Skill Differences and the Effect of Immigrants on the Wages of Natives

by

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Abstract. This paper examines the effects of changes in the supply of immigrant labor on the wages of natives in the 1980s. Estimates are presented indicating that immigrants and natives are nearly perfect substitutes within broad skill categories. This result is then used to calculate the effects of the large influx of immigrants to the U.S. labor market. The calculations presented suggest that immigration depressed the wages of native dropouts by as much as 3 percent and can account for up to 24 percent of the increase in the college – high school wage differential in the 1980s.

Keywords. Immigration, Wage Inequality, Elasticity of Substitution, U.S. Census

1. Introduction

The 1980s witnessed the largest influx of immigrants to the U.S. since the first decade of the twentieth century, with more than 7.3 million immigrants being legally admitted between 1981 and 1990 (*Statistical Abstract of the United States* 1993). The continuing emigration of workers from other countries to the U.S. has rekindled popular interest in the effects of immigration on the native labor force. Much of this discussion has assumed that the labor market outcomes of natives (i.e. wages and employment) are adversely affected by increased levels of immigration. The existing empirical evidence largely suggests, however, that the potential impact of the large influx of immigrant labor during the 1980s on natives' wages was small (Borjas 1994 summarizes these findings).

Although the impact of immigration depends crucially on the relative skill distributions of immigrants and natives (Borjas 1995b), nearly all studies have treated the immigrant population as homogenous with respect to skill, even while looking at the effects of immigration on low–skill natives. In this essay, I disaggregate *both* the immigrant and native populations by skill and examine the differential impact of changes in skill distribution of immigrant labor during the 1980s on low- and high-skill natives. I first estimate the elasticity of substitution between immigrants and natives of the same sex with similar skills, testing Borjas, Freeman, and Katz's (1992, 1996) assumption of perfect substitutability. Even after accounting for various possible biases as well as adjusting for changes in the composition of the immigrant and native work force, I estimate that immigrants and natives possessing similar skills *are*, in fact, nearly perfect substitutes in production. Using this result, along with a simple aggregate production function model, I then calculate that changes in the skill distribution of immigrants can account for up to 24 percent of the increase in the college–high school wage premium that occurred during the 1980s. For native dropouts, immigration had an even greater effect, lowering the real wage level by as much as 3.6 percentage points, accounting for up to one third of the change during the decade.

Section 2 examines changes in the skill composition of natives and immigrants in the labor force. Section 3 documents changes in average productivity within different skill groups. Section 4 estimates the elasticity of substitution between immigrants and natives of the same sex possessing similar skills. Section 5 examines the effects of immigrants on changes in the wage levels of natives with different skills as well as changes in the native college–high school wage differential using a simple aggregate production function framework. Section 6 concludes.

I use the five percent Public Use Microsamples (PUMS) of the 1980 and 1990 Censuses for all of the analyses in the paper. The large samples available in these data permit disaggregation by nativity, sex, and skill within metropolitan areas. Further description of the methods used in constructing the data used can be found in the Data Appendix.

2. Changes in the Skill Composition of the Labor Force

The shift towards higher levels of skill in the U.S. labor force during the 1980s has been well–documented.[1] In this section I examine how changes in the relative skill distributions of immigrants and natives affected the supply of labor with different levels of skill during the decade.[2] Throughout the paper I use educational attainment as a proxy for skill and divide the labor force into four skill groups: high school dropouts, high school graduates, individuals with some (i.e. 1 – 3 years) college, and college graduates.[3]

Between 1980 and 1990 the immigrant share of all skill levels increased, with the largest relative increase occurring among the least-skilled.[4] Two factors affected immigrants' share of each skill level. First, the number of immigrants employed in the U.S. grew by about 55 percent during the decade causing the immigrant share of all skill levels to increase. Second, the movement towards higher levels of skill was greater among natives than among immigrants, causing a larger increase in the immigrant share among high school dropouts and high school graduates than in the immigrant share of higher skill levels. Table 1

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[1] See, for example, Katz and Murphy 1992.

[2] Immigrants are defined as individuals who were not born in the United States or its territories and who were not born to American parents living abroad.

[3] Because the educational attainment question in the Census changed between 1980 and 1990, I use the use the recoding scheme proposed by Jaeger (1995b) to classify individuals into these four categories.

[4] While the Census measures labor supply and income for 1979 and 1989, I will refer to the years in which the data were collected.

presents estimates of the size of native and immigrant employment within four broad levels of skill for both men and women. The immigrant share of total employment increased from 6.7 percent in 1980 to 8.9 percent in 1990, with greater growth occurring among men.[5] This overall growth in the immigrant population would have increased the immigrant share of each of the skill groups, even in the absence of any shifts of the distribution of skill within the immigrant and native populations.[6]

The trend towards higher levels of skill was more pronounced among natives than among immigrants and is summarized in Figure 1. The high school dropout share of the native population fell by about half, from 23.4 percent to 12.5 percent, compared with a drop from 39.5 percent to 30.4 percent among immigrants. An absolute decrease in the number of dropout natives, coupled with an absolute increase in the number of dropout immigrants, led to a near-doubling of the share of immigrants among dropouts from 10.8 to 20.0 percent, and was somewhat more pronounced among men than among women. A similar, although not as dramatic, change occurred for the immigrant share of high school graduates. The immigrant share of the two college groups also increased, although relative increases in the number of natives in these groups slightly reduced the immigrant share from what it would have been in the absence of shifts in the skill distribution.

3. Changes in the Wages and Quality of Native and Immigrant Skill Groups

It is well–known that real wages for the least-skilled fell *absolutely* during the 1980s and as well as *relatively* compared to their more-skilled counterparts (Bound and Johnson

[5] One potential concern in estimating growth rates of the immigrant population, particularly for low-skill subpopulations, is differential undercount of undocumented aliens between the 1980 and 1990 Censuses. However, Warren and Passel (1987) present evidence that approximately 2 million undocumented aliens were enumerated in the 1980 Census. Assuming an undercount rate of 25 percent, approximately the same number would have been enumerated in the 1990 Census (Warren 1995). During the late 1980s the Immigration Control and Reform Act (IRCA) allowed many previously illegal aliens to become citizens, but a majority of these entered the U.S. after 1982 (Warren 1995). Although the IRCA reduced the number of undocumented aliens from the level that would have prevailed in its absence (Woodrow and Passel 1990), there is no evidence to suggest that undercount rates of undocumented aliens were wildly different in the 1980 and 1990 Censuses. Therefore, none of the results in this paper are adjusted for undercount of this group.

[6] The immigrant share of the four skill groups would have been 14.2, 6.1, 7.1, and 9.5 percent for dropouts, high school graduates, individuals with some college, and college graduates, respectively, in the absence of changes in the distribution of skill among both immigrants and natives.

1992; Katz and Murphy 1992; Murphy and Welch 1992; Juhn, Murphy, and Pierce 1993). Table 2 documents these changes for the immigrant and native populations, presenting real hourly wages for the 16 nativity × sex × skill groups of Table 1 within the 50 largest metropolitan areas of the U.S.[7] I limit my attention in this section and Section 3 to these 50 metropolitan areas because approximately 80 percent of immigrants (and 50 percent of natives) lived there in both 1980 and 1990, with essentially no change in the mass of population living in these areas during the 1980s.[8] The last column of Table 2 presents changes in the log relative real wage of natives and immigrants. Within all sex × skill groups, immigrants' wages fell by more or rose by less than their native counterparts. For men, the wage gap between immigrants and natives grew by between 4.1 and 7.5 percent while for women the change was between 1.7 and 4.9 percent.[9] This decline in the relative wages of immigrants continues the trend noted by LaLonde and Topel (1991) for the 1970s.

Changes in the demographic composition of these broadly-defined skill groups may account for some of the changes in relative wages.[10] In particular, if there was a net decrease in the age of the immigrant population or a net increase in the number of immigrants from countries whose educational systems confer relatively less human capital valued by the U.S. labor market (for a given level of educational attainment), the average productivity (and hence wage) of the immigrant population within skill category would have declined. To measure the change in average labor productivity within nativity × sex × skill group

[7] The 50 metropolitan areas are chosen based on 1990 population and are listed in the Data Appendix. They are defined to be geographically consistent in the 1980 and 1990 data. Jaeger (1995a) discusses the methodology to match geographies. The hourly wage of each group is defined as the real total wage, salary, and self-employment income for the group divided by the total hours worked for the group, where the total hours for each individual is the number of weeks worked × the usual number of hours worked per week. Throughout the paper, 1990 nominal wages are converted to 1979$ using the personal consumption deflator. See the Data Appendix for more information.

[8] Limiting the sample to these 50 metropolian areas is driven, in part, by data requirements for the regression analysis that follows. It also abstracts any possible differences in wage trends due to differences in the urban/rural composition of the native and immigrant populations.

[9] Throughout the paper I interpret 100 × change in logs as percentage changes.

[10] The previous literature uses use the term "quality" to refer to the average relative wage value in the U.S. labor market of particular demographic characteristics, such as age, race/ethnicity, country of origin, or time spent in the U.S. These differences may, of course, arise from factors other than efficiency in production, such as discrimination.

I estimated a log wage equation for each of the 16 groups using 1980 data and then used the coefficients from those regressions with average demographic characteristics to predict the average log wage within each group in both 1980 and 1990. The difference between 1990 and 1980 in the mean log predicted wage provides an estimate of how compositional changes within groups would have changed wages if the prices of "skills" (e.g. experience, race/ethnicity, country of origin, etc.) within group had been held constant. Each regression includes a fourth order polynomial in (age - 40) as a proxy for labor market experience and controls for race/ethnicity.[11] For high school dropouts, the regressions include a dummy variable for having 8 or fewer years of education while those for college graduates include a dummy variable for having more than a college education.[12] The regressions for women include a second-order polynomial in number of children born to control for time possibly spent out of the labor market.

