The Impact of Immigration on the Educational Attainment of Natives *

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Abstract

Using a state panel based on census data from 1940–2008, I examine the impact of immigration on the high school completion of natives in the United States. Immigrant children could compete for schooling resources with native children, lowering the return to native education and discouraging native high school completion. Conversely, native children might be encouraged to complete high school in order to avoid competing with immigrant high–school dropouts in the labor market. I find evidence that both channels are operative and that the net effect is positive, particularly for blacks, though not for Hispanics. An increase of one percentage point in the share of immigrants in the population aged 11–64 increases the probability that natives aged 11–17 eventually complete 12 years of schooling by 0.1–0.2 percentage points, and increases the probability for native–born blacks by 0.4–0.5 percentage points. I account for the endogeneity of immigrant flows by using instruments based on 1940 settlement patterns.

The extent to which the children of low–education or low–income parents are able to achieve their full potential in the United States is a cause for concern when viewed from an international perspective. Contrary to popular mythology, there is less intergenerational mobility in earnings and education in the United States than in continental Europe and Canada, and no more than in the United Kingdom.[1] An important step upward for many children from low socio–economic status families is graduation from high school. Yet U.S. high school graduation rates are no longer increasing, and indeed are falling, even for the native–born.[2] In this paper, I contribute to our understanding of the determinants of high school educational attainment by investigating the role of immigration. Increasing immigration in recent decades has led to popular concern that immigration is reducing the quality of K–12 education. If this concern is well founded, rising immigration could reduce native high school graduation rates. Conversely, immigration–induced changes in labor market incentives for educational attainment could have the opposite effect. I seek evidence for these two channels and assess their net effect.

Immigrants and the young children of immigrants generally have a more limited command of English than natives. If immigrants and natives are taught in the same classes, teachers of some subjects may slow the pace of instruction to accommodate non–native speakers. If immigrant students have had worse quality prior education, or have less education than their native classmates, teachers may lower expectations for all students. Immigrant students could also divert financial resources from native students, potentially lowering the quality of their education. For example, Fix and Zimmerman (1993) find that federal Chapter I spending per economically disadvantaged student fell due to the immigration–induced expansion in the number of eligible children. If immigrants are taught separately from natives, resources might also be diverted from natives. For example, federal Title III money for Limited English Proficient education may come at the expense of other federal funding. A lower educational quality for natives will reduce their earnings capacity at a given number of years of education, and this lower return to

[1] Checchi, Ichino and Rustichini (1999); Corak (2006).
[2] Heckman and LaFontaine (2010).

education in turn may induce natives to complete fewer years of high school.

There exists a second channel, however, through which immigration could increase natives' high school educational attainment (Betts 1998). Incentives to complete education are influenced by the wage structure, which is in turn affected by the entry of adult immigrant workers. Immigration will affect wage inequality among natives if the distribution of immigrant skill differs from that of natives. Compared to natives, immigrants to the United States are very disproportionately poorly educated and somewhat disproportionately highly educated. Immigrants are underrepresented among workers with an intermediate level of education, such as a high school diploma. The effect of immigrants entering the labor market should therefore be to increase wage inequality in the lower half of the distribution, particularly the wage gap between high school dropouts and high school graduates. Empirical studies confirm this.[3] The net effect of the changes in the wage structure is likely to be to increase the return to completing high school, and hence native completion rates.

Any negative effects on the schooling quality of natives will affect the children of low socio–economic status (SES) parents more than children of high SES parents. Richer parents may move their child to a learning environment with either fewer immigrants or immigrants with better language skills and educational background, by using private schools (Betts and Fairlie 2003) or by moving to a different school district (Cascio and Lewis 2010). In these schools, those immigrants present may provide positive peer effects. Furthermore, the educational quality of a child's school is likely to have a smaller impact on the children of high SES parents, as such parents can compensate in part for a school's deficiencies by providing the child with instruction at home. Any negative effects of immigrants on educational attainment are therefore more likely to be observable in high school graduation rates, a margin very relevant for children of low SES families, than in the attainment of higher levels of education.

I therefore focus on the impact of immigration on natives' completion of 12 years of schooling. I compare the results for blacks, Hispanics and non–Hispanic whites, with

[3] Borjas and Katz (2007), Ottaviano and Peri (2008).

the expectation that any impact of immigration will be larger for minorities, who attain on average lower education than non–Hispanic whites. I use the decennial censuses of 1940–2000 and the pooled 2006–2008 American Community Surveys (ACS) to construct a state panel. I extend two closely related papers, Betts (1998) and Betts and Lofstrom (2000), in several ways. The most important extensions in practice are the distinction between immigrants of different educational attainment and the closer matching of the immigrant inflows to natives' adolescent years. The extension to the use of instrumental variables based on historical immigrant settlement patterns is important in principle but less important in practice.

I measure the shares of immigrants in the population when natives are aged 11–17, and I measure native educational attainment at ages 21–27. I find that a one percentage point increase in the share of immigrants in the population aged 11–17 reduces the probability natives eventually complete 12 years of school by 0.3–0.5 percentage points for all races and ethnicities, and by 0.4–0.6 percentage points for blacks, though not at all for non–Hispanic whites. For native Hispanics, the child immigrants of more educated parents have beneficial effects, offsetting detrimental effects of children of unskilled immigrants. Immigration of unskilled adults has a beneficial effect for all natives groups: a one percentage point increase in the share of immigrants with less than 12 years school in the population aged 18–64 increases the eventual native completion rate by 0.7–0.9 percentage points for all races and ethnicities, and by 0.8–1.4 percentage points for blacks and Hispanics. Effects of more educated adult immigrants are generally insignificant.

Unlike Betts (1998) and Betts and Lofstrom (2000), who found a detrimental net effect of immigration on native educational attainment for each native racial and ethnic group, I find the net effect of immigration to be positive for natives generally, and especially for blacks: an increase of one percentage point in the share of immigrants in the population aged 11–64 increases the probability natives complete 12 years of schooling by 0.1–0.2 percentage points, and increases the probability for blacks by 0.4–0.5 percentage points. All effects are rather small compared to the average native completion rate of 86.0% (80.5% for blacks) and given the average immigrant share of 8.9% (8.1% for blacks).

I estimate a detrimental net effect for native–born Hispanics of -0.2 percentage points, although this is imprecisely estimated.

These results are consistent with the theoretical prediction that child immigrants would reduce native educational attainment through the school quality channel, while adult unskilled immigrants would increase it through the labor market channel (as would their children, if native youth are sufficiently forward looking). This suggests the need for reform in immigrant education. Reform could include both increased resources for schools in areas with high immigration and the implementation of best practices regarding improving language skills of non–native speakers, remedying educational deficiencies of immigrants, and integrating immigrants with native students (García, Kleifgen and Falchi 2008). Nevertheless, some caution should be exercised in attributing the negative effects of immigration to the school quality channel, as no school–level data are used. The finding that on net, immigration causes natives to upgrade their education sheds light on why immigration appears to have little or no negative effect on the overall wage levels of the native–born.[4]

My results complement some other findings in the literature. Gould et al. (2009) find that increased numbers of immigrants in Israeli schools reduce the native pass rate for the high school leaving exams. Neymotin (2009)'s school–level analysis finds native SAT scores and probability of applying to top colleges are not negatively affected by the share of immigrant test–takers, but she does not instrument to account for the endogeneity of the share of immigrants at the school. Jackson (2009) finds that immigration of skilled workers increases native college enrollment while inflows of immigrants judged to be college students decrease it, and that the ratio of unskilled to skilled immigrant workers has a positive effect (he does not control for the level of unskilled immigration). Llull (2010) estimates a structural model of the U.S. labor market in which he finds evidence that natives increase their years of education in response to immigration.[5]

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[4] See also Peri and Sparber (2008, 2009).

[5] Borjas (2006) finds that immigration does not reduce native enrollment in graduate school, while Hoxby (1998) shows that native students shifted out of the California State system in favor of the University of California system and private colleges when the former became relatively cheaper for undocumented

1 Data and descriptive statistics

The data come principally from the IPUMS micro–data samples for the 1940–2000 decennial censuses and the combined 2006–2008 American Community Surveys.[6] I choose these data for the large sample sizes they afford for the measurement of both state immigrant shares and the shares of native cohorts, particularly minority cohorts, attaining a given educational level. Even with the large census samples, I am forced to consider a native cohort as one spanning several birth years. A disadvantage of the data is that they do not contain information about parental education or income, except for children living with their parents. I supplement them with data from the Bureau of Economic Analysis on state personal income per capita.

In order to have a consistent outcome variable over all years, I define the outcome of interest as the completion of 12 years of schooling, with or without the obtention of a high school diploma, a distinction only possible from 1990 onwards.[7] I focus on the native–born who were aged 11–17 in the previous census: this implies current ages 21–27 except for respondents to the 2006–2008 ACS, who are aged 19–25.[8] Most covariates are lagged one census, and refer to the time when natives were aged 11–17. I construct samples of all races and ethnicities pooled, blacks, Hispanics and non–Hispanic whites. Being black and Hispanic are not mutually exclusive, so there is some overlap in the two minority samples. Immigrants are defined as those born abroad, including those born in U.S. territories. I drop the states of Alaska and Hawaii, as their absence from the 1940 and 1950 censuses complicates the use of the instruments and covariates measured in 1940.

Figure 1 depicts the shares of native–born 21–27 year olds who have completed at least 12 years of schooling, by race and ethnicity for 1940–2008.[9] The shares increase strongly over the early decades then level off around 1990. Minorities begin the period with much

immigrants.

[6] Ruggles et al. (2010).

[7] Betts (1998) uses 12 years of schooling for 1980, and obtention of a high school diploma in 1990. This discrepancy drives the negative effect he finds of immigration on Hispanics.

