mexico, the table reports the coefficient of the emigrant share variable from regressions where the dependent variable is the mean labor market outcome of workers in an education-experience group of Mexican “stayers” at a particular point in time. The regressions estimated in Canada have 240 observations; the regressions estimated in the United States have 200 observations; and the wage (labor force participation) regressions estimated in Mexico have 160 (120) observations when the sample period is 1960-2000 (1970-2000) and 80 observations when the sample period is 1990-2000. All regression models include education, experience, and period fixed effects, as well as interactions between education and experience fixed effects, education and period fixed effects, and experience and period fixed effects.
Table 2. Sensitivity analysis of the relation between wages and the immigrant/emigrant share

<table>
<thead>
<tr>
<th>Sample:</th>
<th>Canada</th>
<th>United States</th>
<th>Mexico</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Log annual earnings</td>
<td>Log weekly earnings</td>
<td>Log annual earnings</td>
</tr>
<tr>
<td>1. Baseline (from table 1)</td>
<td>-.617</td>
<td>-.507</td>
<td>-.845</td>
</tr>
<tr>
<td></td>
<td>(.246)</td>
<td>(.202)</td>
<td>(.472)</td>
</tr>
<tr>
<td>2. Includes women in measure of immigrant supply shock</td>
<td>-.766</td>
<td>-.642</td>
<td>-.999</td>
</tr>
<tr>
<td></td>
<td>(.229)</td>
<td>(.191)</td>
<td>(.576)</td>
</tr>
<tr>
<td>3. Men and women in skill group</td>
<td>-.679</td>
<td>-.510</td>
<td>-.922</td>
</tr>
<tr>
<td></td>
<td>(.195)</td>
<td>(.149)</td>
<td>(.606)</td>
</tr>
<tr>
<td>4. Salaried workers only</td>
<td>-.518</td>
<td>-.398</td>
<td>-.755</td>
</tr>
<tr>
<td></td>
<td>(.222)</td>
<td>(.173)</td>
<td>(.409)</td>
</tr>
<tr>
<td>5. Omits workers with less than 5 years or more than 35 years of experience</td>
<td>-.678</td>
<td>-.567</td>
<td>-.822</td>
</tr>
<tr>
<td></td>
<td>(.158)</td>
<td>(.147)</td>
<td>(.416)</td>
</tr>
<tr>
<td>6. Uses 10-year grouping for experience</td>
<td>-.555</td>
<td>-.487</td>
<td>-1.231</td>
</tr>
<tr>
<td></td>
<td>(.474)</td>
<td>(.378)</td>
<td>(.835)</td>
</tr>
<tr>
<td>7. Uses four education groups, pools college graduates in Canada and U.S.; pools high school dropouts in Mexico</td>
<td>-.654</td>
<td>-.561</td>
<td>-1.112</td>
</tr>
<tr>
<td></td>
<td>(.248)</td>
<td>(.198)</td>
<td>(.527)</td>
</tr>
<tr>
<td>8. Omits high school dropouts</td>
<td>-.769</td>
<td>-.668</td>
<td>-.101</td>
</tr>
<tr>
<td></td>
<td>(.171)</td>
<td>(.145)</td>
<td>(.606)</td>
</tr>
<tr>
<td>9. Omits workers with 8 or fewer years of schooling</td>
<td>-.608</td>
<td>-.506</td>
<td>-.600</td>
</tr>
<tr>
<td></td>
<td>(.215)</td>
<td>(.182)</td>
<td>(.468)</td>
</tr>
<tr>
<td>10. Omits college graduates</td>
<td>-.081</td>
<td>.001</td>
<td>-.910</td>
</tr>
<tr>
<td></td>
<td>(.403)</td>
<td>(.318)</td>
<td>(.445)</td>
</tr>
<tr>
<td>11. Omits workers with post-graduate degrees</td>
<td>-.535</td>
<td>-.423</td>
<td>-1.134</td>
</tr>
<tr>
<td></td>
<td>(.301)</td>
<td>(.247)</td>
<td>(.486)</td>
</tr>
<tr>
<td>12. Adjusts education of persons who migrated from Mexico as children</td>
<td>---</td>
<td>---</td>
<td>---</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: Standard errors are reported in parentheses and are adjusted for clustering within education-experience cells. All coefficients are obtained from regressions weighted by the sample size used in computing the dependent variable. For Canada and the United States, the table reports the coefficient of the immigrant share variable from regressions where the dependent variable is the mean log weekly or log annual earnings of native-born persons in an education-experience group at a particular point in time. For Mexico, the table reports the coefficient of the emigrant share variable from regressions where the dependent variable is the mean log monthly earnings in an education-experience group of Mexican “stayers” at a particular point in time. All regression models include education, experience, and period fixed effects, as well as interactions between education and experience fixed effects, education and period fixed effects, and experience and period fixed effects.
Table 3. Estimates of first- and second-stage equations from the CES model