Borjas (1985, 1987, 1992) has extensively documented that individuals from different countries have different earnings potentials in the U.S., even conditional on other observable characteristics. To control for these differences, the regressions for immigrants include broadly-defined region or country of origin.[13] It has also been well-documented that immigrants' arrival date in the U.S. has an impact on wages. Whether this is due to assimilation (Chiswick 1978) or cohort quality (Borjas 1985, 1995a) is still open to some debate (LaLonde and Topel 1991). Regardless, individuals who have been in the U.S. longer tend to be more productive in the U.S. labor market. I control for this by including

[11] The race/ethncitity categories are: white non-Hispanic, black non-Hispanic, Hispanic, other non-white non-Hispanic.

[12] The coding of the 8 or fewer years dummy variable is straightforward in both Censuses. The more than college indicator is 1 for individuals with 18 or more completed years of education in the 1980 Census and as 1 for individuals with a Master's degree or better in the 1990 Census. Jaeger (1995b) briefly describes the rationale for this coding.

[13] The eight countries/regions employed are: Canada, North America other than Canada, Mexico, Latin America other than Mexico, Europe, Asia, Africa, and Oceania.

measures for time spent in the U.S.[14] Each immigrant regression also includes measures of the individual's ability to speak English.[15]

The total change in average productivity, $d\ell_j$ for nativity × sex × skill group j is

$$d\ell_j = \sum_{k=1}^{L_j} \hat{\gamma}_{jk}(\bar{\ell}_{jk90} - \bar{\ell}_{jk80}) \tag{1}$$

where L_j is the number of skill factors for group j, $\hat{\gamma}_{jk}$ is the estimated coefficient on characteristic k from the 1980–based log wage regressions for nativity × sex × skill group j, and $\bar{\ell}_{jkt}$ is the mean value of characteristic k in nativity × sex × skill group j at time t. The total change can be decomposed into subtotals for age, race/ethnicity, country of origin, years in the U.S., ability to speak English, and number of children to measure the sources of changes in labor productivity over the decade. Table 3 presents the results of this decomposition.[16]

The average productivity of dropout and high school graduate immigrants fell during the 1980s among both men and women. This is largely due to changes in the country of origin of the stock of immigrants. The share of Mexican immigrants among dropouts increased by 14 percentage points for men and 10 percentage points for women while the share of European immigrants fell by 18 percentage points for men and 15 percentage points for women. Similar changes took place among high school graduates, where the share of individuals born in Mexico increased by 8 percentage points for men and 5 percentage points for women. At the same time, the share of those born in Europe decreased by 19 percent for men and 14 percent for women. These changes, combined with increases in the number of low-skill immigrants who do not speak English well or at all, lowered the

[14] The six categories are: 0–5 years, 6–10 years, 11–15 years, 16–20 years, 21–30 years, and more than 30 years.

[15] The four categories are: speaks only English or speaks English very well, speaks English well, speaks English not well, speaks English not at all.

[16] Individuals with less than $1 or greater than $100 (in 1979$) or whose wage and salary income, self-employment income, hours worked, or weeks worked were allocated by the Census Bureau were deleted from the samples used to estimate the regressions. The weights used in calculating the mean values of demographic characteristics are total hours worked in 1980 and the product of total hours worked and PWGT1 (the sample weighting variable) in 1990. Results from the productivity-adjustment regressions and the means of demographic characteristics are available from the author by request.

average productivity of immigrants in the two lower-skill groups. The mean age of the male immigrant dropout and high school graduate labor force also dropped substantially (by 2.8 and 1.9 years, respectively), but because the returns to experience are low for these groups this change did not greatly affect the overall quality of these two groups.[17] Similar effects from changes in the mix of country of origin, English language ability, and race/ethnicity obtained for male immigrant individuals with some college, but changes in their age distribution somewhat compensated for those effects. Changes in time spent in the U.S. and age accounted for most of the change in productivity for male immigrant college graduates. For female immigrants with some college, a 1.2 year increase in average age was largely responsible for changes in productivity, while for college graduates, a shift towards longer time in the U.S. played the dominant role in increasing productivity.

For natives, small changes in the race/ethnicity composition had little effect on average productivity. Changes in age composition of the native labor force had a larger effect, however. In particular, the three highest skill groups were older in 1990 than in 1980. For men, the average age was .4, 1.5, and 1.1 years for high school graduates, individuals with some college, and college graduates, respectively, while for women the average age was 2.3, 2.3, and 1.8 years older for these three groups. A 6.7 percentage point increase in the share of college graduates with more than a college education also played a role in increasing the average productivity of native women.

The last column of Table 3 presents the relative average productivity change within each sex $\times$ skill group, which is simply the difference between the total native and immigrant changes in average producitivity. Within all sex $\times$ skill groups, the average productivity (as measured by wages in the U.S.) of the immigrant labor force declined by more (or increased by less) than that of the native labor force, echoing the changes in relative wages.

[17] The coefficients on the linear term in (age - 40) are .0058 and .0060 for male immigrant dropouts and high school graduates, respectively. The comparable coefficients for male native dropouts and high school graduates are .0124 and .0085, respectively.

4. Substitution between Natives and Immigrants in Production

Before discussing the effects of changes in the supply of immigrants had on the wages of natives during the 1980s, I must first establish the degree to which immigrants and natives are substitutes for each other in production. While previous research has examined the degree of substitution between natives and immigrants from different cohorts (Grossman 1982) or immigrants from different ethnic groups (Borjas 1987), I focus here on whether immigrants and natives possessing similar skills are substitutes for one another. The potential impact of immigration on low-skill natives will be greatest if immigrants and natives in the same skill level are highly substitutable. In the discussion that follows, I test Borjas, Freeman, and Katz's (1992, 1996) assumption of perfect substitutability between immigrants and natives within sex $\times$ skill groups.

Analytical Framework

Throughout, I assume that physical capital, K, is weakly separable from labor, is supplied perfectly elastically, and is mobile within the U.S. and between the U.S. and the rest of the world.[18] That is, I assume that the real interest rate is constant, and that capital supplies adjust in response to changes in the labor market, keeping the capital market in equilibrium. I also assume that the 8 sex $\times$ skill groups identified above are weakly separable from one another, leading to an aggregate production technology characterized by a nested function of the form:

$$\begin{gathered} Q = f(g_{md}(X_{nmd}, X_{imd}), g_{mh}(X_{nmh}, X_{imh}), g_{ms}(X_{nms}, X_{ims}), g_{mc}(X_{nmc}, X_{imc}), \\ g_{fd}(X_{nfd}, X_{ifd}), g_{fh}(X_{nfh}, X_{ifh}), g_{fs}(X_{nfs}, X_{ifs}), g_{fc}(X_{nfc}, X_{ifc}), K) \end{gathered}, \quad (2)$$

where Q is output, X_{ijk} is the quantity of labor supplied by nativity $\times$ sex $\times$ skill group ijk (i = {*n*ative, *i*mmigrant}; j = {*m*ale, *f*emale}; k = {high school *d*ropout, *h*igh school graduate, *s*ome college, *c*ollege graduate}); and g_{jk} is a function that aggregates immigrant and native labor in sex $\times$ skill group jk. The separability assumption implies that changes in the supply of labor groups affect each other only through the labor aggregation functions

[18] The assumption of separable capital and labor inputs is driven largely by a lack of data for capital.

g_{jk}. For example, I assume that changes in the supply of dropout males affect the wage of college graduate females only through the relationship between g_{md} and g_{fc}. This specification allows me to examine the relationship between natives and immigrants within sex $\times$ skill groups independently of the other groups and capital.

Under the additional assumption that the g_{jk} are linear homogeneous, the gross elasticity of substitution, σ_{jk}, between immigrants and natives along the jk isolabor curve (Hamermesh 1993, pp. 67–68) is

$$\sigma_{jk} = -\frac{d\log(X_{ijk}/X_{njk})}{d\log(W_{ijk}/W_{njk})}. \tag{3}$$

A simple estimate of the elasticity of substitution between immigrants and natives can be obtained by using the changes in relative wages and relative labor supply during the 1980s. This estimator is unbiased if the *only* factor influencing the change in relative wages between immigrants and natives is the change in relative supplies.

Naive Estimates of the Elasticities of Substitution

Table 4 presents changes in relative labor supply, changes in relative wages, and the elasticity of substitution that those changes imply for the 8 sex $\times$ skill groups. The first two columns present changes in the log relative quantity of immigrant and native labor, measured in hours, both unadjusted and adjusted for changes in relative average productivity.[19] The third and fourth columns present unadjusted and average productivity-adjusted changes in the log of the relative wages of immigrants and natives, where wages are measured as in Table 2.[20] The last two columns contain estimates of σ_{jk} implied by the actual and adjusted changes in relative wages and relative supplies, as given by equation (3).

[19] Labor quantities are measured as the sum of individual annual hours for each group. The productivity adjustment adds the change in the relative average productivity of immigrants (the last column of Table 3) to the change in relative hours. For example, the change in log quality-adjusted relative hours for high school graduate men is .516 (the change in log relative actual hours) + ($-.091$) (the change in relative average productivity of immigrants) = .425.

[20] The productivity adjustment for wages is similar to that for hours, except that the adjustment is substracted from the change in relative wages. For example, the change in log productivity-adjusted relative wages for high school graduate men is $-.048$ (the change in log relative real wages) - ($-.091$) (the change in relative average productivity of immigrants) = .043.

Without exception, the estimated σ_{jk} using actual quantities and wages are large, implying that immigrants and natives are close substitutes. The productivity-adjusted estimates of σ_{jk} are generally larger, but because changes in average productivity more than compensated for the change in relative wages, the estimates for both male and female high school graduates and those with some college are negative. Still, the changes in relative wages are small compared to the changes in relative hours. The magnitude of these relatively crude estimates gives some indication that immigrants and natives are relatively close substitutes at all skill levels for both sexes.

Regression Estimates of the Elasticities of Substitution

The naive estimates of the elasticity of substitution in Table 4 are likely to be biased because factors other than changes in relative supplies affected the change in the relative wages of immigrants and natives during the 1980s. In particular, shifts in relative demand or changes in relative average productivity not captured by the adjustment procedure might have pushed relative wages in either direction. This section presents regression estimates of the within–sex $\times$ skill group elasticity of substitution between immigrants and and natives.

