Age at Immigration and the Educational Attainment of Foreign-born Children in the United States: The Confounding Effects of Parental Education *

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Abstract

A substantial empirical literature confirms an educational disadvantage for foreign-born children that arrive in their host countries at older ages. However, parental characteristics are not randomly distributed across the age at immigration distribution. In the presence of a negative correlation between parental education and age at immigration, estimates of the educational attainment-age at immigration gradient, neglecting controls for parental education, will tend to over-estimate this disadvantage. Exploiting the large sample size in the 2000 census, the results indicate a considerable over-estimation (up to almost 28%) of the disadvantage for immigrant children that arrive at older ages, consistent with a negative correlation between parental education and age at immigration. Moreover, the results presented in the paper provide two additional insights: (1) a considerable portion (66%) of the educational advantage for early arrivals arises from the self-selection of immigrant parents reflecting their favourable unobserved characteristics, relative to their peers in their source country and (2) the educational disadvantage for late arriving immigrant children might be eliminated, provided their immigrant parents achieve sufficiently high levels of educational attainment.

Keywords: Education, immigrant children, age at immigration, parental education
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1 Introduction

A substantial empirical literature, such as Schaafsma & Sweetman (2001), Gonzalez (2003), Bleakley & Chin (2004), Chiswick & DebBurman (2004), Corak (2012), and Beck et al. (2012), confirms the disadvantage, in terms of educational attainment, for immigrant children that arrive at older ages. For example, using the 2000 census and a similar identification strategy to Bleakley & Chin (2004), Beck et al. (2012) estimate that for each additional year of arrival after the age of eight the probability of being a high school dropout increases by approximately two probability points per year, relative to a child immigrant from a control group (English Canada).

The principal conclusion from this research is that this educational disadvantage is consistent with neuro-biologic and linguistic research which indicates a certain 'critical period' up until which young children are able to acquire a native level of proficiency of new languages almost effortlessly (Newport, 2003). This form of the 'critical period hypothesis' implies there is no age-dependent relationship in language outcomes if the age at first exposure to the new language is before the onset of a distinct threshold. If the age of exposure to the new language occurs after this threshold there is a decay in the capacity to learn the language. Following Bleakley & Chin (2004), Böhlmark (2008), Beck et al. (2012), and van den Berg et al. (2014), this critical period occurs somewhere between the ages of eight and eleven as children make the transition from 'learning to read' to 'reading to learn'. Note that this 'critical period' might also arise in the presence of psychological and social factors, rather than biological factors, that favour younger children such as greater motivation and pressure to use the second language on a daily basis.[1]

One difficulty with identifying a 'causal' effect of age at immigration upon educational attainment is that age at immigration, language skills, and completed years of education are most likely jointly determined and correlated with parental characteristics. For example, parents with a strong preference for their children's education might tend to immigrate with

[1](McLaughlin, 1992): http://people.ucsc.edu/~mclaugh/MYTHS.htm

young children. By definition, these parents are also more likely to possess characteristics that are beneficial to their children's educational attainment. Consequently, the educational attainment of immigrant children who arrive at younger ages will tend to be higher not necessarily because they arrived at a younger age but because of the positive environment they experienced as children.

Although parental preferences for their children's education are generally unobserved, they will likely be positively correlated with observed parental characteristics, such as educational attainment. Guryan et al. (2008) present robust evidence that one measure of parental preferences, time spent with children in productive 'educational' activities, is positively correlated with both maternal employment and maternal educational attainment. This positive correlation between parental preferences for education and observed parental education would represent one mechanism driving the well-documented intergenerational correlation between parental education and the educational attainment of their children. Despite this, there is some evidence (Card et al. (2000), Aydemir et al. (2013), Luthra & Soehl (2015)) for a weaker relationship between parental and child educational attainment in immigrant families, perhaps consistent with an 'immigrant paradox' where less educated parents with strong parental preferences for education are motivated to emigrate to exploit enhanced educational opportunities for their children.

To the extent that parental educational attainment is correlated with these parental preferences for their children's education and correlated with children's age at immigration, existing studies such as Schaafsma & Sweetman (2001), Gonzalez (2003), Chiswick & DebBurman (2004), Corak (2012) and Beck et al. (2012), will most likely suffer from an omitted variable bias associated with omitting controls for parental educational attainment. In the presence of a negative correlation between parental education and their children's age at immigration, estimates of the educational attainment-age at immigration gradient, neglecting controls for parental educational attainment, will tend to be biased downwards and over-estimate the disadvantage in terms of the educational attainment for child immigrants associated with arriving at a later age.

There are several relevant studies that attempt to address this omitted variable bias using Swedish register data. Böhlmark (2008) exploits within-family variation in age at immigration to control for time-invariant family specific characteristics (parental preferences) using grade point average (GPA) at age 16 as the outcome variable.[2] In general, the coefficients on age at immigration for this 'sibling difference' specification are statistically less negative compared to a pooled cross-section specification, consistent with a negative correlation between unobserved parental characteristics and age at immigration. Interestingly, Böhlmark (2008) show that the results from the 'sibling difference' specification with controls for the first-born child can be more or less replicated in a pooled cross-section specification with controls for parental education, country of birth, and birth order. Böhlmark (2009) presents results for a specification similar to that used in this paper where final education attainment at age 30 is the outcome variable. However, differences in specific Swedish institutions affecting the inter-generational correlation in education for immigrants, including educational institutions, together with differences in the source country composition of immigrants to Sweden, limit the application of these results to child immigrants in the United States.

Moreover, in these papers parental characteristics are measured at the same time as measurement of the outcome variable. These parental characteristics are likely changing over time with duration of residence making it difficult to disentangle acculturation effects from age at immigration effects. In contrast, this paper attempts to mitigate this problem by measuring parental education as close as possible to the time of immigration, admittedly with a cost in terms of the precision of measurement.

Exploiting the large sample size in the 2000 census, this paper applies a grouped estimator, similar to that used by Borjas (1992), Borjas (1993), Card et al. (2000), Card (2005), Aydemir et al. (2009), and Aydemir et al. (2013) to quantify the bias in the estimated gradient in educational attainment by age at immigration associated with neglected controls for

[2]van den Berg et al. (2014), using a similar identification strategy, examine adult height and final educational attainment as outcome variables.

parental education. The results indicate a considerable over-estimation of the disadvantage for immigrant children that arrive at older ages, consistent with a negative correlation between parental education and age at immigration. The results indicate that the systematic variation of parental education with age at immigration explains as much as 28% of the disadvantage associated with arriving at ages 10-14 years, relative to arriving at ages 0-4 years.

After providing evidence for a considerable negative correlation between parental education and age at immigration, this paper provides further insights with respect to this negative correlation. First, the paper investigates one potential source for this negative correlation—positive parental educational selectivity. This recognizes that immigrant parents are not necessarily randomly-selected within source countries but generally positively selected on education and that this positive selection might vary inversely with the age at immigration of their children. The results indicate that approximately two-thirds of the bias associated with omitted controls for parental education might reasonably be attributed to the non-random educational selection of immigrant parents across the age at immigration distribution, reflecting their favourable unobserved characteristics relative to their peers in their source country.

Second, the results from an interacted model indicate that an additional year of parent's schooling has the greatest effect on their children's education for immigrants that arrived at ages 10-14 years, consistent with reduced opportunities for intergenerational assimilation of educational opportunities for late arriving children.. Together, with a considerable negative correlation between parental education and age at immigration, the estimates imply that immigrants who arrive at later ages might be able to overcome, or at least diminish, their educational disadvantage provided their parents are sufficiently educated. The estimates indicate that this sufficient education is not particularly high—approximately 12.5 years of father's schooling, just beyond those of a high school graduate, are enough to eliminate any educational disadvantage for late arriving immigrant children.

The remainder of the paper is organized as follows. The next section describes the data utilized in the paper and the method used to construct measures of parental education. A subsequent section explains the empirical model used to provide estimates of the age at immigration-educational attainment gradient with controls for parental educational attainment. A further section presents the empirical results after which the paper concludes with a discussion of the main results.

2 Data

The analysis utilizes data from the 2000 United States Census Integrated Public Use Microdata Sample (IPUMS) files (Ruggles et al. (2010)). One advantage of the population census, compared to more specialized surveys such as the New Immigrant Survey, is that the data contain information on documented as well as undocumented immigrants. In addition, the availability of a large native-born sample provide a control group that can be used to identify (native-born) birth cohort effects in educational attainment. The large sample sizes, particularly for immigrants that arrived in the United States as children, provide additional advantages in terms of statistical precision over more recent population surveys such as the American Community Survey (ACS).

The data contains information on each individual's country of birth, year of arrival in the United States, and educational attainment. The principal advantage of the 2000 census, relative to earlier census files, is the availability of year of arrival as a continuous variable.[3] Using the individual's current age, together with their year of arrival allows an exact age at immigration to be calculated for each foreign-born individual.

The focus in the paper is upon the completed educational attainment of child immigrants, defined as an individual not born in the United States who came to live in the United States anytime before the age of 15. Within the range of compulsory school attendance in all states,

[3]Prior to the 2000 census, year of arrival is reported in intervals. For example, the 1990 census reports year of arrival according to the following non-uniform intervals: Before 1950, 1950-1959, 1960-1964, 1965-1969, 1970-1974, 1975-1979, 1980-1981, 1992-1984, 1985-1986, and 1987-1990.

this age restriction ensures child immigrants complete at least some schooling in the United States. The sample is restricted to adults, aged between 25-49 years, currently not attending an educational institution. It is clear from an examination of Table 1 that this age restriction arises naturally from the restriction that immigrant children arrived in the United States before the age of 15.