[8] I drop natives aged 17 and 18 in 2006–2007.

[9] For the purposes of the graph I use 21–27 for all years including 2006–2008.

lower education, and converge towards non–Hispanic white rates from 1960 until 1980 or 1990. At the start of the period, both blacks and Hispanics (concentrated in different regions) were educated in segregated, inferior schools. As a result of court decisions in the 1940s and 1950s, the Civil Rights Act of 1964 and the Coleman Report (Coleman et al. 1966), educational quality, integration and attainment increased for minorities.[10] Heckman and LaFontaine (2010) have cautioned that both the increase in high school education generally and the convergence between whites and minorities specifically mask an increasing share of individuals receiving a GED (and presumably reporting 12 years of education). However, a regular high school diploma and a GED cannot be distinguished in the decennial censuses.[11]

Figure 2 shows the evolution of the share of immigrants over the period, by age group. The share of immigrants in the population of working age, 18–64, traces out a U–shape, falling from 12.1% in 1940 to 6.1% in 1970, before rising to 17.8% in 2008 (top line). The share of immigrants in the school–age population, 11–17, traces a different path, rising almost monotonically from 1.6% in 1940 to a still modest 7.1% in 2008 (bottom line).

Figure 3 shows the time paths of three additional key covariates: the shares of the population aged 18–64 composed of immigrants with less than 12 years of education, exactly 12 years of education, and more than 12 years of education. The share of the lowest education immigrants falls from a high of 10.1% in 1940 to a low of 2.8% in 1980, before rising again to 4.7% in 2008. The shares of the immigrants from the two more educated groups rise monotonically from 0.7-1.1% in 1940 to 7.6% in 2008 for those with more than 12 years education, and 5.5% for those with exactly 12 years of education. Appendix Table 1 gives further means of variables measured at the individual level, while Appendix Table 2 give means of variables measured at the state level.

[10] MacDonald and Monkman (2005), Valencia et al. (2002).

[11] Heckman and LaFontaine (2010) combine census data and administrative data on GED awards to compute the regular high school graduation rate at the national level. Possibly something similar could be done at the state level.

2 Estimation

While the main analysis is conducted on a panel of states, I first adjust at the individual level for variation in age and gender structure across states and years. Specifically, for individuals aged 21–27 (19–25 in 2006–2008) at time t and born in state s, I run the following linear probability regression for each of the samples:

$$P(E_{ist} \geq 12) = \alpha_0 + \alpha_1 F_{ist} + \sum_{a=22}^{a=27} \gamma_a A^a_{ist} + \sum_s \sum_t \lambda_{st}(\delta_s \times \nu_t) + \eta_{ist}, \tag{1}$$

where i indexes individuals and s individuals' birth state, E represents years of education, F is a gender dummy, A^a are dummy variables for age, δ_s are state dummies and ν_t are year dummies. I weight this regression using weights based on the census weights. The census weights sum to the U.S. population of the census year, while I wish the standard errors to reflect the variation in sample sizes from year to year. I adjust the census weights so that the ratios of their sums for each year reflect the ratio of the census sample sizes, resulting in considerably more weight being put on recent years. The average year in the weighted data is 1989 for non–Hispanic whites, 1991 for blacks and 1995 for Hispanics.[12]

In a second step, I use the coefficients $\hat{\lambda}_{st}$ as the dependent variable in the subsequent state panel analysis:

$$\hat{\lambda}_{st} = \beta_0 + \beta_1 I^{11-17}_{s,t-10} + \beta_2 I^{E<12}_{s,t-10} + \beta_3 I^{E=12}_{s,t-10} + \beta_4 I^{E>12}_{s,t-10} + \beta_5 X_{s,t-10} + \gamma_s + \nu_t + \epsilon_{st}. \tag{2}$$

I weight the regressions with the inverse of the squared standard errors on the $\hat{\lambda}_{st}$ in the first step, and cluster the standard errors by state. $I^{11-17}_{s,t-10}$ represents the share of the population aged 11–17 that is foreign–born in the previous census, when the native–born cohort was itself aged 11–17, and is designed to capture natives' exposure to immigrant classmates. Ideally, an additional covariate would capture the presence of immigrants when natives were of elementary school age, but the ten–year spacing of the census precludes this. The immigrant covariates are affected by measurement error due to small

[12] The 1940, 1960 and 1970 censuses are 1% samples, the 1980–2000 censuses are 5% samples, the ACS has a more complicated sampling scheme which results in the pooled 2006–2008 sample being smaller than the 2000 sample, and the 1950 census only asked education questions of a subset of the main 1% sample.

samples, and due to being measured for the native's birth state: using birth state avoids endogenous moves of young adults that would plague the use of state of current residence, yet the native may have moved in childhood. The null hypothesis to be tested is that β_1 is negative because immigrant children reduce school quality. However, it is possible that immigrant children increase school quality, particularly for non–Hispanic whites who may share schools with immigrants whose English skills and prior education are better.

$I_{t-10}^{E<12}$ represents the share of the population aged 18–64 when natives were aged 11–17 that was immigrants with less than 12 years of schooling, and $I_{t-10}^{E=12}$ and $I_{t-10}^{E>12}$ are defined similarly. The null hypothesis to be tested is that β_2 is positive, because the presence of immigrants with less than 12 years education increases the return to completing 12 or more years of education. The necessity of using multi–year birth cohorts is likely to bias β_2 towards zero, since the younger members of the age range 11–17 are likely to base their years of schooling decision on the wage structure, and hence immigration rates, of later years. The signs of β_3 and β_4 are ambiguous, as the inflows of more educated immigrants have opposite effects on the return to exactly 12 years of education versus more than 12 years of education (relative to less than 12 years).

This regression suffers from endogeneity problems, however. Native high school educational attainment and high shares of low–education immigrants in a state may be spuriously negatively correlated. What makes the state economically attractive for immigrants, such as the availability of low–skill jobs, may by the same token mean that natives have a low incentive to complete 12 years of schooling. For example, a downturn in a state's low–skill industries could deter immigrants from moving to the state and encourage its natives to graduate from high school, leading β_2 to be biased down (the same direction as the measurement error bias). Similar reasoning suggests that β_3 and β_4 could be biased up by endogeneity. β_1 could be biased up if immigrants with children choose states with high educational attainment (the same direction as the measurement error bias), but there may be other biases due to endogeneity in their parents' choice of state if these have not been controlled for properly.

Due to these endogeneities, I implement an instrumental variables strategy using ten–

year differences of equation (2):

$$\Delta\hat{\lambda}_{st} = \tau_0 + \beta_1 \Delta I^{11-17}_{s,t-10} + \beta_2 \Delta I^{E<12}_{s,t-10} + \beta_3 \Delta I^{E=12}_{s,t-10} + \beta_4 \Delta I^{E>12}_{s,t-10} + \beta_5 \Delta X_{s,t-10} + \nu_t + \Delta\epsilon_{st}. \quad (3)$$

I estimate this using weights $1/(1/w_{s,t} + 1/w_{s,t-10})$, where w is the weight used in equation (2). I devise instruments for the differenced immigration covariates, based on the flows of immigrants to a state that would have been expected given the 1940 geographic distribution of immigrants from different regions and the subsequent national inflows from those regions.[13] To illustrate, if immigrants from Europe prefer the northeastern United States because it is closer to home and because other Europeans are already there because of geography, and Mexican immigrants prefer the southern border states for analogous reasons, the large national increase since 1940 in the share of immigrants that are Mexican will be associated with an increase in immigration to the southern border states relative to the northeast. The predicted flows captured in the instrumental variable will therefore be strongly, though not perfectly, correlated with actual immigrant flows to states. Furthermore, since the national increase in Mexican immigration appears to be the result of increasingly large birth cohorts entering the Mexican labor market[14], and the national decrease in European immigration is due to Europe's having become richer, the decrease in immigration to the Northeast relative to the border states is unrelated to non–immigration factors affecting native education choices.

I define an instrument for each of the education–specific immigration variables as follows. For a state s, the predicted change in the number of immigrants of education level E (aged 18–64), caused by changing origin regions k, can be written as

$$\Delta\hat{M}^E_s = \sum_k \frac{M_{sk}}{M_k} \Delta M^E_k = \sum_k \mu_{sk} \Delta M^E_k, \quad (4)$$

where μ_{sk} is state s's share in 1940 of the national total of immigrants who originate from region k, and ΔM^E_k is the national change in the number of immigrants with education E (aged 18–65) from that region. I use 18 source regions or countries, listed in Appendix

[13] These instruments are similar to the instrument developed by Card (2001), and also used by Jackson (2009) and Hunt and Gauthier–Loiselle (2010).

[14] Hanson and McIntosh (2007).

Table 3. Because the variables to be instrumented are percentage point changes, I convert $\Delta\hat{M}_s^E$ to percentage points by dividing by the population level (aged 18–65) at the start of the period to which Δ refers, to define the final instrument as:

$$Z_s^E = \frac{\Delta\hat{M}_s^E}{POP_s}. \tag{5}$$

I deliberately base the μ_{sk} on immigrants of all educations (and ages) to emphasize the role of geography and taste and minimize the role of economic factors that might disproportionately attract workers of a specific education level. The instrument will be invalid if non–immigration shocks to high school completion are correlated with 1940 immigrant densities; for example, if improvements to the California school system caused a national–level increase in Mexican immigration.[15] By defining an instrument for each education level, I assume that improvements to the California school system did not cause a national–level increase in Mexican immigration of any education group.

It is easy to construct instruments for different immigrant age groups, in particular for ΔI^{11-17}, by replacing the education–specific variables in equations (4) and (5) with age–specific variables, and I do so. However, the intuition of the instruments extends less easily to subdivisions by age group, as immigrant numbers in an age group are strongly influenced by aging as well as immigration, and furthermore, for the 11–17 age group in particular, it is obvious that changes in their inflows will be closely linked to those of adult immigrants.