As in the usual simultaneous equations framework in which only changes in relative quantities, $h = d\log(H_{ijk}/H_{njk})$, and changes in relative wages, $w = d\log(W_{ijk}/W_{njk})$, are observed, the relative and supply and demand elasticities are underidentified without additional information about the parameters themselves and/or the variance-covariance matrix of supply and demand shocks. Without this additional information, the coefficient on h from a regression of w on h, gives an estimated parameter which is a weighted average of the relative supply and relative demand elasticities (Maddala 1977, p. 244). Ordinary least square (OLS) will consistently estimate the relative demand elasticity (1) if the relative demand curve is stable, i.e. if the variance of relative demand shocks is equal to zero, or (2) if the relative supply curve is perfectly inelastic and supply and demand shocks are uncorrelated. That is, the variation in observed quantities, h, must come from solely from supply shocks rather than from demand shocks.

Even if relative quantity changes are orthogonal to relative demand shocks, OLS regression of w on h may lead to biased estimates of σ_{jk}. Because observed hours would be used both as a regressor and to calculate hourly wages, the definitional relationship between the right and left sides of the regression of w on h might introduce bias in the estimated parameters. While this issue is likely to be more problematic with individual rather than aggregate data, the bias might be quantitatively important for estimates in which both changes in relative supplies and changes in relative wages were adjusted for changes in relative quality. To obviate the problem, I replace w with $e = d\log(E_{ijk}/E_{njk})$ and instead estimate

$$e = \alpha + \beta h + \epsilon, \tag{4}$$

where E_{ijk} is the earnings of nativity group i and ϵ is an error term. Then (dropping the jk subscripts):

$$\sigma = \frac{1}{1-\beta} = \frac{1}{c}, \tag{5}$$

where c is the within–sex $\times$ skill elasticity of complementarity between immigrants and natives. Hypothesis tests can be performed on $\hat{c} = 1 - \hat{\beta}$. In particular, the test of the null hypothesis that $c = 0$ is essentially a test that $\sigma = \infty$.

Like much of the previous literature, I use geographic differences in the concentration of immigrants to identify regression estimates of the elasticity of substitution between immigrants and natives. There is substantial variation in both the share of immigrants in the labor force as well as relative growth rates of immigrant and native immigrant labor across the 50 largest metropolitan areas used in my analysis. The change in log total relative quantities, $d\log(H_i/H_n)$, measured in actual hours, ranged from -.146 in Buffalo to .778 in Atlanta; the standard deviation across the 50 metropolitan ares is .231. While there are some differences, this degree of variability across metropolitan areas is representative of the 8 sex $\times$ skill groups.

Recently, Friedberg and Hunt (1995) and Borjas, Freeman, and Katz (1996) have criticized so-called "area analyses" which use variation in the geographic concentration of immigrants to identify the impact of immigration. In particular, they point to evidence presented by Filer (1992) and Frey (1995) that natives, particularly those with low skills, may migrate in response to inflows of immigrants to particular labor markets. This criticism is possibily quite important for the OLS results of Grossman (1982), Borjas (1987), LaLonde and Topel (1991), Altonji and Card (1991), and Schoeni (1996) in which the level of natives' labor market outcomes are regressed on immigrant concentrations. In those studies, immigration and internal migration of natives work in opposite directions on natives' labor market outcomes, biasing the results. The framework of equation (4) does not suffer from the same biases. however. Except in the unlikely case that natives migrate in response to changes in their wages *relative* to similarly-skilled immigrants, native migration does not induce a correlation between ϵ and h. A similar argument can be made regarding the geographic location choice of immigrants.[21] Note too, that the problem of growth in immigrant share being correlated with price levels recently noted by Schoeni (1996) is not an issue when using changes in relative wages (or earnings) as the dependent variable. Man of the endogeneities that plague the previous literature are therefore unlikely to induce much, if any, bias in the estimates from equation (4).

Estimates of c and σ for the 8 sex $\times$ skill groups using OLS and changes in actual relative labor quantities are presented in columns 1 and 2 of Table 5. Heteroskedasticity-consistent standard errors estimated using the jackknife (Efron 1982) are in parentheses.[22] I assume that the ϵ (and hence demand shocks) are orthogonal to relative quantity changes, h; i.e. I assume that the relative demand elasticity is identified. For men, the estimates of c are significantly different from zero at the 5 percent level for dropouts and high school graduates. The estimated σ for all four groups are relatively large, however, indicating that native and immigrant men are likely to be close substitutes in production. For women, $\hat{c}$ is statistically significant at the 1 percent level for high school graduates, but is large

[21] While literature on immigrant location choice is somewhat inconclusive about the responsiveness of immigrants to economic conditions (see Bartel 1992), Borjas (1987), Altonji and Card (1991), and Schoeni (1996) employ a variety of instruments to address this problem. One of the instruments that Borjas (1987) employs in his IV estimates if the proportion of the labor force in one-digit industry groups. The validity of this instrument rests on the implausible assumption taht capital is immobile and cannot relocate in response to changes in supplies.

[22] MacKinnon and White (1985) show that the small sample performance of jackknife standard errors in OLS is superior to other heteroskedasticity-consistent standard errors such as those suggested by Hinkley (1977) or White (1980).

enough to suggest a high degree of substitutability. The estimates of c for women with some college is perversely signed but are not significantly different from zero, also implying a high degree of substitutability between immigrant and native women within skill groups.

Table 3 shows that there were substantial changes in the average relative productivity of immigrants and natives, particularly for men. Columns 6 and 6 of Table 5 present estimates of the elasticities of substitution for the 8 sex $\times$ skill groups in which the change in relative quantities in each metropolitan area has been adjusted for the change in relative average productivity. That is, the changes in productivity-adjusted relative hours, $\tilde{h}$, are substituted for h in estimating equation (4).[23] As with the naive estimates, adjusting for changes in productivity generally increases the etimates of σ, strenthening the conclusion that immigrants and natives are close substitutes.

The OLS estimates of β used to estimate c and σ in Table 5 are almost surely inconsistent, however, due to measurement errors in h. Unlike the usual errors-in-variables problem, however, the same measurement errors are also present in e. Within each sex $\times$ skill group, the observed sample may over- or underestimate the change in the relative sizes of the immigrant and native populations. the change in relative hours, or the change in relative wages. Errors in measuring relative size of the immigrant and native labor force and the in relative hours affect both e and h while measurement error in wages affects only e. In Appendix A, I show that OLS estimates of β will be biased towards 1, in turn biasing the estimates of σ towards higher degrees of substitution.

One natural solution to this errors-in-variables problem is to use a population measure of the right hand side variable rather than sample estimates. Ideally, I would like use the population change in relative hours to instrument for h, because it would be free from measurement error both in the change in the relative size of the immigrant population and in changes in relative hours. While it is not possible to obtain this instrument, a close substitute, the population change in the log relative number of immigrants, i.e. $d\log(N_i/N_n)$, can be constructed for each metropolitan area from the STF3C summary files for both the 1980 and 1990 Censuses. This is a valid instrument for h as long as any

[23] The productivity adjustments for each metropolitan area are analogous to those for the agggregate and are calculated by using within-metropolitan area means in equation (1), giving $\tilde{h} = h + (d\ell_i - d\ell n)$.

mismeasurement in the way this quantity approximates the population change in hours is uncorrelated with any of the measurement errors present in h or e.

Estimates of c and σ using instrumental variables (IV) are presented in columns 3 and 4 (instrumenting for h) and columns 8 and 9 (instrumenting $\tilde{h}$) of Table 5. For men, the IV estimates of c for high school dropouts graduates are close to the OLS estimates (although larger, as expected) and the estimates of σ remain large enough to continue to suggest that immigrants and natives are close to perfect substitutes. As with OLS, the estimates of c for male individuals with some college are not statistically significantly different from zero. The estimates of c for college graduates are statistically significantly different from zero using IV, giving estimates of σ close to 4. For women, two of the elasticities switch signs using actual quantities while one switches sign with adjusted quantities. The estimates of c are statistically significantly different from zero only for high school graduates, however, with the corresponding estimates of σ still large enough to suggest a high degree of substitutability. Correcting for errors-in-variables bias does not alter the conclusion that the relative demand curve within sex $\times$ skill groups is essentially perfectly elastic.[24]

Because the measurement errors in e and h are essentially sampling errors, it is possible to derive estimators for their variances based on observable quantities and correct the estimates of β for errors-in-variables bias. In Appendix A, I derive the corrected estimator for β and present corrected estimates of c and σ. They are roughly comparable to those estimated using IV and do not change the conclusion based on the OLS and IV estimates that immigrants and natives are close to perfect substitutes. That the measurement errors in h are sampling errors also suggests that weighted regressions would be more efficient, since the sampling errors induce heteroskedasticity in ϵ that is approximately proportional to the sum of the inverse of the sample size for the four samples used to create e and h. I estimated both the OLS and IV models from Table 5 weighted by $(1/N_{i90} + 1/N_{n90} + 1/N_{i80} + 1/N_{n80})^{-.5}$ (supressing the sex $\times$ skill subscripts). The results were very similar

[24] Bound, Jaeger, and Baker (1995) show that the inverse of the F statistic on the instrument in the first stage regression is approximately equal to the finite sample bias of IV relative to OLS. The F statistics of the instrument in the regressions h or $\tilde{h}$ on $d\log(N_i/N_n)$ ranged from 21.8 to 109.6, indicating that finite sample biases are not an issue in these estimates.

to those in Table 5 and in no way changed the conclusion that immigrants and natives are close substitutes.

The preceding analysis has rested somewhat delicately on the assumption that the conditions for identifying the relative demand elasticity in equation (4) are met. While it is impossible to know the exact magnitude of the simultaneity bias, there are several reasons to expect that it is not large. First, much of the equation error in (4) appears to be due to sampling errors in h and e. I calculated the approximate variances of demand shocks for each of the OLS models in Table 5 using estimates of the variances of the various measurement error components in h and e and found that they were all small relative to the mean squared errors from those regressions. This suggests that a sufficient condition for identification of the relative demand elasticity, namely that the variance of demand shocks is zero, may be approximately met. Second, I estimated very weak correlations between changes in relative quantities and approximate changes in relative demand shocks.[25] Even if the correlations between actual changes in relative demand shocks and observed changes in relative quantities were substantially stronger than the estimated correlations between demand proxies and h, the effect on the estimated elasticicies would be quite small. It seems unlikely, therefore, that simultaneity biases in the estimates of Table 5 are large.