As noted above, the potential for omitted variable bias arises from the correlation between age at immigration and parental characteristics, such as educational attainment, at the time of immigration. Unfortunately, the 2000 census does not provide direct measures of parental educational attainment for individuals not currently residing with their parents. However, earlier census files may be used to obtain information on parental educational attainment close to the time of arrival for a sample of immigrants who first arrived as children (and likely residing with their parents). For example, the 1970 census may be used to obtain parental educational attainment for immigrant children that arrived in the period 1965-1970, the 1980 census used to obtain parental educational attainment for immigrant children that arrived in the period 1971-1980, and the 1990 census used to obtain parental educational attainment for immigrant children that arrived in the period 1981-1990.[4]

Although immigrant children might be matched to their parents in either the 1970, 1980, or 1990 census, there is no way to match these immigrant children to the adults who are observed in the 2000 census. However, the 2000 census provides information on an individual's country of birth, year of immigration, and age at immigration, so it would be possible to assign to this individual a measure of the mean parental educational attainment of all individuals from the same source country that also arrived in the same period and report the presence of a child with the same gender and with the same age at immigration. There is considerable precedent for constructing a proxy variable for parental education in a

[4]The 1970 census is the earliest census in which year of arrival information is available so it is not possible to use earlier census files to identify more arrival cohorts. Ideally, the strategy would work best when there are quinquennial censuses so that parental education could be measured within five years of arrival in the United States. In the absence of these quinquennial censuses there is a tradeoff between measuring parental education as close as possible to the time of arrival and discarding relevant data.

similar way, both for the United States and other countries.[5] For the United States, Borjas (1992), Borjas (1993), Card et al. (2000), Card (2005) and Aydemir et al. (2009) have all used mean parental educational attainment as a proxy variable for individual parental educational attainment to study the intergenerational transmission of education from immigrants to their native-born (second generation) children. A contribution of this paper is to allow this mean parental educational attainment for potential parents to vary by their children's age at immigration

In order to accommodate the underlying heterogeneity in parental educational attainment across source countries, arrival cohorts, and their children's age at immigration, a researcher would want to allow for as many different cells as possible when calculating mean parental education. This has the advantage of improving the match between the potential parents from the 1970, 1980, or 1990 census and the actual parents of the immigrant children observed as adults in the 2000 census, mitigating the problem of 'potential slippage' identified by Card et al. (2000). However, calculating the mean for cells with a relatively small sample size will compromise the precision of the estimated mean parental education.[6] Despite this, a well understood result in the measurement error literature and noted by several authors, such as Borjas (1992) and Luthra & Soehl (2015), group mean parental education will provide estimates that are more robust to the presence of measurement error in individual reports of parental education. Striking a balance between these opposing considerations, this paper matches the observations in the 2000 census to 960 different types of potential parents from the 1970, 1980, and 1990 census defined over 40 different countries of birth, 5 different arrival cohorts (1965-69, 1970-74, 1975-79, 1980-84, 1985-89), 3 different age at immigration categories (0-4, 5-9, and 10-14), and the gender of the child.[7] Separate estimates are calculated for the mean educational attainment for each parent.

[5] Luthra & Soehl (2015) provide a lengthy list of these international studies for Canada, Denmark, France, Germany, Switzerland, and the United Kingdom.

[6] Aydemir & Borjas (2011) show that sampling variation in mean parental education will create a form of attenuation bias in the group-level estimator.

[7] As shown in Table 1 the age restriction imposed on the sample implies that not all age at immigration categories are observed for all of the arrival cohorts.

Completed educational attainment for the immigrants that arrived as children is measured using the 2000 census while the educational attainment of their (potential) parents is measured using either the 1970, 1980, or 1990 census, conditional upon their arrival cohort.[8]

This strategy has several advantages. First, it uses measures of parental educational attainment measured as close as possible to the time of immigration. Otherwise, both parental educational attainment and duration of residence for immigrant children might be increasing over time and it would be difficult to disentangle the two. For example, immigrant children that arrived at earlier ages might exhibit an educational advantage simply because their parents have also invested in domestic human capital during their years of residence. Second, the strategy provides a larger and more representative sample for analysis than the alternative of restricting attention to the sample where current parental education is observed—individuals aged 25 and older still residing with their parents. Finally, the strategy uses parent's own reports of their education, rather than relying on children's reports of their parent's education, minimizing the effects of measurement error.

A negative correlation between parental educational attainment and their children's age at immigration might arise when there are (birth) cohort effects in education in immigrant source countries. Immigrant parents from more recent birth cohorts are more likely to have higher education and more likely to have young children (because they themselves are younger). Therefore, a negative correlation between parental education and their children's age at immigration might arise even if immigrant parents are randomly selected within source countries across their children's age at immigration distribution. As a result, in cross-sectional data, some of the negative correlation between parental education and age at immigration will reflect these source country birth cohort effects, rather than unobserved parental characteristics.[9]

[8]For the 1970 census, the 1% metro (Form 1) sample is combined with the 1% neighborhood (Form 1) sample and the 1% state (Form 1) sample. For the 1980 and 1990 census files, the 5% weighted samples are combined with the weighted 1% samples.

[9]The availability of a repeated cross section ('quasi panel') or longitudinal data would permit the identification of (parental) age at immigration effects and birth cohort effects.

In order to investigate the importance of these source country birth cohort effects in parental education, this paper utilises the Barro & Lee (2010) data on educational attainment. This data, which provide mean years of schooling in 146 countries by 5 year age groups for every 5 years in the period 1950-2010, facilitate an examination of these birth cohort effects in education. Specifically, using the Barro & Lee (2010) data, mean educational attainment is calculated by parental age at immigration (birth cohort), gender, period of immigration, and source country and matched to these variables in the individual level data in the 1970, 1980, and 1990 census. The mean of this mean source country education is then calculated for each of the 960 groups, defined by country of birth, arrival cohort, (children's) age at immigration, and child gender, and then used as an additional explanatory variable in the empirical model.

Descriptive Statistics

Table 2 provides measures of the mean years of schooling for immigrants that arrived as children from the 2000 census, by age at immigration. These statistics replicate the basic empirical result that mean years of schooling are monotonically lower for immigrant children that arrived at later ages. Immigrant children that arrived ages 10-14 years have, on average, 21 months less schooling than immigrant children that arrived prior to age 5. Consistent with a 'critical period' of language acquisition, this result is observed for all regions of birth with the exception of the developed English speaking countries of birth (Australia, Canada, Ireland, New Zealand, South Africa, and the United Kingdom).

An examination of Table 3, which provides the age at immigration distribution by region of birth, reveals that this educational disadvantage associated with arriving at later ages is primarily driven by a differing region of birth composition within age at immigration categories. Considering male immigrants that arrived before the age of five, approximately 25% come from Europe, 20% come from Asia, 25% come from Mexico & Central America, 12% come from the Caribbean, and 10% come from English speaking countries. In contrast, for the group of male immigrant children that arrived after the age of nine, approximately

8% come from Europe, 19% come from Asia, 46% come from Mexico & Central America, 14% come from the Caribbean, and only 3% come from English speaking countries. An important factor driving the educational advantage for early arriving immigrants is the relatively larger share of child immigrants from English Speaking Countries and Europe who achieve higher mean levels of final educational attainment. Similarly, an important factor driving the educational disadvantage for late arriving immigrants is the relatively larger share of child immigrants from Mexico & Central America who achieve lower mean levels of final educational attainment. Similar results hold for female immigrant children.

It is possible to explore the relationship between parental education and the age at immigration of their children using individual level data from the 1970, 1980, and 1990 census files. The first panel of Table 4 provides the mean education of fathers by age at immigration of their eldest child for male immigrant children.[10] Mean years of schooling for father's who arrived with immigrant children is declining monotonically with the age of their eldest child. This provides evidence for a negative correlation between father's education and their children's age at immigration. While there is some variation in the strength of the relationship between father's education and the age at immigration of their eldest child across regions of birth, interestingly mean parental educational attainment is also decreasing in age at immigration for parents from developed English speaking regions of birth, a 'natural' control group.

The evidence presented Tables 2 & 4 is suggestive of two different groups of child immigrants—(1) later arrivals (aged 10-14) with lower completed education and lower parental education and (2) earlier arrivals (aged 0-4) with higher completed education and higher parental education. This dichotomy is consistent with age at immigration, children's language skills, and children's completed years of education being jointly determined and correlated with parental education, making it extremely difficult to disentangle the causal independent effects of any of these factors upon children's final educational attainment.

[10]An equivalent table for the mean education of mothers, available from the author, indicates similar results to that for fathers.

3 Empirical Model

Ideally, the desired statistical model would relate children's years of schooling to their age at immigration, controlling for parental characteristics. In the absence of controls for these typically unobserved parental characteristics, parental education will be used as a 'proxy' variable for the vector of these unmeasured parental characteristics.[11] As noted by Luthra & Soehl (2015), it is likely that parental education is only an imperfect proxy for unmeasured parental characteristics in the sense that not all of the correlation between age at immigration and parental characteristics can be attributed to parental education. This will generally lead to a biased estimate of the age at immigration-educational attainment profile. However, it is likely that controls for parental education lower the correlation between age at immigration and parental characteristics such that this bias will be smaller than the bias associated with (completely) omitted controls for parental characteristics.

Given the structure of the data, it is feasible to estimate an econometric model that relates individual level variation in years of schooling to variation in group mean parental educational attainment:

$$\mathrm{YR}\widetilde{\mathrm{SCHOO}}\mathrm{L}_{ij} = \beta_0 + \beta_1\,\mathrm{AAI}_{ij} + \beta_2\,\mathrm{P}\widetilde{\mathrm{AR\ E}}\mathrm{D}_j + \delta\,\mathbf{X}_{ij} + \varepsilon_{ij} \tag{1}$$

where $\mathrm{YR}\widetilde{\mathrm{SCHOO}}\mathrm{OL}_{ij}$ denotes the 'age-adjusted' years of schooling for immigrant i in group j that arrived as a child, and observed as an adult.[12] The variable AAI_{ij} denotes the broad age at immigration of individual i in group j, $\mathrm{P}\widetilde{\mathrm{AR\ E}}\mathrm{D}_j$ denotes the mean educational attainment of the (potential) parents of immigrants in group j, $\mathbf{X}$ represents a vector of control variables that vary across individuals, and the error term ε_{ij} captures unobserved

[11] Wooldridge (2001) provides a general discussion on the use of proxy variables while Luthra & Soehl (2015) provides a specific discussion with reference to parental education as a proxy variable for unmeasured parental characteristics.