I choose to use a common first stage for all four race/ethnicity samples, weighting each first stage with the denominator of its dependent variable (the population 11–17 or the population 18–64) in order to improve efficiency. The instruments have weak predictive power when the first stages are weighted by the second stage black or Hispanic weights (essentially the black or Hispanic population aged 21–27). This approach also means the first stage always includes all states and years: some early state–year cells have no native–born blacks or Hispanics aged 21–27.

Immigrant students could have either positive or negative spillovers on their native

[15] See Beaudry et al. 2010 for a formal treatment.

classmates, depending upon the quantity and quality of their prior education, their English skills, their industriousness, and the extent to which their parents contribute to their education. Parental education is likely to be a proxy for some of these characteristics, and because most children aged 11–17 live with their parents, we can observe their parents' education in the census data. It is therefore possible to split the share of the population 11–17 that is immigrant into immigrants whose parent or parents in the household have less than 12 years schooling, immigrants with at least one parent with 12 years or more, and immigrants with neither parent living in the household. The expectation is that children of more educated parents will make better peers than children of less educated parents. I discuss possible biases on the coefficients in the results section below. The difficulty with regressions distinguishing child immigrants according to parental characteristics is that the number of endogenous variables becomes too large for the use of 2SLS.

Child immigrants can also be distinguished according to the region of birth of their parents, but this distinction does not matter once child immigrants are distinguished according to parental education. English language ability is measured only in recent censuses.[16]

3 Results

I examine the impact of immigration on the probability of natives' completing 12 years of schooling, for all native–born, non–Hispanic whites, blacks and Hispanics, first assessing the net impact of immigration, then decomposing the impact into school quality versus labor market channels.

[16] I have experimented with using the (OLS) state–year return to completing 12 years schooling as the independent variable of interest, instrumenting it with actual or predicted immigrant flows. The coefficient on the return is always very imprecisely estimated, and the first–stage immigrant coefficients are often wrongly signed.

3.1 The net effect of immigrants ages 11–64

In Table 1, the first six columns show the impact of the immigrant share of the population aged 11–64 using the fixed effects specification of equation (2) and increasing numbers of covariates, column 7 shows the impact using the ten–year differenced specification of equation (3), and column 8 shows the results of applying two stage least squares (2SLS) to the column 7 specification. The first row shows that the principal coefficient of interest is positive, and becomes larger and statistically significant as more covariates are added, with a magnitude of about 0.2. The 2SLS coefficient is 0.23, but significant only at the 10% level, as the standard errors increase despite a strong first stage (the F–statistic associated with the instrument in the first stage is 25). A coefficient of 0.2 implies that a one percentage point increase in the share of immigrants in the population 11–64 increases the native probability of eventually completing 12 years of education by 0.2 percentage points. This is not a large effect considering that the (weighted) mean completion rate is 86.0%, and the share of immigrants in the population 8.9%.

Following existing literature on the determinants of high school completion, the first controls I add in Table 1 column 2 are for the unemployment rate and the cohort size, measured in the previous census when native respondents were aged 11–17. Neither the youth unemployment rate nor the prime–age unemployment rate in the previous census is statistically significant, possibly because many of the respondents were some years from graduation when the unemployment rate was measured.[17] Consistent with Card and Lemieux (2001), I find that members of larger cohorts have statistically significantly lower educational attainment – I define the cohort size based on natives only, to avoid endogeneity, as the share of the native population which is aged 11–17. However, when I control for region–specific trends in column 3 (dummies for eight BEA regions interacted with a trend), the coefficient on cohort size falls to zero.

In column 4, I control for the share of workers 18–64 who were in agriculture in 1940, interacted with a trend. This covariate is included for its effect on minorities, rather than

[17] State–level unemployment rates are not available from other sources for earlier decades, so the unemployment rate cannot be matched to the year the respondent was aged 17, for example.

non–Hispanic whites, and its coefficient is significant only at the 10% level for all natives. In column 5, I control for the 1940 share of non–Hispanic whites aged 21–27 who had less than 12 years education, interacted with a trend, which captures convergence among states and is statistically significantly positive. One would expect richer states to be able to afford better educational systems. However, states with better educational systems should become richer, so the coefficient on a control for state income would be biased up. In column 6, I show that the correlation between log state personal income per capita and completion of 12 years of schooling is indeed positive and statistically significant, and that the coefficient on the immigrant variable, now a lower bound on the true coefficient, is reduced by one third to 0.14. In the differenced specifications of columns 7 and 8, I control for all covariates except state personal income.

I check the sensitivity of the results to the choice of sample in Table 2, reporting results from differenced specifications, both least squares and 2SLS. In column 1, I reproduce the 2SLS result from the last column of Table 1. In columns 2 and 3, I restrict the sample to differences from 1970–80 or later for the outcome of interest, which implies differences from 1960–70 or later for covariates other than those for 1940, avoiding use of the small 1950 sample. The coefficients on the share of immigrants 11–64 are somewhat smaller, at 0.15 and 0.17, than their counterparts of 0.18 (Table 1 column 7) and 0.23 for all years. In column 4 and 5, I drop the 2000–2008 differences, as the spacing is not ten years like for the other differences. This has little effect on the least squares estimate (0.19 in column 4), but reduces the 2SLS to 0.11, rendering it statistically insignificant (column 5). In the remaining columns, I split the sample into men and women, and find the results to be similar: differences between men and women over ten–year periods are small compared to the common changes in the probability of completing 12 years of schooling. Overall, the finding that the coefficient is in the range 0.1–0.2 is robust to changes in the sample.

In Table 3, I analyze natives by race and ethnicity, using the same specifications as Table 1, reporting only the coefficient on the immigrant covariate. For reference, I reproduce the coefficients from the first row of Table 1, for all natives, in the first row of Table 3. The coefficients for non–Hispanic whites, in the second row, are always slightly smaller, and hence slightly less significant, than those for all natives, and the coefficient from 2SLS, in column 8, is statistically insignificant. The point estimate from specifications once convergence is controlled for in column 5 is robust at 0.12–0.18, however. For blacks in the third row, once state–specific trends and convergence are controlled for in columns 4 and 5, there is a robust positive coefficient of 0.43–0.51, compared to a mean completion rate of 80.5% and a black–white completion gap of 8.7 percentage points. For Hispanics, in the fourth row, an effect of 0.32–0.52 is robust in the least squares columns, but it disappears (coefficient of -0.19) with 2SLS in column 8. I conclude that the net effect of immigration on native completion of 12 years of schooling is positive for natives generally, blacks and probably non–Hispanic whites, and is imprecisely estimated, but possibly negative, for Hispanics.[18]

3.2 Decomposing the impact of immigration into school quality and labor market channels

I now turn to decomposing the impact of immigration into school quality and labor market channels. I return in Table 4 to the sample of all natives, presenting the same specifications as in Tables 1 and 3, except with immigrants split into four categories. The first row shows that the effect on natives of immigrants aged 11–17, likely to have been natives' classmates, is negative: once trends are controlled for, the coefficient is in the range -0.32– -0.45, statistically significant except for the 2SLS coefficient of -0.32 in column 8. The first stage information for the instrumenting of this variable is presented in Table 5, column 2 (in column 1, I present the first stage used for immigrants 11–64 in the tables above). Although the predicted share of the population 11–17 has a statistically significant coefficient, it is not more significant than those of the other excluded instruments, suggesting that I have not managed successfully to instrument the share of immigrants in

[18] For blacks and Hispanics, there is a large positive coefficient on the share of agricultural workers in 1940, which captures convergence among states: agricultural states in 1940 had large shares of either blacks or Hispanics in the population, who were poorly educated. I do not control for the educational attainment of blacks and Hispanics in 1940, as they are based on very small samples for many states. White educational attainment in 1940 is statistically insignificant in the regressions for minorities.

the population 11–17. The Angrist and Pischke (2008) F–statistics (which are adapted for multiple endogenous variables) for the excluded instruments suggest the same; however, if the critical values are the same as those computed by Stock and Yogo (2001) for the Cragg and Donald (1993) test, the bias on the second stage coefficient is approximately 25%, not enough to change the sign. The bias on the least squares coefficient cannot be confidently signed, but is probably towards zero (see above), so I conclude that the true coefficient is -0.3 or more negative, consistent with the hypothesis that having immigrant classmates reduces high school attainment.

The second row of Table 4 shows that the effect on natives' acquiring 12 years of schooling of immigrants aged 18–64 with less than 12 years of schooling is positive and statistically significant in every specification. Once trends are controlled for, the coefficient is in the range 0.74–0.91. This is consistent with the hypothesis that the presence of unskilled immigrants in the labor market alters the wage structure in such a way as to give natives an incentive to complete 12 years of schooling. A comparison of columns 7 and 8 shows that using 2SLS does not increase the coefficient greatly, despite the expectation it would be biased down in least squares. Yet column 3 of Table 5 shows that the predicted share of immigrants with less than 12 years' schooling in the population 18–64 is a strong predictor in the first stage, much stronger than the other excluded instruments, and the Angrist–Pischke F–statistic is very high.

The results of rows three and four of Table 4 indicate that the effects of adult immigrants with exactly 12 years of schooling and of adult immigrants with more than 12 years of schooling are sensitive to the specification and estimation method. The instruments associated with these covariates are fairly strong in their respective first stages (Table 5 columns 4 and 5).