5. The Effect of Immigration on the Wages of Natives

The finding that immigrants and natives are approximately perfect substitutes within sex $\times$ skill groups indicates that the large changes in the supply of immigrants during the 1980s may have had a significant effect on changes in natives' wage levels and on the relative wages of low– and high–skill workers. In particular, the substantial change in the immigrant share of the least-skilled work force indicates that immigration may have played a role in depressing the absolute and relative wages of low–skilled natives. To

[25] The demand index used is the "manpower requirements" index first formulated by Freeman (1975) and subsequently used by Murphy and Welch (1991), Bound and Johnson (1992), Katz and Murphy (1992), among others. It measures the change in relative demand within each education group in a metropolitan area as a weighted sum changes in national industry shares within education and sex groups, with the weights being the share of the sex $\times$ education $\times$ nativity groups in the industry and metropolitan area in 1990. Including these proxies for labor demand in the OLS and IV regressions did not qualitatively change the estimates of the coefficients and therefore the estimates of the elasticities of substitution.

examine the effect of immigrants on changes in wage levels and relative wages, I employ a simple aggregate production function framework. Within this framework, the effect of immigration on wages follows directly from the effects on relative quantities. Empirical estimates in this section are not limited to the 50 largest metropolitan areas, but are for the entire contiguous U.S.

While immigrants are, on average, less skilled than natives, Figure 1 demonstrates that they increase the supply of labor at *both* ends of the skill distribution. Because immigrants had the largest (percentage) impact on the supply of high school dropouts during the 1980s, a production function framework that allows for differential effects on dropouts and high school graduates may better capture the effects of immigrants, particularly for native dropouts. To that end, let the aggregate production functon now be given by

$$Q = F(G(L(D, H), C), K). \tag{6}$$

I assume that the production function contains three labor groups: high school dropouts (D), high school graduates (H), and an aggregate of those who attended some college or more (C). The function is nested such that dropouts and high school graduates comprise a low-skill aggregate, $L(D, H)$, which is part of the labor aggregate $G(L(D, H), C)$. Given the results of the previous section, I assume that the relative demand curve for immigrants and natives within skill groups is perfectly elastic. While those results may be biased towards finding high degrees of substitution, it is worth noting that in this framework the maximal estimate of impact of immigration occurs when immigrants and natives within skill groups are perfect substitutes. The results in this section should therefore be interpreted as upper bounds on the effects of immigration. To simplify the analysis, I also assume that men and women within skill groups are perfect substitutes and continue to assume that capital is weakly separable from the labor aggregate.

Under the additional assumptions (1) capital is mobile and the capital market is in long-run equilibrium (i.e. K is supplied perfectly elastically and adjusts to keep K/G constant, implying that F_G is constant) and (2) G and L exhibit constant elasticities of substitution, r and s, respectively, it is easy to show that the change in the log wage level of native dropouts due to changes in the relative supply of immigrants is

$$\begin{aligned} d\log W_{nd} =& -\left(s_d\frac{1-s_l}{r}+\frac{1-s_d}{s}\right)d\log(D/D_n) \\ & -\left((1-s_d)\frac{1-s_l}{r}-\frac{1-s_d}{s}\right)d\log(H/H_n) \\ & +\frac{1-s_l}{r}d\log(C/C_n) \end{aligned} \tag{7}$$

where s_l is the share of the low-skill aggregate in G and s_d is the share of dropouts in L. Similarly,

$$\begin{aligned} d\log W_{nh} =& -\left(s_d\frac{1-s_l}{r}-\frac{s_d}{s}\right)d\log(D/D_n) \\ & -\left((1-s_d)\frac{1-s_l}{r}+\frac{s_d}{s}\right)d\log(H/H_n)\,, \\ & +\frac{1-s_l}{r}d\log(C/C_n) \end{aligned} \tag{8}$$

and

$$\begin{aligned} d\log W_{nc} =& \frac{s_l s_d}{r}d\log(D/D_n) \\ & +\frac{s_l(1-s_d)}{r}d\log(H/H_n)\,. \\ & -\frac{s_l}{r}d\log(C/C_n) \end{aligned} \tag{9}$$

Equation (7) shows that an increase in the relative supply of immigrant dropouts lowers the wage of native dropouts in two ways. First, it shifts out the relative supply curve of dropouts within the low–skill aggregate L (the substitution effect). Second, it shifts out the relative supply curve for low–skill labor within G (the scale effect). An increase in the relative supply of immigrant dropouts has an unambiguously negative effect on native dropout wages. An increase in the relative supply of immigrant high school graduates has an effect of indeterminate sign, however, while an increase in the relative supply of immigrants in the the high–skill aggregate unambiguously raises native dropout wages. The effects on high school graduates and the college aggregate are analogous to those for dropouts.

The methodology developed in Section 3 adjusts labor supplies for changes in quality *within* nativity × sex × skill groups across time, but does not address differences in productivity *between* groups. To convert productivity-adjusted hours into efficiency units of labor, I calculated the mean predicted wage for each of the 16 nativity × sex × groups using the coefficients employed in calculating the quality adjustments with a fixed set of characteristics for all 16 groups.[26] The relative efficiency of each group, π_j, is then given by

$$\pi_j = \frac{\exp(\widehat{\log w_j})}{\exp\left(\sum_j \psi_j \widehat{\log w_j}\right)} \tag{10}$$

where $\widehat{\log w_j}$ is the fixed-characteristic predicted log wage for sex × nativity group j and ψ_j is the employment share of group j in 1980. The supply of labor by group j in efficiency units in 1980 is then given by $\pi_j H_{j80}$. To calculate the supply of labor in efficiency units in 1990, the actual hours must first be adjusted for changes in quality (i.e. calculating the equivalent number of 1980 hours) and then converted to efficiency units: $\pi_j \exp(\log H_{j90} + d\ell_j)$.

Estimates of immigrant and native supplies of D, H, and C in 1980 and 1990 for the contiguous U.S. are presented in Table 6. The supplies have been normalized to the total labor supply within each year, so that entries in the table represent shares in total labor supply within year.[27] The effect of immigration on the supply of each labor group is presented in the third row for 1980 and the sixth row for 1990, while the effect of immigration on the change in the supply of each group is presented in the seventh row. Table 6 shows that, within year, immigration increases the relative supply of dropouts substantially more than the two higher–skill groups. It is also clear that the share of dropout natives declined precipitously during the decade and that immigration worked to mute the effects of that decline on the overall distribution of skill in the workforce.

[26] Those characteristics are: age 40 and the mean levels of race/ethnicity for the entire sample. Wages for immigrants are evaluated at the mean level of their characteristics (country of origin, ability to speak English, and time in the U.S.) of all immigrants while number of children is evaluated at the mean level for all women.

[27] In calculating shares, I have weighted the some college supply by .3 within the college aggregate and the high school dropout supply by .9 within the low–skill aggregate. That is, I approximately follow Katz and Murphy's (1992) weighting scheme in creating high– and low–skill aggregates, with the exception that I do not include individuals with some college in the low–skill group.

Effect of Immigration on Wage Levels

Most of the literature on the effect of immigration on the wages of natives has examined how immigrants affect wage levels. While Grossman (1982) and Borjas (1987) do not specifically estimate elasticities of the wages of low-skill native workers with respect to changes in the supplies of immigrants, they both estimate the elasticity of native wages with respect to changes in the supplies of immigrants to be approximately -.02 to -.03. Given a change in the aggregate log supply of immigrants of .44 (see Table 1) during the 1980s, their results predict that, *ceteris paribus*, the average wage level of natives would have fallen by .8 to 1.3 percent during the decade. In contrast, Altonji and Card's (1991) first-differenced results predict that the increase in the immigrant share of the labor force from 6.7 percent to 8.9 percent (see Table 1) would have an effect on the wages of low-skill natives ranging from an increase of .7 percent to a decrease of 4.2 percent, depending on the estimation method and the particular low-skill native subsample. The predicted effect using the average of their OLS estimates for the four low-skill native subsamples is a decrease of .9 percent while the predicted effect using their mean IV estimates is a decrease of 1.9 percent. Their estimates are rather imprecisely estimated, however, and in all cases except one, they cannot reject the hypothesis that changes in immigrant share of the labor force have no effect on the log earnings of low-skill native subgroups.

Given the model in equation (6), however, it is clear that the effects of immigration will differ across native skill groups and that changes in the supplies of immigrants will have different effects as the relative skill distributions of immigrants and natives shift. Estimates of the effect of changes in supply of immigrants on the log wage level of each of the three labor groups during the 1980s are presented in Table 7 for various values of r and s.[28] The values of r are commonly-used values for the elasticity of substitution between high– and low–skill workers, while those for s are values indicative of a high degree of substitution

[28] Shares are the mean share between 1980 and 1990 from Table 6. Wage changes are average share-weighted changes in native log real wages, using average shares by sex within skill level. For example, the average male share among native dropouts was .656. Productivity-adjusted changes in log real wages for native dropout men were $-.123$ (change in log real wages) $-.004$ (change in average productivity) $= -.127$ and $-.060 - .006 = -.053$ for native dropout women. This gives a change in log wages for native dropouts of $.656 \times -.123 + (1 - .656) \times -.053 = -.102$. Calculations for high school graduates and the college aggregate are analogous, with those with some college being weighted by .3, as above.

between dropouts and high school graduates. The rows with a value of infinity for s are the effects under the assumption that dropouts and high school graduates are perfect subsitutes.

Most of the impact of immigration on natives' wages derives from the damping effect immigration had on the sharp reduction in the supply of dropout labor to the U.S. economy. For dropouts, immigration can account for approximately 1.5 to 3 percentage points of a 10 percent point decline in real wages, or approximately 15 to 30 percent of the change over the decade. Nearly all of the effects of immigration on dropouts come from relative changes in the supplies of immigrant and native dropouts. The impact on high school graduates is substantially smaller, on the order of a reduction of 1 percent, or about 11 percent of the total decline, and is spread somewhat evenly among the three skill groups. For college equivalents, the effects of immigration account for approximately one quarter of the increase in wage levels. The impact of changes in supplies of high school graduates and those with some college are approximately equal and of opposite sign, leaving dropouts to account for the majority of the change.