[12] A discussion of the methodology for constructing this 'age-adjusted' years of schooling is provided later in this section.

determinants of educational attainment for individual i in group j.[13] The broad age at immigration categories are 0-4 years, 5-9 years, and 10-14 years and have been selected to identify key stages in child development and educational milestones.

The actual estimates on the parental education variables are of limited interest in the reduced form model and simply represent the reduced form contribution of parent's education to their children's education but cannot be interpreted as causal parameters for several reasons. First, mother's and father's education is highly correlated such that the joint effect of parental educational attainment upon their children's completed schooling is well identified but the individual effects of each parent's education might not be well identified. Consequently, this paper focuses upon the joint effect of parental educational attainment. Second, mean parental education in the model (1) can affect child educational attainment both directly through their own parent's education as well as through the mean group characteristics of the group defined by country of birth, arrival cohort, age at immigration, and gender. Despite this, it is anticipated that the group estimator utilized in this paper, constructed by defining groups based upon multiple characteristics, implies a weaker role for 'social capital' in determining the intergenerational elasticities in education. Finally, as shown by Wooldridge (2001), even if parental education is a 'perfect proxy' for unmeasured parental characteristics, the estimates of parental education will still be biased.

Proceeding with due caution, controlling for age at immigration, higher levels of parental educational attainment will be associated with higher levels of educational attainment at for their children ($0 < \beta_2 < 1$). If $\beta_2 = 1$ in model (1) there is no intergenerational assimilation since parental educational outcomes are fully reflected in their children's educational attainment. On the other hand, if $\beta_2 = 0$ there is complete intergenerational assimilation since parental education has no effect upon their children's educational attainment.

Conditioning on parental education, estimates of β_1 in the econometric model (1) will tend to be negative— on average, child immigrants that arrive at later ages will tend to have

[13] The group mean $\widetilde{\mathrm{PAR\ ED}}_j$ is calculated using the available IPUMS person level sample weights.

lower levels of educational attainment. This is particularly more likely since, as a result of data limitations, the model (1) does not contain direct controls for language skills or birth order.[14] Since membership in each group is defined by a country of birth, broad arrival cohort, broad age at immigration, and gender, it might be argued that the group mean for parental education also very crudely controls for language skills at arrival.

For the underlying individual level model there are three types of age that will be important—current age of the immigrant, age at the time of immigration, and years since immigration. These three age variables are perfectly collinear such that given knowledge of any two, the effect of the third can be trivially calculated.[15] This implies that the separate effects of current age, age at immigration, and arrival cohort are not identified for the individual level version of model (1). One (common) solution to this collinearity problem is to impose the identifying restriction that the (current) age effects are identical for immigrants and the native born. In this case, the current age effects for the adult immigrant sample might be identified from the age effects for the native born such that age at immigration and arrival cohort can be identified. An examination of Table 1 confirms that, for a given current age, both age at immigration and arrival cohort are (in principle) identified.

One strategy is to pool the samples for both the native born and the immigrants and estimate a single age profile in educational attainment. As noted by Schaafsma & Sweetman

[14] There are several compelling reasons to exclude controls for language skills. First, there are two sources of variation in language skills that are important—language skills at arrival and current language skills. The census data only contain a crude measure of the latter with no information on the former. More importantly, to the extent that age at immigration, language skills, and completed years of education are jointly determined and correlated with parental characteristics, in model (1) language skills would represent an endogenous variable. The identification of the causal effect of age at immigration on adult educational attainment would require an appropriate instrument for language skills. A popular instrument that exploits the variation implied by the critical period hypothesis of language acquisition used by Bleakley & Chin (2004) is based upon the interaction of age at immigration with country of birth. However, this is likely to be an invalid instrument that fails the exclusion restriction when age at immigration is correlated with (omitted) parental education. In the absence of an appropriate instrument, this paper focuses upon the correlation between parental education and age at immigration and its implications for estimation of the educational attainment-age at immigration gradient. As a further check on the baseline results, Table A.5 presents estimates from a version of model (1) that includes controls for language skills.

[15] For example holding current age constant, years since immigration and age at immigration represent the same variation. Alternatively, holding years since immigration constant, current age and age at immigration represent the same variation. Finally, holding age at immigration constant, current age and years since immigration represent the same variation.

(2001), in the presence of non-linearities in the age profile, the estimates of the age at immigration profile would then represent deviations from this common age profile, a mixture of the native born and immigrant age profiles. Since the focus in this paper is upon the age at immigration profile, it is not desirable to have these estimates contaminated by the native-born age profile. In addition, under this strategy, the native-born become the control group. In the context of child immigration it seems more appropriate to have child immigrants that arrived say at age zero as the control group.

Instead, this paper adapts the two-step approach described in Schaafsma & Sweetman (2001). The first stage involves regressing individual years of schooling on a full set of dummy variables indicating membership in each age group (birth cohort) on a sample of native born individuals aged 25-49 years. As shown in Table 1, this age restriction corresponds to the age restriction in the sample of child immigrants who arrived between 1965 and 1990 who are currently over the age of 25 and arrived in the United States before the age of 15. This model is estimated separately for male and females to allow for the age (birth-cohort) profile to vary by gender. The estimated age coefficients represent the mean years of schooling for a native born individual, relative to a native born reference group, in this case a 25 year old individual.

In the second stage, the difference between each immigrant's years of schooling and their predicted years of schooling (if this immigrant were the same age and gender but born in the United States instead of abroad) is constructed. The predictions are constructed using the estimated coefficients from the first stage statistical model. It is this predicted schooling that is used as the dependent variable ($\mathrm{YR}\widetilde{\mathrm{SCHOOL}}_{ij}$) in model (1).

There are two main issues associated with statistical inference on the estimated parameters in model (1). First, there is a 'generated regressor' associated with using estimated group-level parental education. A failure to account for this sampling variation will lead to misleadingly small standard errors. Second, given this group structure for parental education, it seems reasonable to assume that the model errors are uncorrelated across clusters but correlated within clusters. It is well understood that failing to account for this 'clustering' problem can lead to misleadingly small standard errors, narrow confidence intervals, and low p-values. Cameron & Miller (2015) provide an excellent review of the current methods for conducting cluster-robust statistical inference.

In light of both of these issues, the standard errors for model (1) are estimated using bootstrap (re-sampling) methods. Although the data have a large number of clusters, there is considerable variation in the (sample) sizes as each cluster. As suggested by Cameron & Miller (2015), bootstrapping methods would be advised in this case. Tables A.1 and A.2 in the appendix compare various methods for estimating the standard errors for model (1). Comparing the bottom three panels with the first two panels of Table A.1 indicate that the standard errors which ignore the clustering dramatically understate the 'correct' standard errors by at least a factor of five. While the p-values for the Wald test using the Wild bootstrap are, as expected, larger than those obtained under the paired bootstrap, there is no evidence of considerably different p-values compared to the paired bootstrap. Based upon these results, and taking into account the additional computational complexity involved with the Wild bootstrap, the standard errors for the model (1) will be estimated using the paired bootstrap.

The parameter of interest is β_1, the difference in the mean schooling, conditional on $\mathbf{X}$, between immigrant children that are members of age at immigration cohort AAM_k, relative to those that are members of some omitted age at immigration cohort (0-4 years). Table 1 verifies that both age at immigration effects and arrival cohort effect can, in principle, be identified within the single cross section used in this paper. In order to quantify the extent of the omitted variable bias, the estimate of β_1 in model (1) is compared with a restricted version of model (1) that excludes controls for parental education ($\beta_2 = 0$).

While it is relatively straightforward to calculate the difference in the estimated coefficient $\hat{\beta}_1$ between the unrestricted and restricted model ($\beta_2 = 0$), some care must be exercised when computing a test of the significance of this difference in the estimated coefficients.

Since the samples are dependent, a Wald test for $\hat{\beta}_1$ assuming independent samples would be inappropriate since it ignores the covariance between the estimators for different equations.[16]

Clogg et al. (1995) provides an analytical solution to this problem. However, this cannot be implemented since the standard errors for model (1) are obtained using bootstrapping techniques (paired bootstrap). Consequently, the 'correct' confidence intervals for the difference in the estimated coefficient $\hat{\beta}_1$ between the restricted ($\beta_2 = 0$) and unrestricted model are also calculated using bootstrapping techniques.[17]

As discussed above, one advantage of the grouping estimator is that it measures parental educational attainment as close as possible to the time of immigration. However, given the data structure, there could be at most a difference of 9 years between the year of immigration and the year of observation. For example, educational attainment for parents who arrived in 1981 is not measured until the census in 1990. In the presence of parental 'upskiing' between the period of immigration and the period of observation, both parental educational attainment and duration of residence for immigrant children might be increasing over time and it would be difficult to disentangle the two. Immigrant children that arrived at earlier ages might exhibit an educational advantage simply because their parents have also invested in domestic human capital during their years of residence. This would imply that, a comparison of the estimate of β_1 in model (1) with a restricted version of model (1) that excludes controls for parental education ($\beta_2 = 0$), would overstate the extent of bias due to omitted controls for parental education on arrival since some of this bias would reflect the accumulation of

[16] Denote the estimate of β_1 in the restricted and unrestricted models as $\hat{\beta}_1^R$ and $\hat{\beta}_1^{UR}$ respectively. The standard error of the difference is calculated as:

$$\text{se}(\hat{\beta}_1^{UR} - \hat{\beta}_1^{RR}) = \sqrt{\text{VAR}(\hat{\beta}_1^{UR}) + \text{VAR}(\hat{\beta}_1^{R}) - 2\ \text{COV}(\hat{\beta}_1^{UR}, \hat{\beta}_1^{R})}$$

In the absence of omitted variable bias, $\hat{\beta}_1^R$ and $\hat{\beta}_1^{UR}$ would be perfectly positively correlated. Therefore, in the presence of bias associated with omitted controls for parental education, $\hat{\beta}_1^R$ in the restricted model would be positively (but not perfectly) correlated with $\hat{\beta}_1^{UR}$ in the unrestricted model. The estimated covariance of the point estimates might not need to be particularly large to sufficiently reduce the standard error. Ignoring the covariance term might result in sufficiently large standard errors and wide confidence intervals leading to non-rejection of the null hypothesis $\hat{\beta}_1^R = \hat{\beta}_1^{UR}$ while accounting for this covariance might lead to a smaller standard error and narrower confidence intervals leading to rejection of the null hypothesis.