<table>
<thead>
<tr>
<th></th>
<th>Canada</th>
<th>United States</th>
<th>Mexico</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>First stage, estimates -1/σ_χ</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1. Simple specification</td>
<td>-.030</td>
<td>-.321</td>
<td>-.286</td>
</tr>
<tr>
<td></td>
<td>(.015)</td>
<td>(.120)</td>
<td>(.136)</td>
</tr>
<tr>
<td>2. Generalized specification</td>
<td>-.135</td>
<td>-.122</td>
<td>-.234</td>
</tr>
<tr>
<td></td>
<td>(.037)</td>
<td>(.038)</td>
<td>(.072)</td>
</tr>
<tr>
<td><strong>Second stage, estimates -1/σ_E</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>3. Linear trend, CES-weighted employment</td>
<td>-.047</td>
<td>-.273</td>
<td>-.2020</td>
</tr>
<tr>
<td></td>
<td>(.119)</td>
<td>(.153)</td>
<td>(1.594)</td>
</tr>
<tr>
<td>4. Nonlinear trend, CES-weighted employment</td>
<td>-.416</td>
<td>-.327</td>
<td>-.355</td>
</tr>
<tr>
<td></td>
<td>(.137)</td>
<td>(.379)</td>
<td>(.099)</td>
</tr>
<tr>
<td>5. Nonlinear trend, actual employment</td>
<td>-.391</td>
<td>-.348</td>
<td>-.342</td>
</tr>
<tr>
<td></td>
<td>(.210)</td>
<td>(.361)</td>
<td>(.089)</td>
</tr>
</tbody>
</table>

Notes: Standard errors are reported in parentheses. The standard errors reported in rows 1 and 2 are adjusted for clustering within education-experience cells. The dependent variable is log weekly earnings in Canada and the United States, and log monthly earnings in Mexico. All coefficients are obtained from regressions weighted by the sample size used in computing the dependent variable. The first-stage regressions have 240 observations in Canada, 200 observations in the United States, and 160 observations in Mexico. The second-stage regressions have 30 observations in Canada and the United States, and 20 observations in Mexico. The “nonlinear trend” in Canada includes the education-specific linear trend variables and education-specific splines for the 1980s and the 1990s. The nonlinear trend in the United States includes the education-specific linear trend between 1960 and 2000 and an education-specific spline in the trend line for the 1990s. The nonlinear trend in the Mexico includes an education-specific linear trend between 1960 and 2000 and an education-specific spline in the trend line for the 1990s. All second-stage regressions in all countries also include education-specific shifters for the 1990s to allow for possible level breaks in the trends. To conserve degrees of freedom, the Mexican analysis allows the education-specific trend lines to vary only among three groups: high school dropouts with 0 to 8 years of schooling, high school dropouts with 9 to 11 years of schooling, and all other workers.
### Table 4. Predicted percent wage impacts of 1980-2000 immigrant/emigrant supply shock