Immigration would appear to haved played a large role in the decline of the wages of the least–skilled natives, but played a smaller role for high school graduates and actually increased the wage levels of college equivalent workers. These magnitudes are somewhat variable depending on the assumptions made about the degree of substitutability between dropouts and high school graduates, particularly for the effects on dropouts. Nevertheless, the results in Table 7 indicate that, with plausible economic assumptions, the wages of dropouts would have been substantially higher if the distribution of skill had changed in the same way for immigrants as it did for natives during the 1980s.

Effect of Immigration on Relative Wages

Much attention has recently been focused on the growth in the difference in wages between high– and low–skill workers during the 1980s. Borjas, Freeman, and Katz (1992, 1996) explicitly examined the role of immigration and trade on the increasing wage gap and found that immigration explained little of the increase between high school equivalents and college equivalents, but could explain much more of the growth in the gap between high school dropouts and all other workers. Equation (6) is a more general form of their two–factor model. The model used for their first set of results, in which labor is aggregated into high school equivalents and college equivalents, is a special case of equation (6) (with D and H being perfect substitutes). They rely on the somewhat unrealistic assumption that all workers with a high school degree or better are perfect substitutes in their second set of results examining the drop in the relative wages of high school dropouts.

Under the assumption that dropouts and high school graduates are perfect substitutes and that the elasticity of substitution between high– and low–skill workers is 1.5, immigration can explain approximately 2 percentage points of a 12.4 percentage point increase (or about 16 percent) in the growth of the wage disparity between high school and college workers.[29] This result is consistent with those of Borjas, Freeman, and Katz (1992, 1996). Allowing for less-than-perfect substitution between high school dropouts and high school graduates, and using $r = 1.5$ and $s = 6$, I find that immigration explains about 2.9 percentage points of the 13.4 percentage point increase (22 percent) in the native dropout – college differential, but only 1.6 percentage points of a 12.0 percentage point increase (13 percent) in the native high school – college premium. Note, too, that immigration more than accounts for the change in the dropout – high school graduate wage differential for any value of s shown in Table 7.

6. Conclusion

During the 1980s there were large changes in the relative quantity of immigrant and native labor within sex × skill groups. I found, however, that the relative wages of immigrants and natives within those groups changed little. Using an aggregate production function in which 8 sex × skill groups are weakly separable from one another, I estimated that immigrants and natives within those groups are essentially perfect substitutes. This finding was consistent across naive and regression estimates of the relevant elasticities, as well as estimates corrected for measurement error in changes in relative supplies.

[29] The change in the wage of the high school aggregate is the weighted average of the change in the log wage of dropouts and high school graduates, using average employment shares, both as shown in Table 7. That is, the change is equal to $.310 \times -.102 + .690 \times -.088 = -.092$.

While my results are somewhat sensitive to the choice of substitution parameters, I find that immigration accounts for approximately 15 to 25 percent of the increase in the wage gap between low– and high–skill workers during the 1980s. The impact on native high school dropouts was even more substantial, with immigration accounting for as much as 3 percentage points (roughly one-third) of the decline in their real wages. The effects on the wage levels of other skill groups were comparatively smaller. These effects are unlikely to be uniformly distributed across the U.S., however. Immigrants tend to locate in a relatively few metropolitan areas, and their labor market effects will be concentrated in those areas. Future research will examine the effects of immigration in specific geographic areas.

Appendix A: Errors-in-Variables Bias With Common Error Components

The estimating equation under consideration is

$$e_c^* = \alpha + \beta h_c^* + \epsilon_c^* \tag{A.1}$$

where e_c^* and h_c^* are the true values of changes in relative earnings and hours, respectively, and ϵ_c^* is an error term. The OLS estimator of β is possibly inconsistent due to measurement error h_c, the observed values of changes in log relative relative hours. For purposes of exposition I will assume that the h_c^* are orthogonal to any shocks to demand. That is, I examine the effects of mismeasurement in e_c and h_c on the consistency of $\hat{\beta}$ in isolation of identification issues. I will employ metropolitan area subscripts c in this Appendix to make clear the heteroskedastic nature of the error components.

There are two possible sources of measurement error which affect both e_c and h_c and one which affects only e_c. First, the sample may over- or underestimate the number of individuals in the particular metropolitan area $\times$ nativity $\times$ sex $\times$ skill cell in either year. I represent this error by η_{jtc}, which is present in both e_c and h_c. Second, the sample of individuals in the metropolitan area $\times$ nativity $\times$ sex $\times$ skill cell may have worked more or less hours on average than the population in that cell. This sampling error is represented by ν_{jtc} and is is also present in observed earnings because earnings are assumed to be the product of hours and wages. Lastly, there is sampling error component in wages, represented by ω_{jtc}. The dependent and independent variables in equation (4) are each constructed from four components:

$$e_c = e_{i90c} - e_{n90c} - e_{i80c} + e_{n80c} \tag{A.2}$$

and

$$h_c = h_{i90c} - h_{n90c} - h_{i80c} + h_{n80c} \tag{A.3}$$

where $e_{jtc} = \log(E_{jtc})$, $h_{jtc} = \log(H_{jtc})$, i indicates immigrants and n indicates natives. I assume that each of the components of e_c and h_c are measured with errors of the form described above, that is

$$\begin{aligned} e_{jtc} &= e_{jtc}^* + \eta_{jtc} + \nu_{jtc} + \omega_{jtc} \\ h_{jtc} &= h_{jtc}^* + \eta_{jtc} + \nu_{jtc} \end{aligned}. \tag{A.4}$$

Substituting equations (A.4) into equation (A.1) gives

$$e_c = \alpha + \beta h_c + (1-\beta)(\eta_c + \nu_c) + \omega_c + \epsilon_c^* \tag{A.5}$$

where

$$\eta_c = \eta_{i90c} - \eta_{n90c} - \eta_{i80c} + \eta_{n80c}, \tag{A.6}$$

$$\nu_c = \nu_{i90c} - \nu_{n90c} - \nu_{i80c} + \nu_{n80c}, \tag{A.7}$$

and

$$\omega_c = \omega_{i90c} - \omega_{n90c} - \omega_{i80c} + \omega_{n80c}. \tag{A.8}$$

I assume that η_{jtc}, ν_{jtc}, ω_{jtc}, and ϵ^*_c have zero mean and that all are independent of one another except ν_{jtc} and ω_{jtc}. I assume that $\sigma_{\nu\omega_{jtc}} > 0$, i.e. if the sample of individuals has unusually high (or low) hours they are also more likely to have unusually high (or low) wages. I also assume that all of the error components are temporally and spatially uncorrelated with themselves and with each other. That is,

$$\begin{aligned}
\mathrm{E}\ \eta_{jtc} = \mathrm{E}\ \nu_{jtc} = \mathrm{E}\ \omega_{jtc} = \mathrm{E}\ \epsilon^*_c = 0 \qquad & \forall\ j, t, \text{ and } c \\
\mathrm{E}\ \eta_{jtc}\nu_{k\tau d} = \mathrm{E}\ \eta_{jtc}\omega_{k\tau d} = \mathrm{E}\ \eta_{jtc}\epsilon^*_d = \mathrm{E}\ \nu_{jtc}\epsilon^*_d = 0 \qquad & \forall\ j, k, t, \tau, c, \text{ and } d \\
\mathrm{E}\ \eta_{jtc}\eta_{k\tau d} = \mathrm{E}\ \nu_{jtc}\nu_{k\tau d} = \mathrm{E}\ \omega_{jtc}\omega_{k\tau d} = 0 \qquad & \forall\ j \neq k \text{ or } t \neq \tau \text{ or } c \neq d \\
\mathrm{E}\ \eta^2_{jtc} = \sigma^2_{\eta_{jtc}} & \\
\mathrm{E}\ \nu^2_{jtc} = \sigma^2_{\nu_{jtc}} & \\
\mathrm{E}\ \epsilon^*_c = \sigma^2_{\epsilon^*_c} & \\
\mathrm{E}\ \nu_{jtc}\omega_{jtc} = \sigma_{\nu\omega_{jtc}} > 0 &
\end{aligned} \quad . \tag{A.9}$$

Under assumptions (A.9) the OLS estimator of β is

$$\begin{aligned}
\hat{\beta} &= \frac{\sum \dot{h}_c \dot{e}_c}{\sum \dot{h}^2_c} \\
&= \beta + \frac{\frac{1}{C}\sum(1-\beta)(\dot{h}^*_c + \dot{\eta}_c + \dot{\nu}_c)(\dot{\eta}_c + \dot{\nu}_c) + (\dot{h}^*_c + \dot{\eta}_c + \dot{\nu}_c)(\dot{\omega}_c + \dot{\epsilon}^*_c)}{\frac{1}{C}\sum(\dot{h}^*_c + \dot{\eta}_c + \dot{\nu}_c)},
\end{aligned} \tag{A.10}$$

where $\dot{x} = x - \bar{x}$. Taking the plim of $\hat{\beta}$ yields

$$\begin{aligned}
\text{plim } \hat{\beta} &= \beta + (1-\beta)\frac{\bar{\sigma}^2_{\eta_c} + \bar{\sigma}^2_{\nu_c} + \bar{\sigma}_{\nu\omega_c}}{\sigma^2_h} \\
&= \beta + (1-\beta)\frac{\bar{\sigma}^2_{\eta_c} + \bar{\sigma}^2_{\nu_c} + \bar{\sigma}_{\nu\omega_c}}{\sigma^2_{h^*} + \bar{\sigma}^2_{\eta_c} + \bar{\sigma}^2_{\nu_c}} \\
&= \beta + (1-\beta)\pi
\end{aligned} \tag{A.11}$$