[17] As a check, these 'bootstrapped' confidence intervals are compared to those obtained using the technique described in Clogg et al. (1995). Very little discernible difference can be detected.

domestic human capital after arrival.

In order to investigate the existence of parental ‘upskilling’, an alternative model to (1) is estimated that uses a counter-factual measure of parental education ($\widetilde{\text{POPSCHOOL}}$) which itself is estimated using the following model with individual level data from the 1970, 1980, and 1990 census:

$$\text{POPSCHOOL}_i = \beta_0 + \sum_{j=2}^{40} \beta_{1j}\,\text{COB}_{ij} + \beta_{80}\,\text{C80}_i + \beta_{90}\,\text{C90}_i + \beta_5\,\text{D5}_i + \sum_{j=2}^{40} \beta_{5j}\,\{\text{COB}_{ij} * \text{D5}_{ij}\}$$
$$+ \sum_{j=2}^{40} \beta_{80j}\,\{\text{COB}_{ij} * \text{C80}_{ij}\} + \sum_{j=2}^{40} \beta_{90j}\,\{\text{COB}_{ij} * \text{C90}_{ij}\} + \varepsilon_i \tag{2}$$

where POPSCHOOL_i represents individual level father’s schooling, COB_{ij} is an indicator variable denoting the country of birth of individual i, C80_i is an indicator variable denoting the ten year arrival cohort 1971-1980 while C90_i is an indicator variable denoting the ten year arrival cohort 1981-1990, and D5_i is an indicator variable that takes on a value of zero if parental education is observed within 0-4 years of arrival and a value of one if parental education is observed within 5-9 years of arrival. The predictive model allows for interactions of country of birth with D5 as well as interactions of country of birth with ten-year arrival cohort.

With an identifying restriction that year of arrival cohort effects are constant over a ten year period, counter-factual predicted parental education ($\widetilde{\text{POPSCHOOL}}$) is constructed for parents whose education is observed at least five years after arrival ($\text{D5} = 1$). In this case, the parent is assumed to have arrived within their same ten year arrival cohort and same source country but their education was measured five years earlier, closer to their year of immigration. The estimated model would provide suggestive evidence for the presence of ‘upskilling’ when actual group mean parental education exceeds this counter-factual group mean education. Counter-factual mother’s schooling is constructed in a similar way.

4 Results

Table 5 presents estimates from econometric model (1), separately by the gender of the child immigrant, where the vector of controls **X** includes indicator variables for five year arrival cohort (1965-69, 1970-74, 1975-79, 1980-84, and 1985-89) and region of birth.

First, consider the estimates that do not control for parental education given by specification (1) in Table 5. The estimates imply that average educational attainment of male immigrants that arrived at ages 10-14 years compared to those that arrived at ages 0-4 years is approximately one year (12 months) lower for males and almost 1.09 years (13 months) lower for females. These results reproduce the stylized result that immigrant children arriving at a later age have lower levels of adult educational attainment, consistent with a 'critical period' of second language acquisition somewhere between the ages of eight to eleven years of age.

Specification (2) in Table 5 introduces controls for mean parental educational attainment for each parent and should be compared to specification (1) without parental controls. As noted above, the actual estimates on the parental education variables are of limited interest in this reduced form model and simply represent the reduced form contribution of parent's education to their children's education but cannot be interpreted as causal parameters for several reasons.

An examination of the results presented in Table 5 indicate that a considerable portion of the educational disadvantage from arriving at a later age might reasonably be attributed to variation in parental educational attainment. For example, comparing columns (1) and (2) in Tables 5, relative to immigrants that arrived aged 0-4 years, the mean educational disadvantage for male immigrants arriving at ages 10-14 years decreases from 1.005 years (12.06 months) of schooling to only 0.724 years (8.69 months). This is a statistically significant difference of 0.280 years (3.36 months). Alternatively, relative to immigrants that arrived aged 0-4 years, approximately 28% of the educational disadvantage for male immigrants arriving at ages 10-14, can be attributed to the lower parental education of these later arrivals, rather than their later age at immigration. Similarly, controlling for parental education improves mean relative educational attainment of female immigrants that arrived at aged 10-14 years by 0.278 years (3.33 months).

Although, the results presented in Table 5 indicate that neglected controls for parental education produce a considerable bias in estimates of the age at immigration-education gradient, caution must be exercised when interpreting these results. Group mean (total) parental education can be represented as the sum of two terms: (1) mean educational attainment by birth cohort, sex, and period of immigration in their source country and (2) the deviation of group mean parental education from this mean source country education.

Mean source schooling would be an appropriate measure of parental schooling when immigrant parents are randomly selected (relative to non-immigrants) within source countries. In the presence of birth cohort effects in educational attainment, it is feasible that, within source countries, mean source schooling might be negatively correlated with age at immigration. However, most of the important variation in mean source country schooling is across source countries rather than across birth cohorts within source countries. Indeed results (available from the author) from a specification that uses mean source country schooling as the measure of parental education reveal that even with crude source region controls, the marginal effect of mean source country schooling upon children's educational attainment is extremely small and statistically insignificant, suggesting that source country schooling mainly proxies for source country characteristics. Consequently, the result in column (2) in Table 5 likely over-state the extent of the bias associated with neglected controls for parental education since (total) parental education includes a component which mainly reflects source country characteristics. A portion of the omitted variable bias also reflects omitted controls for source country characteristics.

In light of this, the results in columns (3) and (4) in Table 5 include additional controls for mean source country schooling defined as the group average of mean source country education for each of the 960 groups, defined by country of birth, arrival cohort, (children's) age at immigration, and child gender, as an additional explanatory variable.[18] This represents a parsimonious way of capturing omitted (from column 2) source country characteristics. However, the interpretation of the estimated coefficients is now different. Controlling for mean source schooling, higher values of total parental education now represent years of education beyond the mean schooling of an equivalent non-immigrant from the same source country. This essentially represents parental educational selection relative to their non-immigrant peers. The coefficient on parent's schooling represents the marginal effect of an additional year of education beyond the mean schooling of an equivalent non-immigrant from the same source upon children's educational attainment. The similarity of the coefficients on parental education in columns (2) and (3) in Table 5 suggests that the 'intergenerational coefficient' on parental education is mainly determined by the extent of parental education selectivity.

It is worth noting that the estimated coefficient on source country education in Table 5 is negative. Compare two children who arrived at the same age with the same (total) parental education. The estimated coefficient for source country education then gives the marginal difference in children's schooling if one of these individual has higher source country education. However, controlling for total parental education, a higher source country mean education implies less parental educational selectivity.

Comparing columns (2) and (3) in the top panel Table 5 (which both control for parental education), additional controls for source country mean education for males *increase* the estimated mean educational disadvantage for later arriving children by almost a whole month from 0.724 years (8.69 months) of schooling to 0.812 years (9.74 months), a statistically significant difference of 0.088 years (1.06 months). While (total) parental education accounts for approximately 28% of the education disadvantage for later arriving immigrants (column 2), non-random parental education selection accounts for 18% of this education disadvantage (column 3). This suggests that approximately one third of the bias associated with

[18]This mean source country schooling is constructed using the Barro & Lee (2010) data.

omitted controls for parental education (0.280 years) might be reasonably be attributed to omitted source country characteristics, 'proxied' by mean source country schooling. After controlling for source country birth cohort effects in education, differences in parental education between individuals now reflect differences in the degree of parental education selection, relative to non-immigrants in the source country. Consequently, the remaining two-thirds mainly reflects the non-random educational selection of immigrant parents across the age at immigration distribution, reflecting their favourable unobserved characteristics relative to their peers in their source country. Quantitatively similar results hold when using controls for source country mean education for females (column 4) and for female immigrant children (bottom panel).

The results in Table 5 are consistent with parental education as an omitted variable that is (negatively) correlated with age at immigration and (positively) correlated with children's educational attainment. This suggests that the educational disadvantage from arriving at a later age might, to some extent, be reduced provided parental education is sufficiently high. This is investigated using a version of model (1) where father's schooling is interacted with age at immigration.[19] This interacted model provides for the return to an additional year of father's schooling to vary by their children's age at immigration. For the specification that includes controls for mean source country schooling, these estimated returns for an additional year of father' schooling are 0.084, 0.188, and 0.396 for children that arrived at aged 0-4 years, 5-9 years, and 10-14 years respectively. Importantly, these results indicate that an additional year of father's schooling has the greatest effect on their children's schooling for immigrants that arrive aged 10-14 years. This provides suggestive evidence that children's age at immigration might also be an important determinant of the intergenerational transmission of group level educational attainment for first generation immigrant families. Specifically, this result is consistent with reduced opportunities for intergenerational assim-

[19] An examination of Table A.3 and Figures A.1 and A.2 reveals quantitatively similar results using mother's schooling.

ilation of educational opportunities for late arriving immigrant children.