<table>
<thead>
<tr>
<th></th>
<th>High school dropouts</th>
<th>High school graduates</th>
<th>Some college</th>
<th>College graduates</th>
<th>Post-graduate degree</th>
<th>All workers</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Canada</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Supply shock (%)</td>
<td>-5.6</td>
<td>3.2</td>
<td>13.3</td>
<td>29.0</td>
<td>32.0</td>
<td>10.4</td>
</tr>
<tr>
<td>Short-run, $\sigma_{KL} = 0.5$</td>
<td>-0.6</td>
<td>-4.2</td>
<td>-7.6</td>
<td>-13.0</td>
<td>-14.8</td>
<td>-6.6</td>
</tr>
<tr>
<td>Short-run, $\sigma_{KL} = 1.0$</td>
<td>2.7</td>
<td>-0.9</td>
<td>-4.3</td>
<td>-9.7</td>
<td>-11.5</td>
<td>-3.3</td>
</tr>
<tr>
<td>Short-run, $\sigma_{KL} = 1.5$</td>
<td>3.8</td>
<td>0.2</td>
<td>-3.2</td>
<td>-8.6</td>
<td>-10.4</td>
<td>-2.2</td>
</tr>
<tr>
<td>Long-Run</td>
<td>6.0</td>
<td>2.4</td>
<td>-1.0</td>
<td>-6.4</td>
<td>-8.2</td>
<td>0.0</td>
</tr>
<tr>
<td>Non-structural elasticity</td>
<td>1.8</td>
<td>-1.3</td>
<td>-4.4</td>
<td>-8.7</td>
<td>-10.2</td>
<td>-4.2</td>
</tr>
<tr>
<td>Actual % wage change</td>
<td>-19.3</td>
<td>-16.2</td>
<td>-8.7</td>
<td>-2.2</td>
<td>-7.0</td>
<td>-6.1</td>
</tr>
<tr>
<td><strong>United States</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Supply shock (%)</td>
<td>22.8</td>
<td>8.1</td>
<td>9.2</td>
<td>12.0</td>
<td>12.8</td>
<td>11.1</td>
</tr>
<tr>
<td>Short-run, $\sigma_{KL} = 0.5$</td>
<td>-10.5</td>
<td>-5.7</td>
<td>-6.0</td>
<td>-6.9</td>
<td>-7.2</td>
<td>-6.6</td>
</tr>
<tr>
<td>Short-run, $\sigma_{KL} = 1.0$</td>
<td>-7.1</td>
<td>-2.3</td>
<td>-2.7</td>
<td>-3.6</td>
<td>-3.9</td>
<td>-3.3</td>
</tr>
<tr>
<td>Short-run, $\sigma_{KL} = 1.5$</td>
<td>-6.0</td>
<td>-1.2</td>
<td>-1.6</td>
<td>-2.5</td>
<td>-2.8</td>
<td>-2.2</td>
</tr>
<tr>
<td>Long run</td>
<td>-3.8</td>
<td>1.0</td>
<td>0.6</td>
<td>-0.3</td>
<td>-0.6</td>
<td>0.0</td>
</tr>
<tr>
<td>Non-structural elasticity</td>
<td>-8.2</td>
<td>-2.9</td>
<td>-3.3</td>
<td>-4.3</td>
<td>-4.6</td>
<td>-4.0</td>
</tr>
<tr>
<td>Actual % wage change</td>
<td>-19.8</td>
<td>-14.8</td>
<td>-4.1</td>
<td>8.2</td>
<td>20.0</td>
<td>1.7</td>
</tr>
<tr>
<td><strong>Mexico</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Supply shock (%)</td>
<td>-11.6</td>
<td>-11.6</td>
<td>-35.6</td>
<td>-29.7</td>
<td>-5.8</td>
<td>-14.6</td>
</tr>
<tr>
<td>Short-run, $\sigma_{KL} = 0.5$</td>
<td>8.6</td>
<td>8.6</td>
<td>17.1</td>
<td>15.0</td>
<td>6.5</td>
<td>9.6</td>
</tr>
<tr>
<td>Short-run, $\sigma_{KL} = 1.0$</td>
<td>3.8</td>
<td>3.8</td>
<td>12.3</td>
<td>10.2</td>
<td>1.7</td>
<td>4.8</td>
</tr>
<tr>
<td>Short-run, $\sigma_{KL} = 1.5$</td>
<td>2.2</td>
<td>2.2</td>
<td>10.7</td>
<td>8.6</td>
<td>0.1</td>
<td>3.2</td>
</tr>
<tr>
<td>Long run</td>
<td>-1.0</td>
<td>-1.1</td>
<td>7.5</td>
<td>5.4</td>
<td>-3.1</td>
<td>0.0</td>
</tr>
<tr>
<td>Non-structural elasticity</td>
<td>6.5</td>
<td>6.5</td>
<td>19.9</td>
<td>16.6</td>
<td>3.3</td>
<td>8.2</td>
</tr>
<tr>
<td>Actual % wage change, 1990-2000</td>
<td>10.7</td>
<td>3.0</td>
<td>7.0</td>
<td>24.7</td>
<td>18.7</td>
<td>16.1</td>
</tr>
<tr>
<td>Actual % wage change in urban sample, 1990-2000</td>
<td>4.7</td>
<td>3.8</td>
<td>7.5</td>
<td>23.7</td>
<td>18.0</td>
<td>13.8</td>
</tr>
</tbody>
</table>

Notes: The simulation uses the elasticity of substitution across experience groups reported in row 2 of Table 3, and the elasticity of substitution across education groups reported in row 5 of Table 3. The “non-structural elasticity” estimates uses the own wage elasticity estimated in Table 1 (-0.32 for Canada, -0.36 for the United States, and -0.56 for Mexico), and sets all cross-elasticities equal to zero. The variable measuring the group-specific immigrant supply shock in Canada and the United States is defined as the number of immigrants arriving between 1980 and 2000 divided by a baseline population equal to the average size of the native workforce (over 1980-2000) plus the number of immigrants in 1980. The variable measuring the emigrant supply shock in Mexico is defined as the number of persons who emigrated between 1980 and 2000 divided by a baseline native population equal to the average size of the Mexican stayer workforce (over 1990-2000) plus the number of persons who emigrated between 1980 and 2000. We used the share of income accruing to each of the cells to calculate the weighted averages of predicted wages reported in this table. The reported percent wage changes refer to the product of the log wage change times 100.