where

$$\begin{aligned}
\bar{\sigma}^2_{\eta_c} &= \bar{\sigma}^2_{\eta_{i90c}} + \bar{\sigma}^2_{\eta_{n90c}} + \bar{\sigma}^2_{\eta_{i80c}} + \bar{\sigma}^2_{\eta_{n80c}}, \\
\bar{\sigma}^2_{\nu_c} &= \bar{\sigma}^2_{\nu_{i90c}} + \bar{\sigma}^2_{\nu_{n90c}} + \bar{\sigma}^2_{\nu_{i80c}} + \bar{\sigma}^2_{\nu_{n80c}}, \\
\bar{\sigma}_{\nu\omega_c} &= \bar{\sigma}_{\nu\omega_{i90c}} + \bar{\sigma}_{\nu\omega_{n90c}} + \bar{\sigma}_{\nu\omega_{i80c}} + \bar{\sigma}_{\nu\omega_{n80c}}, \\
\pi &= \frac{\bar{\sigma}^2_{\eta_c} + \bar{\sigma}^2_{\nu_c} + \bar{\sigma}_{\nu\omega}}{\sigma^2_{h^*} + \bar{\sigma}^2_{\eta_c} + \bar{\sigma}^2_{\nu_c}}, \\
\bar{\sigma}^2_{\eta_{jtc}} &= \frac{1}{C}\sum \sigma^2_{\eta_{jtc}}, \\
\bar{\sigma}^2_{\nu_{jtc}} &= \frac{1}{C}\sum \sigma^2_{\nu_{jtc}}, \\
\bar{\sigma}_{\nu\omega_{jtc}} &= \frac{1}{C}\sum \sigma_{\nu\omega_{jtc}},
\end{aligned} \tag{A.12}$$

and $\sigma^2_{h^*}$ is the population variance of h^*. It is clear that $\hat{\beta}$ is biased if $\bar{\sigma}^2_{\eta_c}+\bar{\sigma}^2_{\nu_c}+\bar{\sigma}_{\nu\omega_c}\neq 0$. Estimates of $\hat{\beta}$ will biased towards 1 with the magnitude of the bias increasing as π, the share of the variance of the observed values h_c, σ^2_h, accounted for by $\bar{\sigma}^2_{\eta_c}$, $\bar{\sigma}^2_{\nu_c}$, and $\bar{\sigma}_{\nu\omega_c}$ increases towards 1. This implies that the estimated elasticity of complementarity between immigrants and natives will be biased towards 0, which in turn implies that the estimated elasticity of substitution will be biased towards infinity.

Imposing some structure on $\bar{\sigma}^2_{\eta_{jtc}}$, $\bar{\sigma}^2_{\nu_{jtc}}$, and $\bar{\sigma}_{\nu\omega_{jtc}}$ permits the construction of a consistent estimator of β. First, a consistent estimator of the variance of

$$\eta_c = \eta_{i90c} + \eta_{n90c} + \eta_{i80c} + \eta_{n80c} \tag{A.13}$$

is

$$\hat{\sigma}^2_{\eta_c} = \frac{1}{N_{i90c}} + \frac{1}{N_{n90c}} + \frac{1}{N_{i80c}} + \frac{1}{N_{n80c}} \tag{A.14}$$

(Fienberg 1977, p. 18). Therefore, a consistent estimator of $\bar{\sigma}^2_{\eta_c}$ is given by

$$\begin{aligned}\widehat{\bar{\sigma}^2_{\eta_c}} &= \frac{1}{C}\sum \hat{\sigma}^2_{\eta_c} \\ &= \frac{1}{C}\sum \frac{1}{N_{i90c}} + \frac{1}{N_{n90c}} + \frac{1}{N_{i80c}} + \frac{1}{N_{n80c}}\end{aligned}. \tag{A.15}$$

Second, note that

$$h_{jtc} = \log(H_{jtc}) = \log(N_{jtc}\bar{H}_{jtc}), \tag{A.16}$$

where H_{jtc} is the total number of hours worked by nativity group j in period t in metropolitan area c. The sampling variation in h_{jtc} is therefore equal to the sampling variation in $\log(\bar{H}_{jtc})$. A consistent estimator of $\bar{\sigma}^2_{\nu}$ therefore is

$$\begin{aligned}\widehat{\bar{\sigma}^2_{\nu}} &= \frac{1}{C}\sum \frac{\hat{\sigma}^2_{H_{i90c}}}{N_{i90c}\bar{H}^2_{i90c}} + \frac{\hat{\sigma}^2_{H_{n90c}}}{N_{n90c}\bar{H}^2_{n90c}} + \frac{\hat{\sigma}^2_{H_{i80c}}}{N_{i80c}\bar{H}^2_{i80c}} + \frac{\hat{\sigma}^2_{H_{n90c}}}{N_{n80c}\bar{H}^2_{n90c}} \\ &= \frac{1}{C}\sum \frac{\hat{r}^2_{i90c}}{N_{i90c}} + \frac{\hat{r}^2_{n90c}}{N_{n90c}} + \frac{\hat{r}^2_{i80c}}{N_{i80c}} + \frac{\hat{r}^2_{n80c}}{N_{n80c}}\end{aligned}, \tag{A.17}$$

where r^2_{jtc} is squared coefficient of variation in hours for nativity group j in period t in metropolitan area c. Lastly, the covariances between ν_{jtc} and ω_{jtc} are likely to be quite small. I will carry $\bar{\sigma}_{\nu\omega_c}$ through the analysis, but will assume it is equal to zero empirically.

The variance in the observed values h_c can be partitioned such that

$$\sigma^2_h = \sigma^2_{h^*} + \bar{\sigma}^2_{\eta_c} + \bar{\sigma}^2_{\nu_c} + \bar{\sigma}_{\nu\omega_c}, \tag{A.18}$$

implying an approximately consistent estimator of $\sigma^2_{h^*}$:

$$\hat{\sigma}^2_{h^*} = \hat{\sigma}^2_h - \widehat{\bar{\sigma}^2_{\eta_c}} - \widehat{\bar{\sigma}^2_{\nu_c}} \tag{A.19}$$

where $\hat{\sigma}_h^2 = (1/C)\sum \dot{h}^2$. Therefore, a consistent estimator of β is

$$\tilde{\beta} = \frac{\hat{\beta} - \hat{\pi}}{1 - \hat{\pi}} \tag{A.20}$$

where

$$\hat{\pi} = \frac{\widehat{\bar{\sigma}_{\eta_c}^2} + \widehat{\bar{\sigma}_{\nu_c}^2}}{\hat{\sigma}_{h^*}^2 + \widehat{\bar{\sigma}_{\eta_c}^2} + \widehat{\bar{\sigma}_{\nu_c}^2}}. \tag{A.21}$$

If $\bar{\sigma}_{\nu\omega_c} \neq 0$, this estimator will not fully correct $\hat{\beta}$. Given plausible values for $\bar{\sigma}_{\nu\omega_c}$, however, the degree of inconsistency induced by assuming that $\bar{\sigma}_{\nu\omega_c} = 0$ is quite small.

Table A.1 presents a comparison of the estimates of c and σ using actual labor supply estimated by OLS (columns 1 and 2, taken from columns 1 and 2 of Table 5), OLS adjusted for errors-in-variables bias (columns 3 and 4), and IV (columns 5 and 6, taken from columns 3 and 4 of Table 5). In addition, the last four columns of Table A.1 present estimates of $\sigma_{h^*}^2$, $\bar{\sigma}_{\eta_c}^2$, $\bar{\sigma}_{\nu_c}^2$, and π. The estimates of π indicate that $\bar{\sigma}_{\nu_c}^2$ and $\bar{\sigma}_{\eta_c}^2$ account for a significant portion of σ_h^2, with the majority coming from $\bar{\sigma}_{\eta_c}^2$. The estimates of σ adjusted for errors-in-variables bias in column 4 usually fall between those of OLS and IV and are all large enough (or, if negative, have a corresponding $\hat{c}$ that is not statistically significantly different from zero) to indicate, once again, that immigrants and natives are roughly perfect substitutes in production.

Appendix B: Description of the Data

The data used in this paper are the five percent PUMS of the 1980 and 1990 Censuses. Immigrants are defined as individuals born outside of the U.S. and its territories. Children born abroad of American parents are considered natives. In both Censuses, I used the citizenship variable (CITIZEN) to define immigration status.

Sample Definition

The samples used represent the employed civilian, non-institutionalized population living in the 48 contiguous states and the District of Columbia. I exclude all individuals who lived in group quarters or who were "augmented" (AAUGMENT $\neq 0$ in the 1990 file) individuals; who were in the military or enrolled in school; who had no weeks worked, zero usual hours worked, nonzero farm self-employment income, or negative non-farm self-employment income in the year prior to the Census; whose sum of wage and salary income and non-farm self-employment income was not greater than zero; whose industry code indicates that they were unemployed or in the military (i.e. a code greater than 932 in either Census); or who lived in Alaska or Hawaii.

Weighting

The 1980 data are a 5/100 random sample of the population. When population estimates are presented each observation from these data is given a weight of 20. The 1990 data are a nonrandom sample and weights are used for all analysis with them. The weighting variable used is PWGT1.

Definition of Income and Topcoding

Throughout the paper, nominal amounts for 1990 are deflated using the change in the personal consumption expenditures index (*Economic Report of the President* 1995, Table B-3) between 1979 and 1989, that is they are multiplied by a factor of .59195. The income measure I use is the sum of wage and salary income (INCOME1) and non-farm self-employment income (INCOME2). I include self-employment income both because the number of individuals reporting self-employment income has increased over time (from approximately 10 percent of males in 1980 to 11 percent of males 1990 and from approximately 4 percent of females in 1980 to 8 percent in 1990) and because there are differences between immigrants and natives within sex $\times$ skill groups in the incidence of self-employment income. The topcode values for wage and salary income and self-employment income changed between 1980 and 1990. The topcode value for wage and salary income was \$75,000 in 1980 and \$140,000 (\$82,873 in 1979\$) in 1990 while the topcode value for non-farm self-employment income was \$75,000 in 1980 and \$90,000 (\$53,275 in 1979\$) in 1990. To avoid any biases in changes in real income due to inflation or differences in the topcoding procedure, I recode all wage and salary incomes greater than or equal to \$75,000 in 1979\$ (in either Census) to $1.6 \times \$75,000 = \$120,000$. Thus, some individuals whose wage and salary income was not topcoded in 1990 have an imputed value instead. For self-employment income, I recode all values greater than or equal to \$53,275 in 1979\$ to $1.6 \times \$53,275 = \$85,240$, implying that some individuals for whom self-employment income was not topcoded in 1980 have an imputed value. I use a factor of 1.6 to approximate the conditional mean of income for those whose income was censored. Topcoding of either or both income variables affected less than 1 percent of all female nativity × skill subsamples in both Censuses. This was also true of male dropout and high school graduate immigrants and natives in both Censuses. For males with some college, the topcoding rate was less than 2 percent in both Censuses while for those with a college diploma or more the topcoding rate was 6.4 percent or less in both Censuses.