Figure 1 presents the estimated marginal effects for the interacted version of model (1) that includes controls for mean source country schooling for immigrants that arrived aged 10-14 years, relative to those that arrived aged 0-4 years. An examination of Figure 1 indicates that the educational disadvantage associated with arriving at ages 10-14 years can be more or less eliminated for immigrants whose father's have an education above the 72nd percentile (12.305 years) and 73rd (12.958 years) for male and female immigrant children respectively. Interestingly, male immigrants arriving at ages 10-14 years whose father's have an education above the 88th percentile (14.288 years) exhibit an educational advantage. For female immigrants, this educational advantage arises for levels of father's schooling above the 93rd percentile (15.055 years). Specifically, the presence of reduced opportunities for intergenerational assimilation for late arriving immigrant children implies that the late arriving children of parents with sufficiently high levels of education might experience an educational advantage, relative to children arriving at earlier ages.

Checks for Robustness

In the presence of parental 'upskiling' between the period of immigration and the period of observation, both parental educational attainment and duration of residence for immigrant children might be increasing over time and it would be difficult to disentangle the two. A comparison of the estimate of the age at immigration coefficients in model (1) in a restricted version of model (1) that excludes controls for parental education ($\beta_2 = 0$), would overstate the extent of bias due to omitted controls for parental education on arrival since some of this bias would reflect the accumulation of domestic human capital after arrival.

In order to quantify the likely impact of 'upskilling' on the estimates provided in Table 5, Table 6 presents the estimates from model (1) that uses the predicted counter-factual group mean parental education, discussed in Section 3 above.[20] A comparison of the estimates in Tables 5 and 6 indicate that a small portion of the bias associated with omitted controls for parental education and provided in Table 5 might reasonably be attributed to parental 'upskilling'. However, perhaps unsurprisingly, in a model using group mean parental education, this 'upskilling' is unlikely to have a statistically significant impact.

It is possible to measure the (actual) educational attainment of parents for the sample of individuals that arrived as immigrant children, now observed as adults, who also still reside with their parents. This will represent a small, non-random sample of all immigrants who arrived as children. Additionally, the outcome of matching these children to their actual parents would only provide information on current parental educational attainment and not parental educational attainment at the time of immigration. Nonetheless, this sample provides a natural 'control' group that can be used to measure the accuracy of the approximation implied by the proxy variable for parental education, in particular the precision of the match between actual parents and potential parents.

Three alternative measures of parental education are used to estimate model (1) using the sample of immigrants that arrived as children who still reside with their parents. One measure uses the actual parental education for this sample. Based upon the groups (country of birth, arrival cohort, age at immigration, and immigrant child gender) that can be identified in the 2000 census for this sample, two measures of group mean parental education are also calculated. The first is the group mean parental education that has been used up to this point, based upon the information in the 1970, 1980, and 1990 census. The second is the group mean parental education based upon the information in the 2000 census. Mainly as a result of the much smaller number of groups and the resulting loss in precision, an examination of the estimates reported in Table 7 indicates there is no quantitative difference in

[20]Results, available from the author, indicate that there is some variation in the apparent incidence of upskilling by region—English speaking and European immigrant parents generally have lower rates of apparent 'upskilling' while immigrant parents from Asian, Central American, and the residual region (Oceania, Africa, and the Middle East) tend to exhibit greater rates of apparent 'upskilling'. Intuitively, immigrant parents from these regions most likely face greater incentives to invest in domestic human capital primarily because they encounter the greatest difficulties in transferability of their foreign human capital.

the results based upon the group mean calculated for actual parental education (column 1) as opposed to group mean potential parental education (columns 2 and 3). At least for this special sample of immigrant children still residing with their parents, there is no evidence that the availability of actual parental education would lead to a vastly different conclusion to those reported in Table 5 using a group mean approximation for parental education.

The estimates of model (1), reported in Table 5, restrict each unit of parent's education to be equally productive for their children's education, regardless of the source region of the parents. Luthra & Soehl (2015) provide convincing evidence against this restriction that intergenerational educational transmission is uniform between immigrants of differing origins. As a crude way of relaxing this restriction, the estimates presented in Table A.4 allow the return to parental education to vary across broad immigrant source regions. These results indicate that those regions which face the greatest educational disadvantage for late arrivals (Mexico & Central America) also face the greatest return to an additional year of parental education. This is consistent with a lower 'intergenerational assimilation' rate for the most disadvantaged immigrant children. Comparing the results in Table 5 with those in Table A.4 provides suggestive evidence that restricting returns to parent's education to be the same across source regions might slightly yet not significantly understate the bias associated with neglected controls for parental education in estimates of the age at immigration-education gradient.

Language skills, completed education, and age at immigration are likely jointly determined and correlated with parental education. Consequently, the estimated model (1) excludes controls for English language ability. While language skills will represents an endogenous variable in model (1) threatening any causal interpretation of the estimates, their inclusion as additional controls provides some insights into the potential mediating effect of controls for language skills for the 'baseline' estimates provided in Table 5.[21] An examination of the estimates provided in Table A.5 indicate that controls for language ability does very little to alter the parameters of interest.[22] Relative to immigrants that arrived aged 0-4 years, approximately 35% of the educational disadvantage for male immigrants arriving at ages 10-14, can be attributed to the lower parental education of these later arrivals (compared to 28% for the model that excludes controls for language ability). Additionally, the estimates indicate that approximately 42% of the bias associated with omitted controls for parental education (0.280 years) might be reasonably be attributed to omitted source country characteristics, 'proxied' by mean source country schooling (compared to 31% for the model that excludes controls for language skills). While English deficiency is associated with approximately 2.5 years less education, there is no evidence that the relatively large main effects identified in this paper can be attributed to omitted controls for English deficiency.

5 Conclusion

A substantial empirical literature confirms the disadvantage, in terms of educational attainment, for immigrant children that arrive at older ages. This paper has provided robust evidence for a considerable negative correlation between parental education and the age of immigration of their immigrant children. In the presence of this negative correlation between parental education and their children's age at immigration, estimates of the educational attainment-age at immigration gradient, neglecting controls for parental educational attainment, are biased downwards and over-estimate the disadvantage in terms of the educational attainment for child immigrants associated with arriving at a later age. The results

[21] There are several reasons why language skills would be endogenous. First, it is likely that unobserved variables that are correlated with both educational attainment and language skills generate an omitted variable bias. Second, since lower completed education decreases expected wages and lowers the incentives to invest in human capital, we might expect causality from education to language skills. Third, the prevalence of substantial measurement error in self-reported measures of language ability creates an attenuation bias.

[22] English proficiency is characterised as the self-reported ability to speak English. From this variable, we construct a binary indicator of English deficiency that takes on the value of 0 if the respondent reports speaking English very well (the highest category) and a value of 1 otherwise.

indicate that the systematic variation of parental education with age at immigration explains as much as 28% of the disadvantage associated with arriving at ages 10-14 years, relative to arriving at ages 0-4 years.

This paper provides several insights with respect to this negative correlation. First, the paper investigates one potential source for this negative correlation—positive parental educational selectivity. This recognizes that immigrant parents are not necessarily randomly-selected within source countries but generally positively selected on education and that this positive selection might vary inversely with the age at immigration of their children. The results indicate that approximately two-thirds of the bias associated with omitted controls for parental education might reasonably be attributed to the non-random educational selection of immigrant parents across the age at immigration distribution, reflecting their favourable unobserved characteristics relative to their peers in their source country.

Second, the results from an interacted model indicate that an additional year of parent's schooling has the greatest effect on their children's education for immigrants that arrived at ages 10-14 years, consistent with reduced opportunities for intergenerational assimilation of educational opportunities for late arriving children. In conjunction with a considerable negative correlation between parental education and age at immigration, the estimates imply that immigrants who arrive at later ages might be able to overcome, or at least diminish, their educational disadvantage provided their parents are sufficiently educated. The estimates indicate that this sufficient education need not be particularly high—approximately 12.5 years of father's schooling, just beyond those of a high school graduate, would be enough to eliminate any educational disadvantage for late arriving immigrant children. Taking these results a step further suggests that one advantage of an immigrant selection policy that places greater emphasis on (documented) applicant's education lies in its ability to improve the educational outcomes of the applicant's foreign born children. This is particularly relevant since for many immigrants the motivation for migrating to a new country has more to do with providing opportunities for children than advancing their own economic wellbeing.