I repeat the exercise of Table 4 for native non–Hispanic whites in Table 6. The least squares results are qualitatively similar to those for all natives, but the absolute values of the coefficients are smaller by about 0.2: the coefficient on immigrants aged 11–17 is in the range -0.16– -0.30. For this sample, however, 2SLS reduces the point estimate to essentially zero (c.f. columns 7 and 8), suggesting there is no negative effect on native non–Hispanic whites through the schooling channel (or that it is cancelled out by natives' anticipation of their immigrant classmates' future labor market effect). Nevertheless, the WLS and 2SLS coefficients are not statistically significantly different. The coefficients on more educated adult immigrants are qualitatively similar to those in Table 4.

In Table 7, I turn to native–born blacks. For this sample, the negative effect of immigrants 11–17 is robustly statistically significant and larger in absolute value than for non–Hispanic whites, ranging from -0.43– -0.60 in the preferred columns 4–8. The positive effect of adult immigrants with less than 12 years of education is also robust and larger than for non–Hispanic whites, ranging from 0.80–1.04 in columns 4–8. The effects of adult immigrants with exactly 12 years education and with more than 12 years are imprecisely estimated. I conclude that, as expected, the school quality and labor market channels through which immigration might operate are present for blacks and are larger than for non–Hispanic whites.

Finally, I examine native–born Hispanics in Table 8. While the sign of the coefficient on immigrants aged 11–17 is always negative, it is imprecisely measured in many specifications (first row). There is no evidence of a beneficial effect of adult immigrants with less than 12 years of education in the preferred columns of the second row. The results for Hispanics therefore appear to provide only weak support for the hypotheses being tested. In the next section, I show that this is because the Table 8 regressions are misspecified.

3.3 Distinguishing child immigrants according to parental characteristics

Distinguishing among immigrants aged 11–17 according to their parents' education proves helpful for interpreting the effects of immigration on native–born Hispanics, and these results are presented in columns 2 and 3 of Table 9. Column 1 repeats the preferred fixed effects specification of Table 8 (column 5). It is useful to concentrate first on the changes in the adult immigration coefficients. The coefficient on adult immigrants with less than 12 years of education rises from 0.1 in column 1 to a statistically significant 1.4 in column 2, in accordance with both theory and the results for native–born blacks. The coefficient is smaller in the differenced specification of column 3 (0.7), but statistically significant at the 10% level, and very different from the corresponding coefficient of -0.5 in Table 8 column 7. The coefficient on adult immigrants with more than 12 years of education falls from 2.8 in column 1 to 0.6–0.8 (and statistically insignificant) in columns 2 and 3, similar to the values for non–Hispanic native groups. This leaves only the coefficient on adult immigrants with exactly 12 years of education as slightly anomalous (-1.75 and -1.40 in columns 2 and 3).

It appears that in the specification with the undifferentiated child immigrant variable, the unskilled adult immigrant coefficient was capturing the opposing effects of the unskilled adult immigrants and their children: as anticipated, the coefficient on child immigrants with parents with less than 12 years of education is negative and statistically significant, with coefficients of -2.0 in columns 2 and 3. The coefficient on child immigrants with more educated parents is positive and statistically significant, with a relatively large coefficient of 3.2 in column 2 (2.5 in column 3); this effect was being picked up by the educated adult immigrant variable in the simpler specification.

The magnitude of these coefficients implies a one percentage point increase in the relevant immigrant share changes the 12 year completion rate by 2–3 percentage points. This compares to shares of immigrants in the 11–17 population of 3.3% with parents with less than 12 years schooling and 4.6% for those with a parent with at least 12 years schooling, and a Hispanic completion rate of 80.5%.[19] These numbers imply moderate elasticities of 0.08 for children of unskilled parents and 0.14–0.18 for children of more educated parents.

The coefficient on the share of the 11–17 year–old population that is immigrants with no parent in the household is strongly negative in both columns. There are several possible explanations for this. The first is causal: such immigrants, who tend to have very low enrollment rates, drop out early from school and have very negative peer effects in

[19] The shares of immigrants in the population 11–17 are computed using Hispanic weights, and are hence higher than implied in Figure 2. See Appendix Table 2.

the short time they are in school in the United States. If this is so, one might expect a similarly negative impact of children of the lowest educated immigrants. However, unreported regressions splitting parental dropouts into those with more or less than nine years of education do not point to clear differences between the two groups. Two other explanations are endogenous, and imply a negative bias on the coefficient. To the extent that the young immigrants move to the United States without their parents in order to work (Oropesa and Landale 2009), they are the immigrants most likely to move to states with (imperfectly controlled) favorable labor market conditions for youth, conditions which would incite the closely substitutable native students to drop out. To the extent that some young immigrants move to the United States with their parents and initially attend school, (imperfectly controlled) favorable youth labor market conditions are likely to entice both immigrants and natives to drop out of school to work, and hence also to leave home.

Since it is not possible to instrument all the immigrant variables in the expanded specifications of columns 2 and 3, it is not possible to make statements about causal effects. Nevertheless, Table 9 suggests that for Hispanic natives, just as for other natives, adult immigrants with less than 12 years schooling have a beneficial effect on completing 12 years of schooling, and that for Hispanic natives, classmates with educated parents have a beneficial effects, while classmates with poorly educated parents have detrimental effects. It is likely that the large effect of child immigrants with no parent in the household reflects endogeneity. It is not possible to know which group is responsible for the negative overall effect of immigration on native Hispanics found in Table 3 using 2SLS, but it seems likely to be adults with exactly 12 years schooling, or children of adults with less than 12 years schooling.

In Table 10, I present the corresponding regressions for native non–Hispanic whites (columns 1–3) and native–born blacks (columns 4–6). For these natives, there is no indication that grouping all child immigrants introduced misspecification, as the coefficients on the adult immigrant variables do not change much when the child immigrant variable is split into components. Neither is there a clear distinction between the effects of child immigrants with lower and higher education parents, and indeed, these coefficients are all statistically insignificant. For whites, I interpret the generally small coefficients as consistent with the finding of no effect of child immigrants in the 2SLS of Table 6 column 8. For blacks the point estimates are larger, even if insignificant, and I interpret them as consistent with the finding of a statistically significant negative effect of child immigrants in the 2SLS of Table 7 column 8.

The coefficient on child immigrants with no parent in the household is not nearly as negative in Table 10 as for native Hispanics (and is positive for blacks in column 6). Nevertheless, for non–Hispanic whites, it is responsible for the statistically significant negative sign when all child immigrants are grouped in column 1 or in the differenced equivalent (as an unreported regressions show). Also, if the Hispanic weights are used for blacks and non–Hispanic whites, the coefficient on child immigrants with no parent in the household becomes much more negative (other coefficients are less sensitive), though less negative than for Hispanics. In recent decades, which are most heavily weighted in the regressions, such child immigrants are very disproportionately Mexican–born, which may explain why their "effect", whether causal or endogenous, is largest when Hispanic states are weighted more or for native–born Hispanics: within a state, the immigrants may be more concentrated in either schools or labor markets with many native–born Hispanics, and they may be closer labor market substitutes for young native–born Hispanics than for other young natives. For blacks and non–Hispanic whites, 2SLS in Tables 6 and 7 should address the endogenous behavior of these immigrants.

4 Conclusion

In this paper, I have shown that natives' probability of completing 12 years of education is increased by immigration, albeit by a small magnitude. An increase of one percentage point in the share of immigrants in the population aged 11–64 increases the probability natives complete 12 years of schooling by 0.1–0.2 percentage points, and increases the probability for blacks by 0.4–0.5 percentage points. The effects are rather small compared to the average native completion rate of 86.0% (80.5% for blacks) and given the average immigrant share of 8.9% (8.1% for blacks). For native–born Hispanics, by contrast, 2SLS results suggest a negative effect of 0.2 percentage points, though it is imprecisely estimated.

The net positive effect for natives generally and blacks is the result of the additional incentive to complete 12 years provided to natives 11–17 by the presence of unskilled adult immigrants in the labor market, offset by the negative effect of the presence of immigrant children also aged 11–17 who are possibly classmates. I find that a one percentage point increase in the share of immigrants in the population aged 11–17 reduces the probability natives eventually complete 12 years of school by 0.3–0.5 percentage points for all natives and by 0.4–0.6 percentage points for blacks, though not at all for native non–Hispanic whites. On the other hand, a one percentage point increase in the share of immigrants with less than 12 years school in the population aged 18–64 increases the eventual native completion rate by 0.7–0.9 percentage points for natives generally, and by 0.8–1.0 percentage points for blacks. Effects of more educated adult immigrants are generally insignificant.

The mechanisms for native–born Hispanics are more complex. While the presence of unskilled adult immigrants provides an incentive for high school completion similar in magnitude to that for native–born blacks, the effect of child immigrants is much larger, and strongly dependent on the education of the child immigrants' parents: children of parents with less than 12 years of education have a deleterious impact on native completion rates, while children with a parent with 12 or more years of education have a beneficial effect. These results are subject to the caveat that 2SLS could not be employed in the necessary more complex specification.

The results suggest the need for reform in accommodating immigrant students, particularly those with less educated parents and in schools with many native–born minorities, and help explain how natives respond to immigration in such as way as to obviate the potential negative impact on native wages.