Coding of Geography

The samples I use for the regression analysis are taken from the 50 largest metropolitan areas or consolidated metropolitan areas in the U.S., based on 1990 population. These 50 areas are Albany, Atlanta, Baltimore, Birmingham, Boston, Buffalo, Charlotte, Chicago, Cincinnati, Cleveland, Columbus, Dallas, Dayton, Denver, Detroit, Greensboro, Hartford, Houston, Indianapolis, Jacksonville, Kansas City, Los Angeles, Louisville, Memphis, Miami, Milwauke, Minneapolis, Nashville, New Orleans, New York, Norfolk, Oklahoma City, Orlando, Philadelphia, Phoenix, Pittsburgh, Portland, Providence, Richmond, Rochester, Saramento, St. Louis, Salt Lake City, San Antonio, San Diego, San Francisco, Seattle, Tampa, Washington, and West Palm Beach. The 50 areas are defined to be geographically consistent betwen 1980 and 1990. Jaeger (1995a) discusses in detail the methodology used to match geographies between the 1980 and 1990 codings used by the Bureau of the Census.

REFERENCES

Altonji, Joseph G., and David Card. "The Effects of Immigration on the Labor Market Outcomes of Less-skilled Natives." In *Immigration, Trade, and the Labor Market*, edited by John M. Abowd and Richard B. Freeman. Chicago: University of Chicago Press and National Bureau of Economic Research, 1991.

Bartel, A. "Where Do the New U.S. Immigrants Live?." *Journal of Labor Economics* 7 (1989): 371–391.

Borjas, George. "Assimilation, Changes in Cohort Quality, and the Earnings of Immigrants." *Journal of Labor Economics* 3 (1985): 463–489.

Borjas, George. "Immigrants, Minorities, and Labor Market Competition." *Industrial and Labor Relations Review* 40 (1987): 382–392.

Borjas, George. "National Origin and the Skills of Immigrants." In *Immigration and the Work Force: Economic Consequences for the United States and Source Areas*, edited by Borjas, G. and R. B. Freeman. Chicago: University of Chicago Press and National Bureau of Economic Research, 1992.

Borjas, George. "The Economics of Immigration." *Journal of Economic Literature* 32 (1994): 1667–1717.

Borjas, George. "Assimilation and Changes in Cohort Quality Revisited: What Happened to Immigrant Earnings in the 1980s?." *Journal of Labor Economics* 13 (1995a): 201–245.

Borjas, G. "The Economic Benefits from Immigration." *Journal of Economic Perspectives* 92 (1995b): 3–22.

Borjas, George, Richard B. Freeman, and Larry F. Katz. "On the Labor Market Effects of Immigration and Trade." In *Immigration and the Work Force: Economic Consequences for the United States and Source Areas*, edited by Borjas, George and Richard B. Freeman. Chicago: University of Chicago Press and National Bureau of Economic Research, 1992.

Borjas, George, Richard B. Freeman, and Larry F. Katz. "Searching for the Effect of Immigration on the Labor Market." Cambridge, MA: National Bureau of Economic Research Working Paper Number 5454, 1996.

Bound, John, David A. Jaeger, and Regina M. Baker. "Problems with Instrumental Variables Estimation When the Correlation between the Instruments and the Endogenous Explanatory Variable Is Weak." *Journal of the American Statistical Association* 90 (1995): 400–407.

Bound, John and George Johnson. "Changes in the Structure of Wages in the 1980's: An Evaluation of Alternative Explanations." *American Economic Review* 82 (1992): 371–392.

Chiswick, Barry. "The Effect of Americanization on the Earnings of Foreign Born Men." *Journal of Political Economy* 36 (1978): 897-922.

Economic Report of the President. Washington, DC: U.S. Government Printing Office., 1995.

Efron, Bradley. *The Jackknife, the Bootstrap, and Other Resampling Plans*. Philadelphia: Society for Industrial and Applied Mathematics, 1982.

Fienberg, Sidney E. *The Analysis of Cross-Classified Categorical Data*. Cambridge, Massachusetts: MIT Press, 1977.

Filer, Randall K. "The Effect of Immigrant Arrivals on Migratory Patterns of Native Workers." In *Immigration and the Work Force: Economic Consequences for the United States and Source Areas*, edited by G. J. Borjas and R. B. Freeman. Chicago: University of Chicago Press and National Bureau of Economic Research, 1992.

Freeman, Richard B. "Overinvestment in College Training." *Journal of Human Resources* 10 (1975): 287–311.

Friedberg, Rachel M., and Jennifer Hunt. "The Impact of Immigrants on Host Country Wages, Employment, and Growth." *Journal of Economic Perspectives* 92 (1995): 23–44.

Frey, William H. "Immigration and Internal Migration 'Flight' from US Metropolitan Areas: Toward a New Demographic Balkanisation." *Urban Studies* 324–5 (1995): 733–757.

Grossman, Jean. "The Substitutability of Natives and Immigrants in Production." *Review of Economics and Statistics* 64 (1982): 596–603.

Hamermesh, Daniel. *Labor Demand*. Princeton, New Jersey: Princeton University Press, 1993.

Hinkley, D. V. "Jackknifing in Unbalanced Situations." *Technometrics* 19 (1977): 285–292.

Jaeger, David A. "Essays in Empirical Labor Economics." Ph.D. Dissertation, University of Michigan, 1995.

Jaeger, David A. "Reconciling the Old and New Census Bureau Education Questions: Recommendations for Researchers." *Journal of Business and Economic Statistics* 154 (1997): 300-309.

Jaeger, David A., Susanna Loeb, Sarah E. Turner, and John Bound "Census Geocodings." NBER Working Paper 6772, October, 1998.

Juhn, Chinhui, Kevin M. Murphy, and Brooks Pierce. "Wage Inequality and the Rise in Returns to Skills." *Journal of Political Economy* 101 (1993): 410–442.

Katz, Larry F., and Kevin M. Murphy. "Changes in Relatives Wages, 1963–1987: Supply and Demand Factors." *Quarterly Journal of Economics* 107 (1992): 35–78.

LaLonde, Robert J., and Robert H. Topel. "Labor Market Adjustments to Increased Immigration." In *Immigration, Trade, and the Labor Market*, edited by J. M. Abowd and R. B. Freeman. Chicago: University of Chicago Press and National Bureau of Economic Research, 1991.

MacKinnon, J. G., and Halbert White. "Some Heteroskedasticity-Consistent Covariance Matrix Estimators with Improved Finite Sample Properties." *Journal of Econometrics* 29 (1985): 305–325.

Maddala, G. S. *Econometrics*. New York: McGraw–Hill, 1977.

Murphy, Kevin M., and Finis Welch. “The Role of International Trade in Wage Differentials.” In *Workers and Their Wages*, edited by M. Kosters. Washington, D. C.: American Enterprise Institute, 1991.

Murphy, Kevin M., and Finis Welch. “The Structure of Wages.” *Quarterly Journal of Economics* 107 (1992): 285–326.

Schoeni, Robert F. “The Effect of Immigrants on the Employment and Wages of Native Workers: Evidence from the 1970s and 1980s.” Santa Monica, CA: RAND, unpublished draft, 1996.

Statistical Abstract of the United States: 1993 (113th edition). Washington, D. C.: U.S. Government Printing Office, 1993.

Warren, Robert. “Estimates of the Undocumented Immigrant Population Residing in the United States, by Country of Origin and State of Residence: October 1992.” U.S. Immigration and Naturalization Service, unpublished draft, April, 1995.

Warren, Robert and Jeffery S. Passel. “A Count of the Uncountable: Estimates of Undocumented Aliens Counted in the 1980 United States Census.” *Demography* 24 (1987): 375–393.

White, Halbert. “Heteroskedasticity-Consistent Covariance Matrix Estimation and a Direct Test for Heteroskedasticity.” *Econometrica* 48 (1980): 817–838.

Woodrow, K. A., and Jeffery S. Passel. “Post-IRCA Undocumented Immigration to the United States: An Assessment Based on the June 1988 CPS.” In *Undocumented Migration to the United States*, edited by Frank D. Bean, Barry Edmonston, and Jeffery S. Passel. Washington, D. C.: Urban Institute Press, 1990.

Table 1

Immigrant and Native Employment in the U.S. : 1980 and 1990

	Employment (thousands)				Change 1980 to 1990 (%)			Immigrant Share of Skill Level (%)	
	1980		1990						
Skill Level	Immigrant	Native	Immigrant	Native	Immigrant	Native	Total	1980	1990
Men									
Dropouts	1,451	12,214	1,911	7,298	31.7	–40.3	–32.6	10.6	20.7
High School	801	17,411	1,328	18,596	65.9	6.8	9.4	4.4	6.7
Some College	487	8,905	966	13,700	98.2	53.8	56.1	5.2	6.6
College	771	9,570	1,312	12,687	70.1	32.6	35.4	7.5	9.4
Total	3,510	48,100	5,517	52,281	57.2	8.7	12.0	6.8	9.5
Women									
Dropouts	992	7,953	1,153	4,985	16.3	–37.3	–31.4	11.1	18.8
High School	838	16,924	1,210	17,852	44.4	5.5	7.3	4.7	6.3
Some College	435	7,607	872	13,545	100.4	78.0	79.3	5.4	6.1
College	411	5,779	853	9,683	107.6	67.5	70.2	6.6	8.1
Total	2,676	38,263	4,089	46,064	52.8	20.4	22.5	6.5	8.2
Total									
Dropouts	2,443	20,167	3,064	12,282	25.4	–39.1	–32.1	10.8	20.0
High School	1,638	34,335	2,538	36,448	54.9	6.2	8.4	4.6	6.5
Some College	923	16,513	1,839	27,245	99.3	65.0	68.8	5.3	6.3
College	1,182	15,349	2,165	22,370	83.1	45.7	48.4	7.2	8.8
Total	6,186	86,363	9,605	98,345	55.3	13.9	16.6	6.7	8.9

SOURCE: Five percent public-use microsamples of the 1980 and 1990 U.S. Census.