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Table 1: Age at Immigration: 1965-69, 1971-1980 & 1981-1990 Arrivals

		1965-1970 Year of Arrival									
Age in 1970	Age in 2000	1961	1962	1963	1964	1965	1966	1967	1968	1969	1970
0	30										0
1	31									0	1
2	32								0	1	2
3	33							0	1	2	3
4	34						0	1	2	3	4
5	35					0	1	2	3	4	5
6	36					1	2	3	4	5	6
7	37					2	3	4	5	6	7
8	38					3	4	5	6	7	8
9	39					4	5	6	7	8	9
10	40					5	6	7	8	9	10
11	41					6	7	8	9	10	11
12	42					7	8	9	10	11	12
13	43					8	9	10	11	12	13
14	44					9	10	11	12	13	14
15	45					10	11	12	13		
16	46					11	12	13			
17	47					12	13	14			
18	48					13	14				
19	49					14					
		1971-1980 Year of Arrival									
Age in 1980	Age in 2000	1971	1972	1973	1974	1975	1976	1977	1978	1979	1980
5	25					0	1	2	3	4	5
6	26				0	1	2	3	4	5	6
7	27			0	1	2	3	4	5	6	7
8	28		0	1	2	3	4	5	6	7	8
9	29	0	1	2	3	4	5	6	7	8	9
10	30	1	2	3	4	5	6	7	8	9	10
11	31	2	3	4	5	6	7	8	9	10	11
12	32	3	4	5	6	7	8	9	10	11	12
13	33	4	5	6	7	8	9	10	11	12	13
14	34	5	6	7	8	9	10	11	12	13	14
15	35	6	7	8	9	10	11	12	13	14	
16	36	7	8	9	10	11	12	13	14		
17	37	8	9	10	11	12	13	14			
18	38	9	10	11	12	13	14				
19	39	10	11	12	13	14					
20	40	11	12	13	14						
21	41	12	13	14							
22	42	13	14								
23	43	14									
		1981-1990 Year of Arrival									
Age in 1990	Age in 2000	1981	1982	1983	1984	1985	1986	1987	1988	1989	1990
15	25	6	7	8	9	10	11	12	13	14	
16	26	7	8	9	10	11	12	13	14		
17	27	8	9	10	11	12	13	14			
18	28	9	10	11	12	13	14				
19	29	10	11	12	13	14					
20	30	11	12	13	14						
21	31	12	13	14							
22	32	13	14								
23	33	14									

Table 2: Mean Years of Schooling: Immigrant Children

	Age at Immigration					
	0-4 years		5-9 years		10-14 years	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Immigrant Children	MALES					
All Regions	13.276	(2.879)	12.727	(3.264)	11.515	(4.032)
English Speaking[a]	14.107	(2.473)	14.190	(2.469)	14.332	(2.458)
Europe	13.633	(2.390)	13.445	(2.547)	13.163	(2.881)
Asia	14.531	(2.512)	14.411	(2.648)	14.080	(2.928)
Mexico & Central America	11.583	(3.067)	10.834	(3.418)	9.350	(3.960)
Caribbean	12.751	(2.731)	12.690	(2.759)	12.340	(2.993)
South America	13.848	(2.628)	13.538	(2.582)	13.264	(2.750)
Oceania, Africa, & M East	14.108	(2.558)	14.033	(2.622)	14.003	(3.006)
Number of Observations	19,770		21,241		35,140	
Immigrant Children	FEMALES					
All Regions	13.556	(2.742)	12.987	(3.126)	11.930	(3.829)
English Speaking[a]	14.356	(2.408)	14.304	(2.308)	14.256	(2.415)
Europe	13.819	(2.344)	13.505	(2.512)	13.050	(2.865)
Asia	14.709	(2.379)	14.548	(2.623)	14.000	(2.958)
Mexico & Central America	11.940	(2.828)	11.150	(3.282)	9.707	(3.925)
Caribbean	13.014	(2.758)	13.014	(2.725)	12.73	(2.893)
South America	14.320	(2.558)	13.738	(2.386)	13.379	(2.661)
Oceania, Africa, & M East	14.291	(2.388)	14.004	(2.673)	13.829	(2.858)
Number of Observations	20,916		21,547		32,435	

[a] Australia, Canada, Republic of Ireland, New Zealand, South Africa, and the United Kingdom.

Notes: Sample means are weighted utilising the available population weights (PERWT) with standard deviations shown in parentheses. Children's mean years of schooling is calculated using the 2000 census files with total sample sizes of 76,151 and 74,898 for males and females respectively.

Table 3: Age at Immigration Distribution by Region of Birth

	Age at Immigration					
	0-4 years		5-9 years		10-14 years	
Region of Birth	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
	MALES					
English Speaking[a]	0.097	(0.295)	0.064	(0.245)	0.033	(0.178)
Europe	0.246	(0.431)	0.120	(0.325)	0.076	(0.265)
Asia	0.201	(0.401)	0.212	(0.409)	0.190	(0.392)
Mexico & Central America	0.248	(0.432)	0.335	(0.472)	0.469	(0.499)
Caribbean	0.121	(0.326)	0.170	(0.375)	0.141	(0.348)
South America	0.039	(0.194)	0.050	(0.217)	0.046	(0.209)
Oceania, Africa, & M East	0.048	(0.215)	0.049	(0.217)	0.045	(0.208)
	FEMALES					
English Speaking[a]	0.093	(0.29)	0.065	(0.246)	0.039	(0.194)
Europe	0.246	(0.431)	0.123	(0.328)	0.082	(0.275)
Asia	0.213	(0.410)	0.213	(0.410)	0.201	(0.400)
Mexico & Central America	0.232	(0.422)	0.312	(0.463)	0.405	(0.491)
Caribbean	0.133	(0.339)	0.190	(0.392)	0.177	(0.382)
South America	0.037	(0.188)	0.050	(0.217)	0.051	(0.220)
Oceania, Africa, & M East	0.047	(0.211)	0.048	(0.213)	0.045	(0.207)

[a] Australia, Canada, Republic of Ireland, New Zealand, South Africa, and the United Kingdom.

Notes: Sample means are weighted utilising the available population weights (PERWT) with standard deviations shown in parentheses. Children's mean years of schooling is calculated using the 2000 census files with total sample sizes of 76,151 and 74,898 for males and females respectively.

Table 4: Mean Years of Father's Schooling

	Eldest Child Age at Immigration					
	0-4 years		5-9 years		10-14 years	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Father's Schooling[b]	MALES					
English Speaking[a]	14.531	(3.239)	14.086	(3.315)	13.637	(3.116)
Europe	12.831	(4.218)	11.841	(4.342)	10.522	(4.487)
Asia	14.011	(4.166)	13.343	(4.293)	12.674	(4.531)
Mexico & Central America	8.677	(3.815)	8.216	(3.940)	7.562	(3.762)
Caribbean	11.610	(3.354)	11.105	(3.091)	10.561	(3.279)
South America	13.036	(3.546)	12.304	(3.523)	11.832	(3.571)
Oceania, Africa, & M. East	14.502	(3.587)	13.513	(4.019)	12.745	(4.223)
Father's Schooling[b]	FEMALES					
English Speaking	14.409	(3.595)	13.98	(3.255)	13.848	(3.023)
Europe	12.720	(4.167)	11.807	(4.463)	10.403	(4.422)
Asia	13.850	(4.295)	13.478	(4.290)	12.764	(4.402)
Mexico & Central America	8.721	(3.837)	8.189	(4.068)	7.694	(3.843)
Caribbean	11.264	(3.266)	10.807	(3.135)	10.544	(3.264)
South America	12.387	(3.369)	12.503	(3.523)	11.850	(3.509)
Oceania, Africa, & M. East	14.546	(3.593)	13.532	(3.939)	12.880	(3.982)

[a] Australia, Canada, Republic of Ireland, New Zealand, South Africa, and the United Kingdom.

[b] Source: 1970, 1980, and 1990 individual level census files

Notes: Sample means are weighted utilising the available population weights (PERWT) with standard deviations shown in parentheses. Mean father's years of schooling is calculated using the 1970, 1980, and 1990 census files with total sample sizes of 78,392 and 72,838 for males and female immigrant children respectively.

Table 5: Linear Regression: Years of Schooling for Child Immigrants

	(1)		(2)		(3)		(4)	
	Coeff.	Std. Err	Coeff.	Std. Err [p val.]	Coeff.	Std. Err [p val.]	Coeff.	Std. Err [p val.]
MALE CHILDREN								
Father's Schooling	—	—	0.233^a	(0.069)	0.233^a	(0.062)	0.217^a	(0.063)
Mother's Schooling	—	—	−0.050	(0.081)	−0.007	(0.073)	0.009	(0.074)
Mean Source Schooling: Males	—	—	—	—	−0.128^a	(0.033)	—	—
Mean Source Schooling: Females	—	—	—	—	—	—	−0.124^a	(0.031)
Age at Immigration								
5-9 years	−0.226	(0.150)	−0.058	(0.169)	−0.103	(0.178)	−0.105	(0.164)
Difference from (1)	—	—	0.168	[0.026]	0.123	[0.138]	0.122	[0.140]
Difference from (2)	—	—	—	—	−0.045	[0.225]	−0.047	[0.184]
10-14 years	−1.005^a	(0.224)	−0.724^a	(0.228)	−0.812^a	(0.231)	−0.821^a	(0.231)
Difference from (1)	—	—	0.280	[0.001]	0.193	[0.038]	0.183	[0.046]
Difference from (2)	—	—	—	—	−0.088	[0.028]	−0.097	[0.018]
Number of Observations	76,151		76,151		76,151		76,151	
Number of Clusters	479		479		479		479	
R^2	0.257		0.265		0.267		0.267	
FEMALE CHILDREN								
Father's Schooling	—	—	0.193^a	(0.065)	0.164^a	(0.057)	0.153^b	(0.061)
Mother's Schooling	—	—	−0.018	(0.071)	0.052	(0.066)	0.060	(0.070)
Mean Source Schooling: Males	—	—	—	—	−0.124^a	(0.030)	—	—
Mean Source Schooling: Females	—	—	—	—	—	—	−0.114^a	(0.029)
Age at Immigration								
5-9 years	−0.306^b	(0.144)	−0.102	(0.161)	−0.161	(0.154)	−0.162	(0.154)
Difference from (1)	—	—	0.204	[0.006]	0.145	[0.078]	0.144	[0.089]
Difference from (2)	—	—	—	—	−0.059	[0.150]	−0.060	[0.098]
10-14 years	−1.089^a	(0.215)	−0.811^a	(0.207)	−0.915^a	(0.199)	−0.919^a	(0.207)
Difference from (1)	—	—	0.278	[0.001]	0.174	[0.073]	0.170	[0.070]
Difference from (2)	—	—	—	—	−0.104	[0.021]	−0.108	[0.021]
Number of Observations	74,898		74,898		74,898		74,898	
Number of Clusters	479		479		479		479	
R^2	0.232		0.240		0.242		0.242	

Notes: Bootstrap standard errors in parentheses, a, b, c denote statistical significance in a two-tailed test at the 1%, 5%, and 10% levels respectively. The bootstrap p-value for a 'dependent-samples' Wald test of the difference in the regression coefficients is shown in square brackets. The clustered bootstrap results are based upon 1,000 replications. The dependent variable is the age-adjusted, years of schooling for child immigrants that arrived between 0 and 14 years of age. All regressions include indicators for five year arrival cohort and indicators for seven regions of birth.