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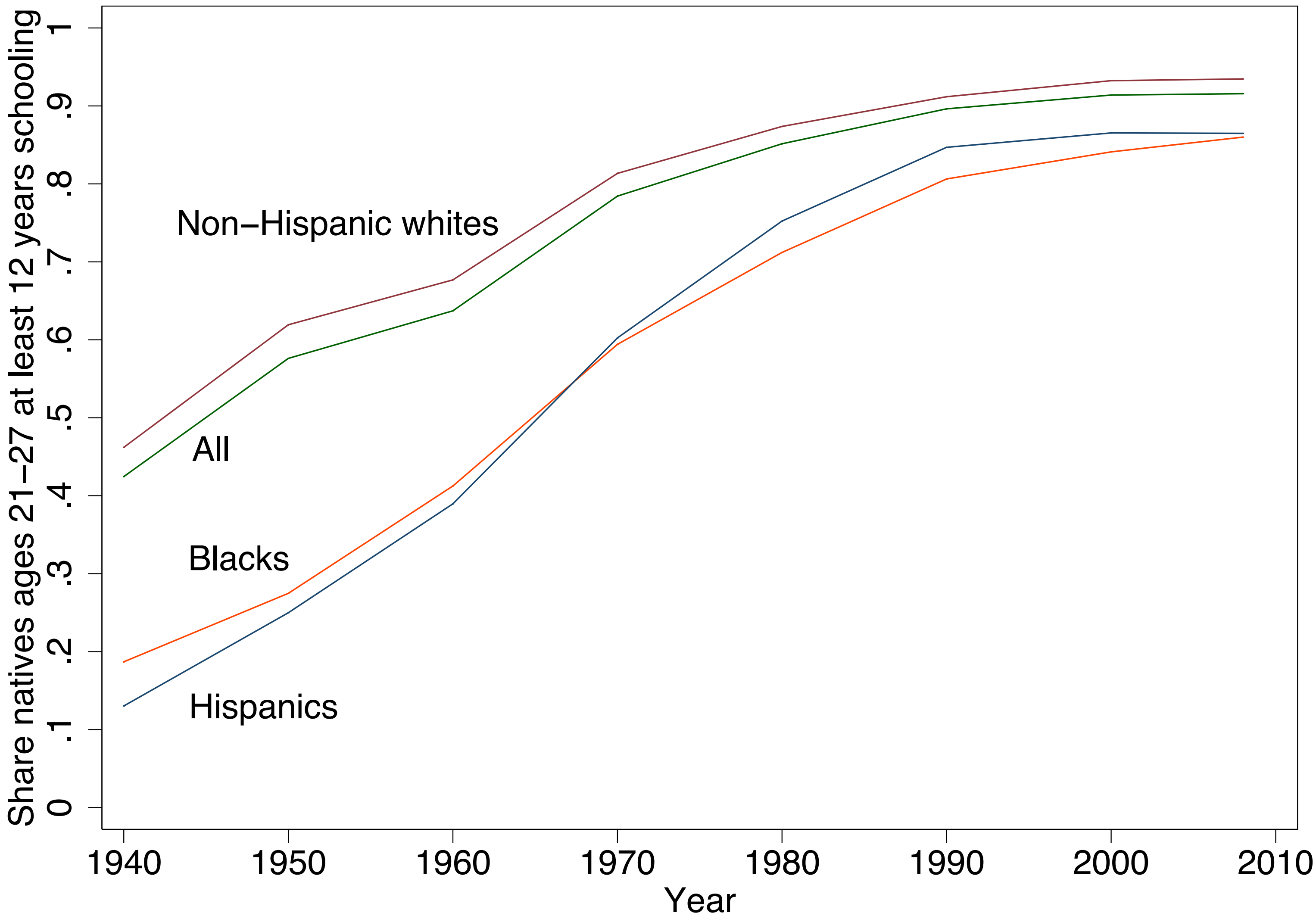
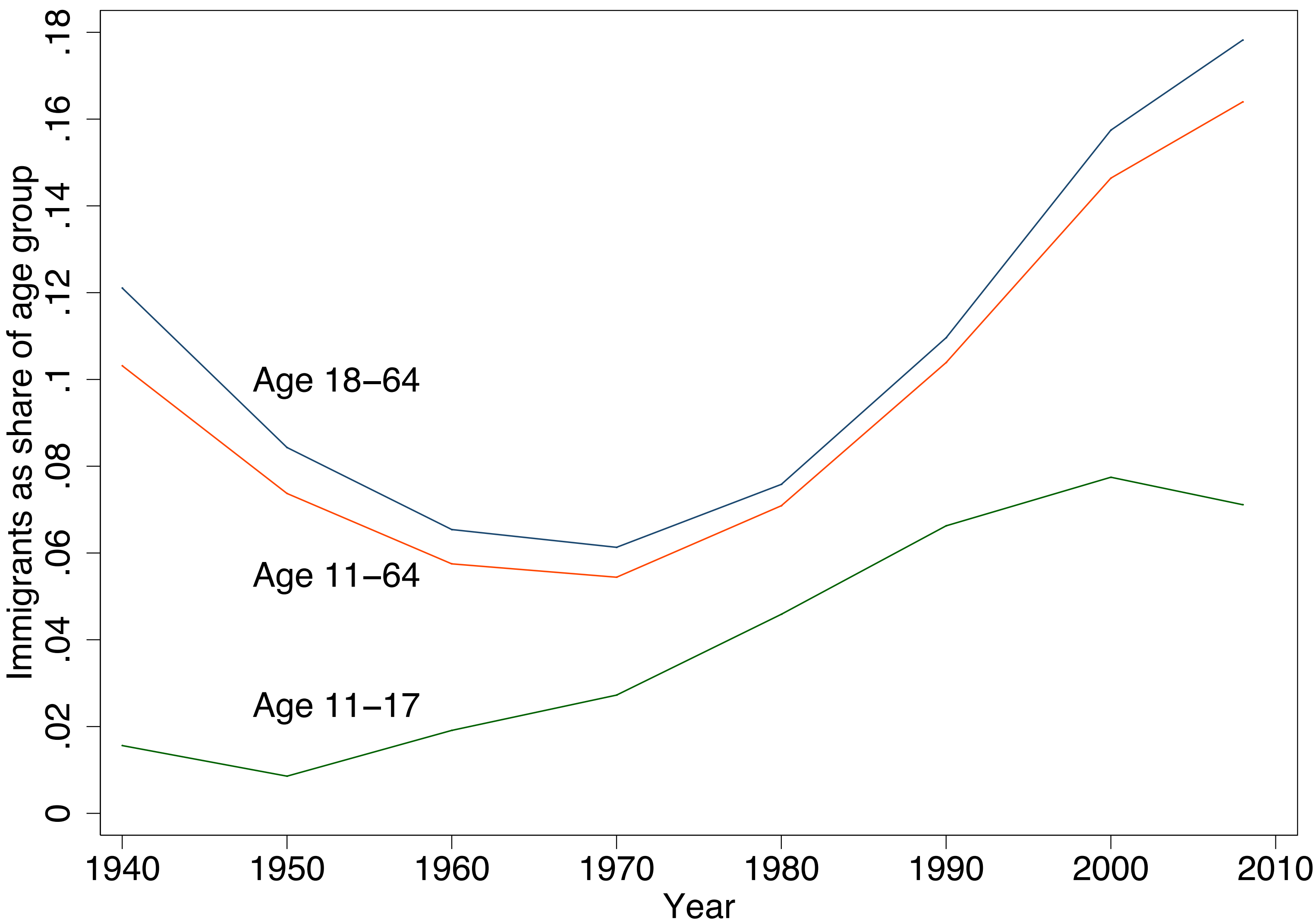
Figure 1: Share of natives with at least 12 years of schooling, by race and ethnicity



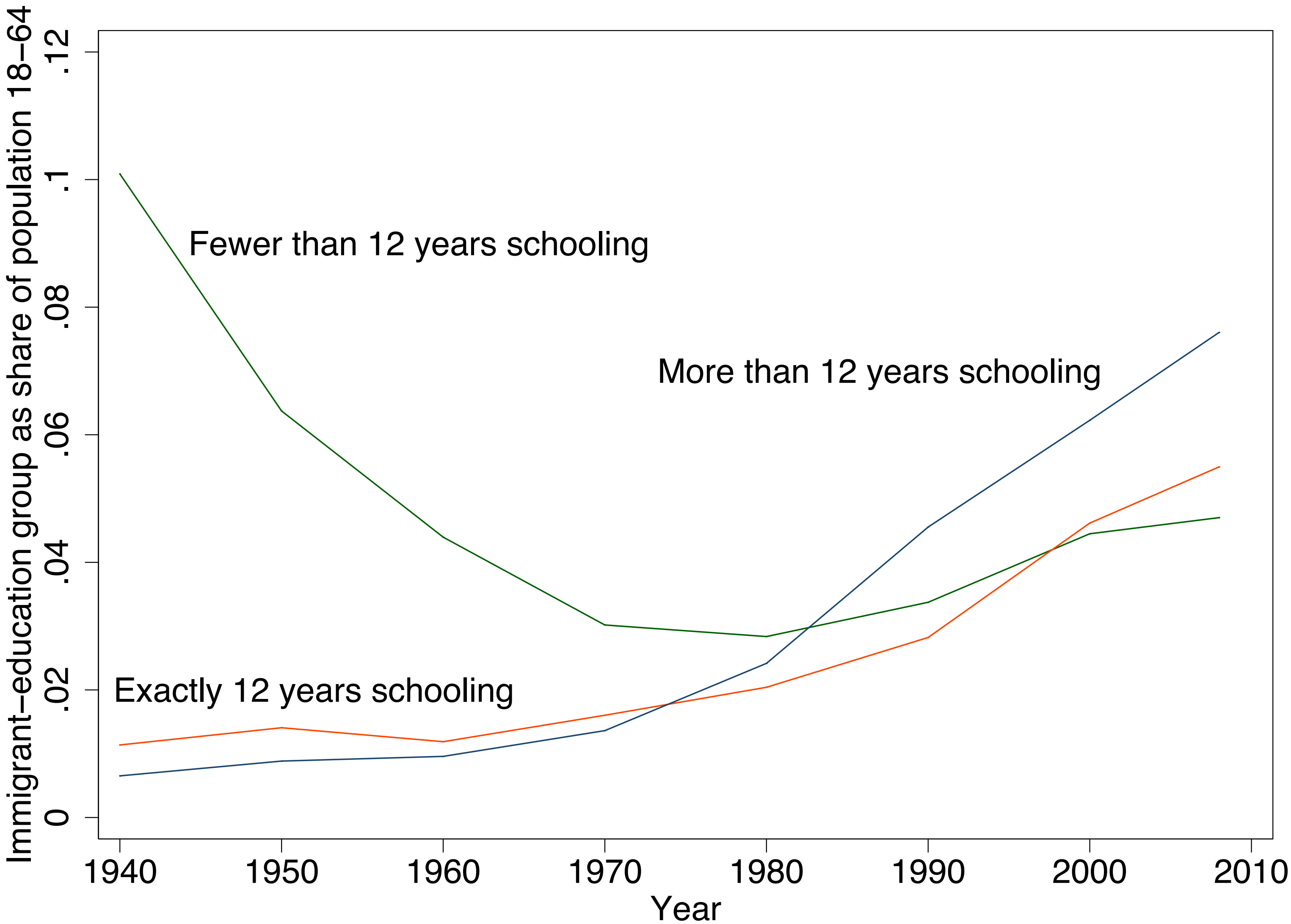
Note: The share of natives aged 21–27 who have completed at least 12 years of schooling.
Source: U.S. Census 1940–2000, American Community Survey 2006–2008.