Table 2

Wages of Immigrants and Natives: 1980 and 1990

	Real Hourly Wage (1979$)				
	Immigrants		Natives		
Skill Level	1980	1990	1980	1990	$d\log(W_i/W_n)$
Men					
Dropouts	6.38	5.33	7.34	6.62	–.075
High School	7.79	6.81	8.23	7.56	–.048
Some College	8.84	8.42	9.25	9.17	–.041
College	12.83	13.47	13.06	14.30	–.041
Women					
Dropouts	4.23	3.93	4.50	4.39	–.049
High School	5.07	5.07	5.14	5.31	–.034
Some College	5.89	6.41	5.76	6.38	–.017
College	7.84	9.04	7.75	9.12	–.021

SOURCE: Five percent public-use microsamples of the 1980 and 1990 U.S. Census.
NOTE: 50 Largest Metropolitan Ares of the U.S.

Table 3
Contributing Factors to Changes in Labor Productivity: 1980 – 1990

	Change in Fixed-Price log(Predicted Wage) Due to													
	Immigrants								Natives					
	Country of Origin	Ability to Speak English	Time in United States	Age	Race/ Ethnicity	Educ.	Num.of Children	Total ($d\ell_i$)	Age	Race/ Ethnicity	Educ,	Num. of Children	Total ($d\ell_n$)	F
							Men							
	−.024	−.013	−.007	−.002	−.016	−.001		−.061	−.013	−.004	.010		−.008	
l	−.012	−.021	−.013	−.001	−.024			−.072	.025	−.006			.019	
ge	−.018	−.013	−.001	.020	−.015			−.027	.049	−.002			.047	
	−.005	−.004	.004	.016	−.009	.003		.006	.040	−.002	.002		.040	
							Women							
	−.005	−.007	.003	.003	−.010	−.001	−.001	−.018	−.001	.001	.004	−.002	.003	
l	.010	−.010	−.003	.010	−.012		−.005	−.011	.036	.002		−.011	.027	
ge	.001	−.005	.004	.022	−.000		−.005	.016	.049	.001		−.009	.041	
	.004	.005	.015	.007	.004	.005	−.000	.039	.044	.000	.013	−.004	.053	

: Five percent public-use microsamples of the 1980 and 1990 U.S. Census.
) Largest Metropolitan Areas of the U.S.

Table 4

Naive Estimates of the Elasticity of Substitution between Immigrants and Natives within Sex and Skill Groups

	Relative Hours $d\log(H_i/H_n)$		Relative Wages $d\log(W_i/W_n)$		Implied $\hat{\sigma}$	
Skill Level	Actual	Adjusted	Actual	Adjusted	Actual	Adjusted
			Men			
Dropouts	.887	.833	–.075	–.022	12	38
High School	.516	.425	–.048	.043	11	–10
Some College	.292	.219	–.041	.032	7	–7
College	.244	.210	–.041	–.007	6	31
			Women			
Dropouts	.718	.697	–.049	–.028	15	25
High School	.427	.389	–.034	.003	13	–115
Some College	.158	.134	–.017	.007	9	–18
College	.161	.147	–.021	–.007	8	21

SOURCE: Five percent public-use microsamples of the 1980 and 1990 U.S. Census.
NOTE: 50 Largest Metropolitan Areas of the U.S. "Adjusted" hours and wages are adjusted using the change in mean log fixed-price wages between 1980 and 1990 from Table 3.

Table 5

Estimated Elasticity of Substitution between Immigrants and Natives within Sex and Skill Groups

	Actual Quantitites				Adjusted Quantities			
	OLS		IV		OLS		IV	
	$\hat{c}$	$\hat{\sigma}$	$\hat{c}$	$\hat{\sigma}$	$\hat{c}$	$\hat{\sigma}$	$\hat{c}$	$\hat{\sigma}$
Skill Level	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Men								
Dropouts	.106	9.4	.130	7.7	.076	13.1	.118	8.4
	(.053)		(.056)		(.037)		(.041)	
High School	.127	7.8	.163	6.1	.113	8.8	.138	7.3
	(.054)		(.046)		(.053)		(.046)	
Some College	.030	33.3	.098	10.2	–.003	–389.7	.050	20.2
	(.101)		(.093)		(.078)		(.093)	
College	.091	11.0	.266	3.8	.079	12.6	.246	4.1
	(.101)		(.123)		(.101)		(.115)	
Women								
Dropouts	.020	50.1	.038	26.6	.005	219.6	.008	120.8
	(.048)		(.057)		(.051)		(.062)	
High School	.168	5.9	.206	4.9	.157	6.4	.182	5.5
	(.063)		(.074)		(.069)		(.074)	
Some College	–.008	–123.9	.066	15.1	–.009	–112.3	.071	14.1
	(.066)		(.101)		(.067)		(.099)	
College	.175	5.7	–.010	–103.3	.186	5.4	.028	35.2
	(.139)		(.173)		(.127)		(.156)	

SOURCE: Five percent public-use microsamples of the 1980 and 1990 U.S. Census.
NOTE: Sample is 50 Largest Metropolitan Areas of the U.S. Jackknife standard errors in parentheses. Instrument for all IV models is the population aggregate $d\log(N_i/N_n)$. F statistic on instrument in first stage of IV models ranges from 21.8 to 109.6.

Table 6

Immigrant and Native Supplies of Three Skill Groups: 1980 and 1990

		Skill Group		
Nativity Group		Dropouts	High School	College
			1980	
Immigrants	(a)	.015	.011	.024
Natives	(b)	.158	.286	.506
Effects of Immigration	(c)=log[(a+b)/b]	.092	.037	.046
			1990	
Immigrants	(d)	.014	.013	.035
Natives	(e)	.071	.243	.623
Effects of Immigration	(f)=log[(d+e)/e]	.184	.052	.055
			Changes	
Effects of Immigration	(f)-(c)	.092	.015	.009

SOURCE: Five percent public-use microsamples of the 1980 and 1990 U.S. Census.
NOTE: Supplies have been normalized to sum of total of dropouts, high school, and college within year. Dropouts are weighted by .9 and some college weighted by .3 in creating shares.

Table 7

Estimated Effect of Immigrants on Native Log Real Wage Levels: 1980 – 1990

Elasticities of Substitution		Change in Log Wage Levels Due to Immigrants' Effect on Changes in Supplies of			
r	s	Dropouts	High School	College	Total
		Change in the Wages of Dropouts			
		(Share in $L = .310$, $d\log(D/D_n) = .092$, $d\log(W_{dn}) = -.102$)			
1	3	−.038	−.003	.005	−.036
1	6	−.028	−.004	.005	−.027
1	9	−.024	−.005	.005	−.024
1	∞	−.017	−.006	.005	−.018
1.5	3	−.033	−.001	.004	−.030
1.5	6	−.022	−.002	.004	−.021
1.5	9	−.018	−.003	.004	−.018
1.5	∞	−.011	−.004	.004	−.012
2	3	−.030	.000	.003	−.027
2	6	−.019	−.001	.003	−.018
2	9	−.016	−.002	.003	−.015
2	∞	−.008	−.003	.005	−.009
		Change in the Wages of High School Graduates			
		(Share in $L = .690$, $d\log(S/S_n) = .015$, $d\log(W_{hn}) = -.088$)			
1	3	−.007	−.007	.005	−.010
1	6	−.012	−.007	.005	−.014
1	9	−.014	−.006	.005	−.015
1.5	3	−.002	−.005	.004	−.004
1.5	6	−.007	−.005	.004	−.008
1.5	9	−.008	−.004	.004	−.009
2	3	.001	−.004	.003	−.001
2	6	−.004	−.004	.003	−.005
2	9	−.005	−.003	.003	−.006
		Change in the Wages of College Equivalents			
		(Share in $G = .594$, $d\log(C/C_n) = .009$, $d\log(W_{cn}) = .032$)			
1		.012	.004	−.004	.012
1.5		.008	.003	−.002	.008
2		.006	.002	−.002	.006

SOURCE: Five percent public-use microsamples of the 1980 and 1990 U.S. Census.

Table A.1

Comparison of OLS, OLS Adjusted for Errors-in-Variables Bias, and IV Estimates of Elasticity of Substitution between Immigrants and Natives within Sex and Skill Groups

	OLS				IV		Variance Components for Errs.-in-Vars. Adjustment			
	Unadjusted		Adjusted for Errs.-in-Vars.							
Skill Level	$\hat{c}$	$\hat{\sigma}$	$\hat{c}$	$\hat{\sigma}$	$\hat{c}$	$\hat{\sigma}$	$\hat{\sigma}^2_{h^*}$	$\widehat{\bar{\sigma}^2_{\eta_c}}$	$\widehat{\bar{\sigma}^2_{\nu_c}}$	$\hat{\pi}$
Men										
Dropouts	.106	9.4	.126	7.9	.130	7.7	.159	.025	.005	.159
High School	.127	7.8	.156	6.4	.163	6.1	.096	.019	.003	.184
Some College	.030	33.3	.044	22.9	.098	10.2	.064	.026	.003	.312
College	.091	11.0	.160	6.3	.266	3.8	.018	.012	.001	.431
Women										
Dropouts	.020	50.1	.025	39.6	.038	26.6	.129	.026	.008	.209
High School	.168	5.9	.227	4.4	.206	4.9	.048	.014	.003	.258
Some College	–.008	–123.9	–.013	–77.5	.066	15.1	.046	.023	.005	.374
College	.175	5.7	.352	2.8	–.010	–103.3	.030	.024	.006	.502

SOURCE: Five percent public-use microsamples of the 1980 and 1990 U.S. Census.

NOTE: Sample is 50 largest metropolitan areas of the U.S. OLS estimates of c and σ are from columns 1 and 2 of Table 5. IV estimates are from columns 3 and 4 of Table 5.

Figure 1
Immigrant Share of Education Category and
Distribution of Education Within Nativity Group: 1980 and 1990