Table 6: Linear Regression: Years of Schooling for Child Immigrants, Counter-factual Parental Education

	Parental Schooling (1)		'Counter-factual' Parental Schooling (2)	
	Coeff.	Std. Err	Coeff.	Std. Err
MALE CHILDREN				
Father's Schooling	0.233^a	(0.069)	0.204^a	(0.063)
Mother's Schooling	-0.050	(0.081)	-0.025	(0.074)
Age at Immigration				
5-9 years	-0.058	(0.169)	-0.067	(0.169)
10-14 years	-0.724^a	(0.228)	-0.739^a	(0.221)
Number of Observations	76,151		76,151	
Number of Clusters	479		479	
R^2	0.265		0.264	
FEMALE CHILDREN				
Father's Schooling	0.193^a	(0.065)	0.162^a	(0.059)
Mother's Schooling	-0.018	(0.071)	0.011	(0.065)
Age at Immigration				
5-9 years	-0.102	(0.161)	-0.115	(0.158)
10-14 years	-0.811^a	(0.207)	-0.825^a	(0.204)
Number of Observations	74,898		74,898	
Number of Clusters	479		479	
R^2	0.240		0.240	

Notes: Bootstrap standard errors in parentheses, a, b, c denote statistical significance in a two-tailed test at the 1%, 5%, and 10% levels respectively. The clustered bootstrap results are based upon 1,000 replications. The dependent variable is the age-adjusted, years of schooling for child immigrants that arrived between 0 and 14 years of age. All regressions include indicators for five year arrival cohort and indicators for seven regions of birth.

Table 7: Linear Regression: Years of Schooling for Child Immigrants, Living at Home Sample

			Actual Parental Schooling		Proxy Parental Schooling			
					2000 Census		1970, 1980,& 1990 Census	
	(1)		(2)		(3)		(4)	
	Coeff.	Std. Err	Coeff.	Std. Err	Coeff.	Std. Err	Coeff.	Std. Err
MALE CHILDREN								
Father's Schooling	—	—	0.106^a	(0.015)	0.094^b	(0.041)	0.292^b	(0.121)
Mother's Schooling	—	—	0.082^a	(0.014)	0.090^b	(0.045)	−0.108	(0.120)
Age at Immigration								
5-9 years	−0.263	(0.188)	−0.214	(0.163)	−0.188	(0.162)	−0.213	(0.165)
Difference from (1)	—	—	0.049	[0.675]	0.075	[0.565]	0.050	[0.678]
Difference from (2)	—	—	—	—	0.025	[0.585]	0.001	[0.991]
10-14 years	-1.146^a	(0.259)	-0.998^a	(0.240)	-1.031^a	(0.245)	-1.025^a	(0.246)
Difference from (1)	—	—	0.148	[0.303]	0.115	[0.446]	0.122	[0.390]
Difference from (2)	—	—	—	—	−0.033	[0.622]	−0.026	[0.770]
Number of Observations	3,399		3,399		3,399		3,399	
Number of Clusters					281		281	
R^2	0.167		0.231		0.180		0.177	
FEMALE CHILDREN								
Father's Schooling	—	—	0.111^a	(0.016)	0.072	(0.055)	0.243^b	(0.103)
Mother's Schooling	—	—	0.064^a	(0.016)	0.062	(0.066)	−0.103	(0.097)
Age at Immigration								
5-9 years	−0.329	(0.234)	−0.203	(0.210)	−0.228	(0.230)	−0.238	(0.219)
Difference from (1)	—	—	0.125	[0.336]	0.100	[0.374]	0.091	[0.350]
Difference from (2)	—	—	—	—	−0.025	[0.667]	−0.035	[0.641]
10-14 years	-1.206^a	(0.350)	-1.044^a	(0.316)	-1.04^a	(0.343)	-1.082^a	(0.353)
Difference from (1)	—	—	0.161	[0.242]	0.166	[0.218]	0.124	[0.265]
Difference from (2)	—	—	0.005	[0.955]	−0.037	[0.670]		
Number of Observations	2,177		2,177		2,177		2,177	
Number of Clusters					250		250	
R^2	0.182		0.247		0.191		0.190	

Notes: Results are based upon a sample of immigrant children that arrived between 0 and 14 years of age and where still residing with their parents in the 2000 census. Bootstrap standard errors in parentheses, a, b, c denote statistical significance in a two-tailed test at the 1%, 5%, and 10% levels respectively. The clustered bootstrap results are based upon 1,000 replications. The bootstrap p-value for a 'dependent-samples' Wald test of the difference in the regression coefficients is shown in square brackets. The dependent variable is the age-adjusted, years of schooling for child immigrants that arrived between 0 and 14 years of age. All regressions include indicators for five year arrival cohort and indicators for seven regions of birth.

Figure 1: Marginal Effects for Age at Immigration, 10-14 years relative to 0-4 years, Males

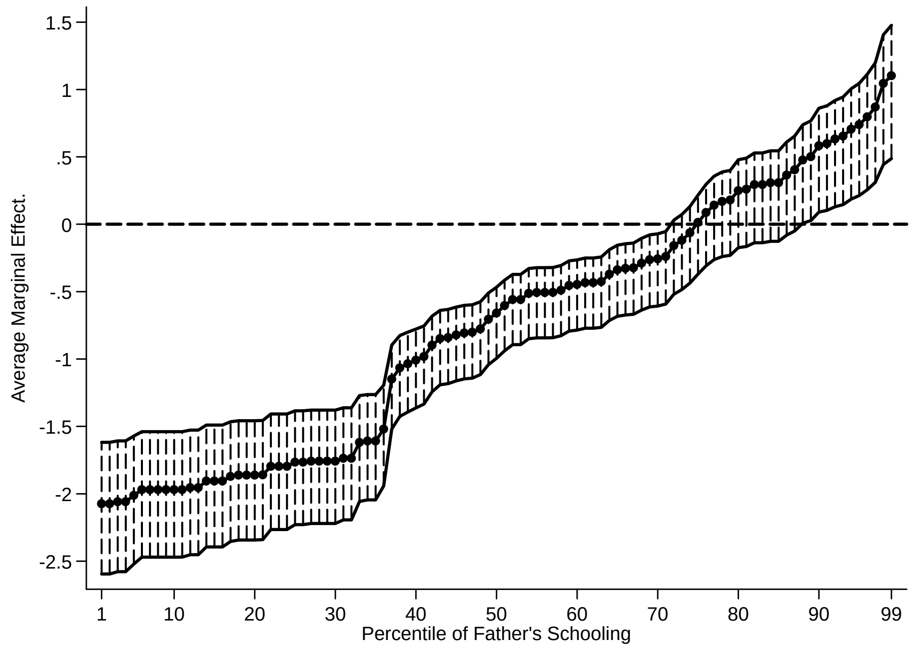
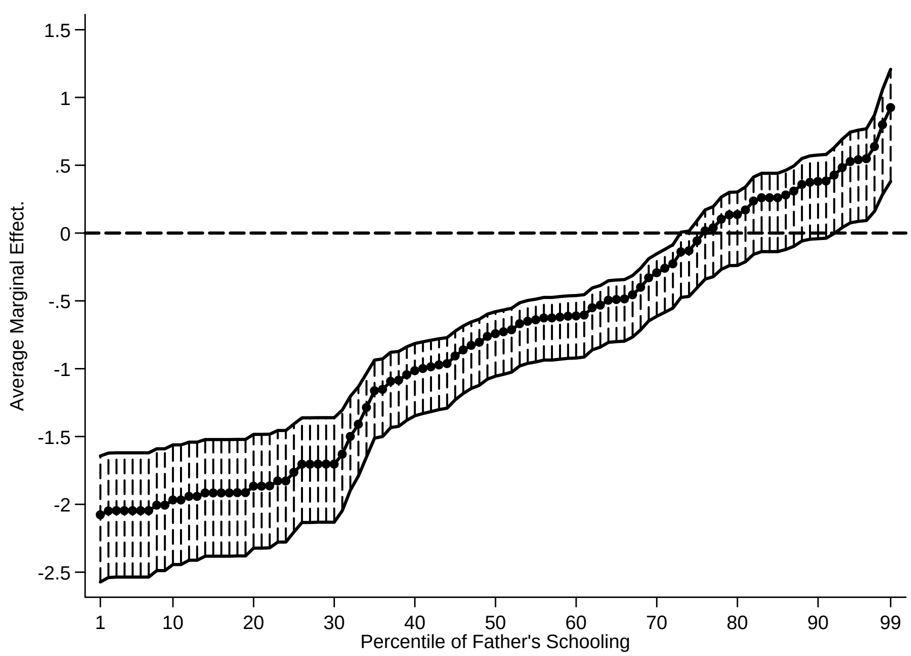


Figure 2: Marginal Effects for Age at Immigration, 10-14 years relative to 0-4 years, Females



Notes: Dashed lines represent the 95% confidence intervals around the marginal effects. The estimated model includes controls for mean source schooling (males) and interaction terms between father's schooling and age at immigration.