Figure 2: Immigrant share in various age groups



Note: Immigrants as a share of each age group.
Source: U.S. Census 1940–2000, American Community Survey 2006–2008.

Figure 3: Immigrant education groups as share of the population, ages 18–64



Note: Immigrants 18–64 with various education levels as share of the total population aged 18–64.
Source: U.S. Census 1940–2000, American Community Survey 2006–2008.

Table 1: Effects of immigrants in population 11-64 on native probability of completing 12 years education

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Fixed effects						10-year differences	
							WLS	2SLS
Share pop 11-64 which is immigrant; t-10	0.04 (0.13)	0.07 (0.08)	0.11 (0.10)	0.16^{*} (0.08)	0.21^{**} (0.07)	0.14^{*} (0.07)	0.18^{**} (0.06)	0.23^{*} (0.13)
Unemployment rate age 18-24; t-10	--	0.65^{*} (0.37)	-0.04 (0.18)	-0.11 (0.18)	-0.15 (0.17)	-0.10 (0.17)	-0.11 (0.11)	-0.12 (0.11)
Unemployment rate age 25-54; t-10	--	-0.59 (0.79)	0.18 (0.38)	0.30 (0.39)	0.39 (0.39)	0.36 (0.37)	0.37 (0.26)	0.39 (0.25)
Share of native population which is age 11-17; t-10	--	-1.33^{**} (0.36)	-0.35 (0.25)	-0.28 (0.24)	-0.23 (0.25)	0.06 (0.23)	0.05 (0.18)	0.06 (0.18)
Share workers in agriculture 1940*year	--	--	--	0.007^{*} (0.004)	0.006^{*} (0.004)	-0.004 (0.003)	0.052 (0.033)	0.056^{*} (0.032)
Share white natives 21-27 less than 12 years school 1940*year	--	--	--	--	0.007^{**} (0.003)	0.006^{**} (0.003)	0.064^{**} (0.027)	0.068^{**} (0.027)
Log personal income per capita; t-10	--	--	--	--	--	0.093^{**} (0.025)	--	--
BEA regions*year (p-value)	--	--	0.00	0.00	0.00	0.00	0.00	0.00
R^2	0.91	0.92	0.97	0.97	0.97	0.98	0.85	--
Observations	343						294	

Notes: The dependent variable is the share of natives ages 21-27 who have completed 12 years of education, adjusted at the individual level for age and sex. Estimation is by weighted least squares columns 1-7, and two-stage least squares column 8, with weights w the inverse of the squared standard errors on the state-year interaction coefficient in the individual regression for columns 1-6, and $1/(1/w_t+1/w_{t+10})$ for columns 7 and 8. All specifications include year dummies; fixed effects specifications also include state dummies. The dependent variable is based on 1950-2008 data, the independent variables on data from 1940-2008. The instrument in column 8 is based on the 1940 distribution of immigrants from different regions (see text). Standard errors are clustered by state and reported in parentheses.

Table 2: Effects of immigrants in population 11-64 on native probability of completing 12 years education, sensitivity to sample

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	All	1970-80 onwards		Without 2008		Females		Males	
	2SLS	WLS	2SLS	WLS	2SLS	WLS	2SLS	WLS	2SLS
Share pop 11-64 which is immigrant; t-10	0.23*	0.15**	0.17*	0.19**	0.11	0.19**	0.25*	0.17**	0.20
	(0.13)	(0.06)	(0.12)	(0.08)	(0.13)	(0.06)	(0.14)	(0.05)	(0.13)
Other covariates	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	--	0.76	--	0.87	--	0.82	--	0.85	--
Observations	294	196		245		294		294	
1st stage instrument coefficient	0.28**	--	0.22**	--	0.41**	--	0.28**	--	0.28**
	(0.05)		(0.05)		(0.07)		(0.05)		(0.05)
1st stage instrument F-statistic	25.2	--	20.8	--	37.1	--	25.2	--	25.2
1st stage partial R^2	0.10	--	0.11	--	0.18		0.10	--	0.10

Notes: The dependent variable is the share of natives ages 21-27 who have completed 12 years of education, adjusted at the individual level for age and (columns 1-4) sex. Estimation is on the 10-year difference of the data; WLS indicates weighted least squares estimation; 2SLS indicates two-stage least squares estimation; both with weights $1/(1/w_t+1/w_{t+10})$, where is the weight described in Table 1. All specifications include year dummies. The instrument is based on the 1940 distribution of immigrants from different regions (see text). Standard errors are clustered by state and reported in parentheses. Other covariates are those in Table 1: unemployment rates, cohort size, BEA region trends, 1940 agriculture share trend, 1940 share white natives with less than 12 years education trend.

Table 3: Effects of immigrants in population 11-64 on native probability of completing 12 years education, by race and ethnicity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Native-born sample			Fixed effects				10-year differences	
							WLS	2SLS
All	0.04	0.07	0.11	0.16*	0.21**	0.14*	0.18**	0.23*
	(0.13)	(0.08)	(0.10)	(0.08)	(0.07)	(0.07)	(0.06)	(0.13)
Non-Hispanic whites	-0.00	0.01	0.07	0.08	0.14*	0.12	0.14**	0.18
	(0.08)	(0.06)	(0.10)	(0.09)	(0.07)	(0.08)	(0.07)	(0.14)
Blacks	0.04	0.12	0.21	0.48**	0.51**	0.44**	0.44**	0.43**
	(0.15)	(0.11)	(0.15)	(0.10)	(0.09)	(0.11)	(0.10)	(0.19)
Hispanics	0.11	0.21	0.42**	0.50**	0.52**	0.32**	0.27*	-0.19
	(0.25)	(0.18)	(0.20)	(0.19)	(0.19)	(0.13)	(0.14)	(0.23)
Unemployment rates; t-10	--	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Share of native pop which is age 11-17; t-10	--	Yes	Yes	Yes	Yes	Yes	Yes	Yes
BEA regions*year	--	--	Yes	Yes	Yes	Yes	Yes	Yes
Share workers in agriculture 1940*year	--	--	--	Yes	Yes	Yes	Yes	Yes
Share white natives 21-27 less than 12 years school 1940*year	--	--	--	--	Yes	Yes	Yes	Yes
Log personal income p.c.; t-10	--	--	--	--	--	Yes	--	--
Observations (all/whites/blacks/Hispanics)			343/343/325/332				294/294/272/283	

Notes: The dependent variable is the share of natives ages 21-27 of a racial/ethnic group who have completed 12 years of education, adjusted at the individual level for age and sex. Estimation is by weighted least squares columns 1-7, and two-stage least squares column 8, with weights w the inverse of the squared standard errors on the state-year interaction coefficient in the individual regression for columns 1-6, and $1/(1/w_t+1/w_{t+10})$ for columns 7 and 8. All specifications include year dummies; fixed effects specifications also include state dummies. The dependent variable is based on 1950-2008 data, the independent variables on data from 1940-2000. The instrument in column 8 is based on the 1940 distribution of immigrants from different regions (see text). Standard errors are clustered by state and reported in parentheses. Each coefficient is from a different regression.

Table 4: Effects of immigrants by age and education on native probability of completing 12 years education

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
			Fixed effects				10-year differences	
							WLS	2SLS
Share pop 11-17 which is immigrant; t-10	-0.61**	-0.67**	-0.46**	-0.45**	-0.44**	-0.39**	-0.33**	-0.32
	(0.21)	(0.16)	(0.11)	(0.11)	(0.11)	(0.11)	(0.08)	(0.21)
Share pop 18-64 which is immigrant < 12 yrs edu; t-10	0.94**	1.01**	0.94**	0.91**	0.86**	0.80**	0.74**	0.77**
	(0.23)	(0.16)	(0.18)	(0.17)	(0.15)	(0.15)	(0.15)	(0.19)
Share pop 18-64 immigrant 12 years edu; t-10	0.46	0.63	-0.46	-0.45	-0.51*	-0.46*	-0.18	0.23
	(0.60)	(0.56)	(0.38)	(0.37)	(0.28)	(0.28)	(0.25)	(0.79)
Share pop 18-64 immigrant > 12 years edu; t-10	-0.23	-0.23	0.53*	0.57*	0.73**	0.61*	0.45*	0.02
	(0.62)	(0.48)	(0.32)	(0.32)	(0.28)	(0.31)	(0.23)	(0.76)
Unemployment rates; t-10	--	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Share of native pop which is age 11-17; t-10	--	Yes	Yes	Yes	Yes	Yes	Yes	Yes
BEA regions*year	--	--	Yes	Yes	Yes	Yes	Yes	Yes
Share workers in agriculture 1940*year	--	--	--	Yes	Yes	Yes	Yes	Yes
Share non-Hispanic whites <12 years education 1940*year	--	--	--	--	Yes	Yes	Yes	Yes
Log personal income p.c.; t-10	--	--	--	--	--	Yes	--	--
R^2	0.94	0.95	0.98	0.98	0.98	0.98	0.87	--
Observations			343				294	

Notes: The dependent variable is the share of natives ages 21-27 who have completed 12 years of education, adjusted at the individual level for age and sex. Estimation is by weighted least squares columns 1-7, and two-stage least squares column 8, with weights w the inverse of the squared standard errors on the state-year interaction coefficient in the individual regression for columns 1-6, and $1/(1/w_t+1/w_{t+10})$ for columns 7 and 8. All specifications include year dummies; fixed effects specifications also include state dummies. The dependent variable is based on 1950-2008 data, the independent variables on data from 1940-2000. The instruments in column 8 are based on the 1940 distribution of immigrants from different regions (see text). Standard errors are clustered by state and reported in parentheses.

Table 5: First stage of two-stage least squares

	(1)	(2)	(3)	(4)	(5)
	Share of immigrants in age group		Share of population aged 18-64 which is immigrant with		
	11-64	11-17	less than 12 years education	exactly 12 years education	more than 12 years education
Predicted share pop 11-64 which is immigrant	0.28** (0.05)	--	--	--	--
Predicted share pop 11-17 which is immigrant	--	0.70** (0.26)	-0.02 (0.07)	-0.03 (0.02)	0.06 (0.06)
Predicted share pop 18-64 which is immigrant less than 12 years edu;	--	-0.02 (0.09)	0.73** (0.07)	0.04 (0.04)	0.10 (0.07)
Predicted share pop 18-64 which is immigrant 12 years edu	--	-1.26** (0.45)	-0.81** (0.21)	0.23** (0.08)	-0.23 (0.13)
Predicted share pop 18-64 which is immigrant more than 12 years edu	--	0.63** (0.22)	0.43** (0.11)	0.09 (0.06)	0.43** (0.08)
F-test for joint significance	25.2	9.5	71.5	36.0	26.0
Angrist-Pischke F-test	--	5.6	91.4	12.7	12.0
Partial R^2	0.10	0.26	0.14	0.10	0.14
Observations	294	294	294	294	294

Notes: Estimation is by weighted least squares on 10-year differenced data, with weights $1/(1/h_t+1/h_{t+10})$, where h is the denominator of the dependent variable in each regression: population age 11-64 in column 1, population age 11-17 in column 2, population age 18-64 with less than 12 years education column 3, population age 18-64 with exactly 12 years education column 3, population age 18-64 with more than 12 years education column 5. All specifications include year dummies and the (differenced) non-immigrant covariates of Table 1 column 8. The excluded instruments (predicted shares) are based on historical patterns of settlement by age and education. The dependent variable and excluded instruments are based on data from 1940-2000 (i.e. from t-10), while most other covariates are based on data from 1950-2008. Standard errors are clustered by state and reported in parentheses.

Table 6: Effects of immigrants by age and education on native non-Hispanic whites' probability of completing 12 years education

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
			Fixed effects				10-year differences	
							WLS	2SLS
Share pop 11-17 which is immigrant; t-10	-0.51**	-0.52**	-0.29**	-0.30**	-0.28**	-0.27**	-0.16**	-0.04
	(0.17)	(0.15)	(0.10)	(0.10)	(0.10)	(0.11)	(0.07)	(0.19)
Share pop 18-64 which is immigrant < 12 yrs edu; t-10	0.61**	0.63**	0.59**	0.60**	0.53**	0.51**	0.45**	0.41**
	(0.17)	(0.13)	(0.17)	(0.18)	(0.16)	(0.17)	(0.15)	(0.17)
Share pop 18-64 immigrant 12 years edu; t-10	0.29	0.41	-0.48	-0.48	-0.55**	-0.53**	-0.26	0.41
	(0.49)	(0.45)	(0.37)	(0.37)	(0.22)	(0.21)	(0.18)	(0.63)
Share pop 18-64 immigrant > 12 years edu; t-10	-0.05	-0.09	0.44	0.43	0.65**	0.62**	0.38*	-0.13
	(0.48)	(0.42)	(0.37)	(0.32)	(0.26)	(0.29)	(0.22)	(0.67)
Unemployment rates; t-10	--	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Share of native pop which is age 11-17; t-10	--	Yes	Yes	Yes	Yes	Yes	Yes	Yes
BEA regions*year	--	--	Yes	Yes	Yes	Yes	Yes	Yes
Share workers in agriculture 1940*year	--	--	--	Yes	Yes	Yes	Yes	Yes
Share white natives 21-27 less than 12 years school 1940*year	--	--	--	--	Yes	Yes	Yes	Yes
Log personal income p.c.; t-10	--	--	--	--	--	Yes	--	--
R^2	0.95	0.95	0.98	0.98	0.98	0.98	0.88	--
Observations			343				294	

Notes: The dependent variable is the share of native-born non-Hispanic whites ages 21-27 who have completed 12 years of education, adjusted at the individual level for age and sex. Estimation is by weighted least squares columns 1-7, and two-stage least squares column 8, with weights w the inverse of the squared standard errors on the state-year interaction coefficient in the individual regression for columns 1-6, and $1/(1/w_t+1/w_{t+10})$ for columns 7 and 8. All specifications include year dummies; fixed effects specifications also include state dummies. The dependent variable is based on 1950-2008 data, the independent variables on data from 1940-2000. The instruments in column 8 are based on the 1940 distribution of immigrants from different regions (see text). Standard errors are clustered by state and reported in parentheses.

Table 7: Effects of immigrants by age and education on native blacks' probability of completing 12 years education

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
			Fixed effects				10-year differences	
							WLS	2SLS
Share pop 11-17 which is immigrant; t-10	-0.86**	-0.87**	-0.72**	-0.60**	-0.59**	-0.60**	-0.43**	-0.46**
	(0.32)	(0.22)	(0.19)	(0.20)	(0.20)	(0.20)	(0.18)	(0.22)
Share pop 18-64 which is immigrant <12 yrs edu; t-10	0.69**	0.76**	1.21**	1.01**	0.99**	1.02**	0.80**	1.04**
	(0.33)	(0.22)	(0.26)	(0.24)	(0.21)	(0.25)	(0.21)	(0.31)
Share pop 18-64 immigrant 12 yrs edu; t-10	-0.18	0.27	0.08	0.03	-0.04	-0.04	0.42	1.51
	(0.70)	(0.78)	(0.74)	(0.77)	(0.72)	(0.72)	(0.61)	(1.30)
Share pop 18-64 immigrant > 12 yrs edu; t-10	0.74	0.65	0.65	1.11	1.21*	1.24*	0.78	-0.46
	(0.75)	(0.71)	(0.63)	(0.69)	(0.68)	(0.69)	(0.58)	(1.21)
Unemployment rates; t-10	--	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Share native pop age 11-17; t-10	--	Yes	Yes	Yes	Yes	Yes	Yes	Yes
BEA regions*year	--	--	Yes	Yes	Yes	Yes	Yes	Yes
Share workers in agriculture 1940*year	--	--	--	Yes	Yes	Yes	Yes	Yes
Share white natives 21-27 less than 12 years school 1940*year	--	--	--	--	Yes	Yes	Yes	Yes
Log personal income p.c.; t-10	--	--	--	--	--	Yes	--	--
R^2	0.96	0.97	0.97	0.98	0.98	0.98	0.86	--
Observations			326				273	

Notes: The dependent variable is the share of native-born blacks age 21-27 who have completed 12 years of education, adjusted at the individual level for age and sex. Estimation is by weighted least squares columns 1-7, and two-stage least squares column 8, with weights w the inverse of the squared standard errors on the state-year interaction coefficient in the individual regression for columns 1-6, and $1/(1/w_t+1/w_{t+10})$ for columns 7 and 8. All specifications include year dummies; fixed effects specifications also include state dummies. The dependent variable is based on 1950-2008 data, the independent variables on data from 1940-2000. The instruments in column 8 are based on the 1940 distribution of immigrants from different regions (see text). Standard errors are clustered by state and reported in parentheses.

Table 8: Effects of immigrants by age and education on native Hispanics' probability of completing 12 years education

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
			Fixed effects				10-year differences	
							WLS	2SLS
Share pop 11-17 which is immigrant t-10	-0.37 (0.27)	-0.60* (0.20)	-0.76** (0.26)	-0.63** (0.25)	-0.57** (0.25)	-0.38 (0.31)	-0.39* (0.22)	-1.04** (0.42)
Share pop 18-64 which is immigrant < 12 yrs edu; t-10	0.91* (0.47)	1.33** (0.48)	0.88* (0.47)	0.32 (0.41)	0.14 (0.44)	-0.05 (0.44)	-0.52 (0.40)	-0.17 (0.63)
Share pop 18-64 immigrant 12 years edu; t-10	1.77 (1.23)	0.67 (1.10)	-1.77* (0.90)	-1.21 (0.87)	-1.21 (0.84)	-1.02 (0.93)	-0.77 (0.90)	-4.97** (1.70)
Share pop 18-64 immigrant > 12 years edu; t-10	-1.18 (1.16)	-0.29 (0.99)	2.34** (0.85)	2.66** (0.79)	2.77** (0.74)	2.21** (0.86)	2.56** (0.87)	4.66** (1.41)
Unemployment rates; t-10	--	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Share of native pop which is age 11-17; t-10	--	Yes	Yes	Yes	Yes	Yes	Yes	Yes
BEA region trends	--	--	Yes	Yes	Yes	Yes	Yes	Yes
Share agricultural workers 1940*year	--	--	--	Yes	Yes	Yes	Yes	Yes
Share white natives 21-27 less than 12 years school 1940*year	--	--	--	--	Yes	Yes	Yes	Yes
Log personal income p.c.; t-10	--	--	--	--	--	Yes	--	--
R^2	0.91	0.91	0.94	0.94	0.94	0.04	0.59	--
Observations			332				283	

Notes: The dependent variable is the share of native—born Hispanics age 21-27 who have completed 12 years of education, adjusted at the individual level for age and sex. Estimation is by weighted least squares columns 1-7, and two-stage least squares column 8, with weights w the inverse of the squared standard errors on the state-year interaction coefficient in the individual regression for columns 1-6, and $1/(1/w_t+1/w_{t+10})$ for columns 7 and 8. All specifications include year dummies; fixed effects specifications also include state dummies. The dependent variable is based on 1950-2008 data, the independent variables on data from 1940-2008. The instruments in column 8 are based on the 1940 distribution of immigrants from different regions (see text). Standard errors are clustered by state and reported in parentheses.

Table 9: Effects of youth immigrants by parental education on native Hispanics' probability of completing 12 years education

	Fixed effects		10-year differences
	(1)	(2)	(3)
Share population 11-17	-0.57**	--	--
which is immigrant; t-10	(0.25)		
Parents less than 12 years edu	--	-2.04**	-1.93**
		(0.64)	(0.56)
One parent 12 or more years ed	--	3.17**	2.49**
		(0.76)	(0.67)
No parent in household	--	-4.77**	-3.07*
		(1.66)	(1.71)
Share population 18-64			
which is immigrant; t-10			
Less than 12 years education	0.14	1.38**	0.72*
	(0.44)	(0.40)	(0.39)
12 years education	-1.21	-1.75**	-1.40*
	(0.84)	(0.59)	(0.81)
More than 12 years education	2.77**	0.58	0.81
	(0.74)	(0.57)	(0.70)
Other covariates	Yes	Yes	Yes
R^2	0.94	0.95	0.62
Observations	332		283

Notes: The dependent variable is the share of native—born Hispanics age 21-27 who have completed 12 years of education, adjusted at the individual level for age and sex. Estimation is by weighted least squares, with weights w the inverse of the squared standard errors on the state-year interaction coefficient in the individual regression for columns 1 and 2, and $1/(1/w_t+1/w_{t+10})$ for column 2. All specifications include year dummies; fixed effects specifications also include state dummies. Other covariates are those in Table 1: unemployment rates, cohort size, BEA region trends, 1940 agriculture share trend, 1940 share white natives with less than 12 years education trend. The dependent variable is based on 1950-2008 data, the independent variables on data from 1940-2000. Standard errors are clustered by state and reported in parentheses.

Table 10: Effects of youth immigrants by parental education on native blacks' and non-Hispanic whites' probability of completing 12 years education

	Native non-Hispanic whites			Native blacks		
	Fixed effects		10-year diffs	Fixed effects		10-year diffs
	(1)	(2)	(3)	(4)	(5)	(6)
Share population 11-17	-0.28**	--	--	-0.59**	--	
which is immigrant; t-10	(0.10)			(0.20)		
Parents less than 12 years edu	--	0.12	-0.02	--	-0.87	-0.80
		(0.18)	(0.15)		(0.67)	(0.65)
One parent 12 or more years	--	-0.29	-0.08	--	-0.20	-0.61
		(0.29)	(0.20)		(0.71)	(0.55)
No parent in household	--	-1.72**	-0.86	--	-0.60	1.10
		(0.56)	(0.52)		(1.49)	(1.82)
Share population 18-64						
which is immigrant; t-10						
Less than 12 years education	0.53**	0.51**	0.46**	0.99**	1.09**	0.77**
	(0.16)	(0.23)	(0.17)	(0.21)	(0.31)	(0.28)
12 years education	-0.55**	-0.55**	-0.27	-0.04	-0.13	0.36
	(0.22)	(0.23)	(0.20)	(0.72)	(083)	(0.70)
More than 12 years edu	0.65**	0.75**	0.38	1.21*	0.98	0.83
	(0.26)	(0.29)	(0.26)	(0.68)	(0.69)	(0.52)
Other covariates	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.98	0.98	0.88	0.98	0.98	0.86
Observations	343		294	326		272

Notes: The dependent variable is the share of native—born non-white Hispanics (columns 1-3) or blacks (columns 4-6) age 21-27 who have completed 12 years of education, adjusted at the individual level for age and sex. Estimation is by weighted least squares, with weights w the inverse of the squared standard errors on the state-year interaction coefficient in the individual regression for columns 1, 2, 4 and 6 and $1/(1/w_t+1/w_{t+10})$ for columns 3 and 6. All specifications include year dummies; fixed effects specifications also include state dummies. Other covariates are those in Table 1: unemployment rates, cohort size, BEA region trends, 1940 agriculture share trend, 1940 share white natives with less than 12 years education trend. The dependent variable is based on 1950-2008 data, the independent variables on data from 1940-2000. Standard errors are clustered by state and reported in parentheses.

Appendix Table 1: Means of individual level variables

	(1) 12 or more years education completed	(2) Female	(3) Age	(4) Observations
A. All				
1950	0.576	0.49	24.0 (2.0)	47,488
2008	0.911	0.49	21.5 (1.8)	492,267
1950-2008	0.860	0.51	23.6 (2.2)	4,067,316
B. Non-Hispanic whites				
1950	0.619	0.49	24.0 (2.0)	41,151
2008	0.933	0.49	21.5 (1.8)	347,252
1950-2008	0.892	0.50	23.7 (2.2)	3,187,014
C. Blacks				
1950	0.250	0.53	24.1 (2.0)	4967
2008	0.850	0.50	21.4 (1.8)	58,569
1950-2008	0.805	0.53	23.5 (2.2)	455,617
D. Hispanics				
1950	0.275	0.52	23.9 (1.9)	969
2008	0.858	0.50	21.4 (1.8)	55,845
1950-2008	0.805	0.51	23.2 (2.2)	252,294

Notes: Weighted with census weights adjusted so that the sum of weights for each year reflects the sample size of the census in that year. 2008 refers to the pooled 2006-2008 ACSs. The sample contains natives aged 21-27, except for 2006-2008, when they are aged 19-25. Standard deviations are in parentheses.

Appendix Table 2: Means of state-level covariates

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Non-Hispanic white weights			Black weights			Hispanic weights		
	1940	2000	1940-2000	1940	2000	1940-2000	1940	2000	1940-2000
Unemployment rate	0.162	0.100	0.105	0.100	0.108	0.106	0.154	0.108	0.104
ages 18-24	(0.053)	(0.018)	(0.030)	(0.041)	(0.018)	(0.026)	(0.038)	(0.015)	(0.022)
Unemployment rate	0.071	0.035	0.046	0.053	0.037	0.045	0.075	0.038	0.043
ages 25-54	(0.019)	(0.007)	(0.014)	(0.015)	(0.008)	(0.012)	(0.015)	(0.006)	(0.010)
Share native population	0.140	0.105	0.113	0.150	0.104	0.112	0.138	0.111	0.111
which is aged 11-17	(0.011)	(0.007)	(0.016)	(0.009)	(0.008)	(0.017)	(0.011)	(0.008)	(0.014)
Share employment in	0.092	0.093	0.090	0.167	0.107	0.116	0.110	0.081	0.086
agriculture 1940	(0.068)	(0.062)	(0.063)	(0.064)	(0.008)	(0.075)	(0.047)	(0.051)	(0.052)
Share native whites with	0.547	0.530	0.532	0.626	0.565	0.574	0.512	0.487	0.489
<12 years education 1940	(0.086)	(0.095)	(0.092)	(0.069)	(0.094)	(0.094)	(0.074)	(0.083)	(0.082)
State personal income per	583	29,490	16,174	380	29,140	17,037	535	30,720	20,398
capita (nominal)	(195)	(4216)	(9662)	(160)	(4667)	(9804)	(180)	(3972)	(9758)
Share population 11-17									
which is immigrant; t-10									
Parents high s dropouts	0.011	0.017	0.015	0.007	0.016	0.013	0.015	0.036	0.033
	(0.007)	(0.017)	(0.018)	(0.004)	(0.016)	(0.017)	(0.013)	(0.019)	(0.022)
One parent high s grad	0.003	0.041	0.028	0.001	0.043	0.030	0.003	0.058	0.046
	(0.002)	(0.025)	(0.023)	(0.001)	(0.027)	(0.025)	(0.002)	(0.023)	(0.026)
No parent in household	0.001	0.006	0.004	0.001	0.006	0.004	0.002	0.010	0.009
	(0.001)	(0.004)	(0.005)	(0.001)	(0.004)	(0.005)	(0.001)	(0.004)	(0.006)
Observations	49	49	343	45	49	321	47	49	330

Notes: The weights are the inverse of the squared standard errors on the state-year interaction coefficient in the individual regression for attainment of 12 years of education, by race/ethnicity. The individual regressions are weighted with census weights adjusted so that the sum of weights for each year reflects the sample size of the census in that year.

Appendix Table 3: 1940 shares of national-level immigrants from various origins, all ages and educations

Origin	(1) Share
United Kingdom	0.050
Ireland	0.059
Italy	0.062
Germany	0.052
Poland	0.060
Russia	0.064
Other Europe	0.050
Mexico	0.043
Puerto Rico	0.096
Canada	0.038
Central America	0.052
South America	0.069
Other Caribbean	0.073
Cuba	0.047
China	0.058
India	0.047
Other Asia	0.045
Rest of world	0.054