Supporting Information

Table A.1: Linear Regression: Years of Schooling for Male Child Immigrants

	(1)			(2)			(3)			(4)		
	Coeff.	Std. Err	p val.	Coeff.	Std. Err	p val.	Coeff.	Std. Err	p val.	Coeff.	Std. Err	p val.
Age at Immigration												
OLS Standard Errors												
5-9 years	−0.226	(0.029)	[0.000]	−0.058	(0.030)	[0.052]	−0.103	(0.030)	[0.001]	−0.105	(0.030)	[0.000]
10-14 years	−1.005	(0.031)	[0.000]	−0.724	(0.032)	[0.000]	−0.812	(0.033)	[0.000]	−0.821	(0.033)	[0.000]
Paired Bootstrap, Non-Cluster												
5-9 years	−0.226	(0.030)	[0.000]	−0.058	(0.029)	[0.048]	−0.103	(0.030)	[0.001]	−0.105	(0.029)	[0.000]
10-14 years	−1.005	(0.032)	[0.000]	−0.724	(0.030)	[0.000]	−0.812	(0.034)	[0.000]	−0.821	(0.033)	[0.000]
Cluster Robust Variance Estimator												
5-9 years	−0.226	(0.147)	[0.124]	−0.058	(0.171)	[0.735]	−0.103	(0.165)	[0.532]	−0.105	(0.165)	[0.527]
10-14 years	−1.005	(0.202)	[0.000]	−0.724	(0.213)	[0.001]	−0.812	(0.211)	[0.000]	−0.821	(0.211)	[0.000]
Paired Cluster Bootstrap												
5-9 years	−0.226	(0.150)	[0.132]	−0.058	(0.169)	[0.732]	−0.103	(0.178)	[0.563]	−0.105	(0.164)	[0.524]
10-14 years	−1.005	(0.224)	[0.000]	−0.724	(0.228)	[0.001]	−0.812	(0.231)	[0.000]	−0.821	(0.231)	[0.000]
Wild Bootstrap[a]												
5-9 years	−0.226	—	[0.188]	−0.058	—	[0.796]	−0.103	—	[0.624]	−0.105	—	[0.618]
10-14 years	−1.005	—	[0.002]	−0.724	—	[0.004]	−0.812	—	[0.002]	−0.821	—	[0.002]
Parental Education		No			Yes			Yes			Yes	
Mean Source Schooling: Males		No			No			Yes			No	
Mean Source Schooling: Females		No			No			No			Yes	
Number of Observations		76,151			76,151			76,151			76,151	
Number of Clusters		479			479			479			479	

[a] see Cameron & Miller (2015)

Notes: The clustered bootstrap results are based upon 1,000 replications. The p values for a two-tailed Wald test around zero are reported in square brackets.

Table A.2: Linear Regression: Years of Schooling for Female Child Immigrants

	(1)			(2)			(3)			(4)		
	Coeff.	Std. Err	p val.	Coeff.	Std. Err	p val.	Coeff.	Std. Err	p val.	Coeff.	Std. Err	p val.
Age at Immigration												
OLS Standard Errors												
5-9 years	−0.306	(0.028)	[0.000]	−0.102	(0.028)	[0.000]	−0.161	(0.028)	[0.000]	−0.162	(0.029)	[0.000]
10-14 years	−1.089	(0.030)	[0.000]	−0.811	(0.031)	[0.000]	−0.915	(0.032)	[0.000]	−0.919	(0.032)	[0.000]
Paired Bootstrap, Non-Cluster												
5-9 years	−0.306	(0.027)	[0.000]	−0.102	(0.029)	[0.000]	−0.161	(0.028)	[0.000]	−0.162	(0.028)	[0.000]
10-14 years	−1.089	(0.029)	[0.000]	−0.811	(0.030)	[0.000]	−0.915	(0.032)	[0.000]	−0.919	(0.032)	[0.000]
Cluster Robust Variance Estimator												
5-9 years	−0.306	(0.136)	[0.025]	−0.102	(0.157)	[0.516]	−0.161	(0.148)	[0.28]	−0.162	(0.150)	[0.281]
10-14 years	−1.089	(0.189)	[0.000]	−0.811	(0.195)	[0.000]	−0.915	(0.190)	[0.000]	−0.919	(0.191)	[0.000]
Paired Cluster Bootstrap												
5-9 years	−0.306	(0.144)	[0.034]	−0.102	(0.161)	[0.528]	−0.161	(0.154)	[0.296]	−0.162	(0.154)	[0.294]
10-14 years	−1.089	(0.215)	[0.000]	−0.811	(0.207)	[0.000]	−0.915	(0.199)	[0.000]	−0.919	(0.207)	[0.000]
Wild Bootstrap[a]												
5-9 years	−0.306	—	[0.028]	−0.102	—	[0.574]	−0.161	—	[0.364]	−0.162	—	[0.368]
10-14 years	−1.089	—	[0.002]	−0.811	—	[0.004]	−0.915	—	[0.002]	−0.919	—	[0.002]
Parental Education		No			Yes			Yes			Yes	
Mean Source Schooling: Males		No			No			Yes			No	
Mean Source Schooling: Females		No			No			No			Yes	
Number of Observations		74,898			74,898			74,898			74,898	
Number of Clusters		479			479			479			479	

[a] see Cameron & Miller (2015)

Notes: The clustered bootstrap results are based upon 1,000 replications. The p values for a two-tailed Wald test around zero are reported in square brackets.

Table A.3: Marginal Effects: Linear Regression Years of Schooling for Child Immigrants, Interacted Model

	Marginal Effect Additional Year of Parent's Schooling					
	coef.	Std. Error	p. val.	coef.	Std. Error	p. val.
Age at Immigration						
Father's Schooling						
Males						
0-4 years	0.048	(0.029)	[0.094]	0.084	(0.029)	[0.003]
5-9 years	0.156	(0.026)	[0.000]	0.188	(0.027)	[0.000]
10-14 years	0.369	(0.036)	[0.000]	0.396	(0.041)	[0.000]
Females						
0-4 years	0.050	(0.026)	[0.050]	0.077	(0.026)	[0.003]
5-9 years	0.164	(0.026)	[0.000]	0.186	(0.027)	[0.000]
10-14 years	0.346	(0.034)	[0.000]	0.363	(0.035)	[0.000]
Mother's Schooling						
Males						
0-4 years	0.004	(0.036)	[0.915]	0.046	(0.035)	[0.190]
5-9 years	0.126	(0.038)	[0.001]	0.166	(0.039)	[0.000]
10-14 years	0.347	(0.054)	[0.000]	0.381	(0.055)	[0.000]
Females						
0-4 years	0.017	(0.032)	[0.591]	0.062	(0.034)	[0.064]
5-9 years	0.135	(0.038)	[0.000]	0.178	(0.038)	[0.000]
10-14 years	0.350	(0.046)	[0.000]	0.386	(0.049)	[0.000]
Mean Source Schooling: Males		No			Yes	

Notes: Bootstrap standard errors in parentheses. The clustered bootstrap results are based upon 1,000 replications. The dependent variable is the age-adjusted, years of schooling for child immigrants that arrived between 0 and 14 years of age. All regressions include indicators for five year arrival cohort and indicators for seven regions of birth. The estimated model contains interactions of father's years of schooling with age at immigration.

Table A.4: Linear Regression: Years of Schooling for Child Immigrants, Parental Schooling Interacted With Source Regions

	(1)		(2)		(3)		(4)	
	Coeff.	Std. Err	Coeff.	Std. Err	Coeff.	Std. Err	Coeff.	Std. Err
MALE CHILDREN								
Age at Immigration								
5-9 years	-0.226	(0.154)	-0.110	(0.175)	-0.176	(0.161)	-0.180	(0.165)
10-14 years	-1.005^{a}	(0.221)	-0.717^{a}	(0.208)	-0.799^{a}	(0.201)	-0.809^{a}	(0.199)
Father's Schooling*Region	No		Yes		Yes		Yes	
Mother's Schooling*Region	No		Yes		Yes		Yes	
Mean Source Schooling: Males	No		No		Yes		No	
Mean Source Schooling: Females	No		No		No		Yes	
Number of Observations	76,151		76,151		76,151		76,151	
Number of Clusters	479		479		479		479	
FEMALE CHILDREN								
Age at Immigration								
5-9 years	-0.306^{b}	(0.144)	-0.133	(0.159)	-0.247^{c}	(0.147)	-0.248	(0.151)
10-14 years	-1.089^{a}	(0.215)	-0.827^{a}	(0.193)	-0.984^{a}	(0.182)	-0.985^{a}	(0.186)
Father's Schooling*Region	No		Yes		Yes		Yes	
Mother's Schooling*Region	No		Yes		Yes		Yes	
Mean Source Schooling: Males	No		No		Yes		No	
Mean Source Schooling: Females	No		No		No		Yes	
Number of Observations	74,898		74,898		74,898		74,898	
Number of Clusters	479		479		479		479	

Notes: Bootstrap standard errors in parentheses, a, b, c denote statistical significance in a two-tailed test at the 1%, 5%, and 10% levels respectively. The bootstrap p-value for a 'dependent-samples' Wald test of the difference in the regression coefficients is shown in square brackets. The clustered bootstrap results are based upon 1,000 replications. The dependent variable is the age-adjusted, years of schooling for child immigrants that arrived between 0 and 14 years of age. All regressions include indicators for five year arrival cohort and indicators for seven regions of birth.

Table A.5: Linear Regression: Years of Schooling for Child Immigrants, Controls for English Deficiency

	(1)		(2)		(3)		(4)	
	Coeff.	Std. Err	Coeff.	Std. Err [p val.]	Coeff.	Std. Err [p val.]	Coeff.	Std. Err [p val.]
MALE CHILDREN								
Father' Schooling	—	—	0.225^a	(0.060)	0.225^a	(0.054)	0.210^a	(0.055)
Mother's Schooling	—	—	−0.111	(0.072)	−0.072	(0.066)	−0.057	(0.066)
Mean Source Schooling: Males	—	—	—	—	-0.116^a	(0.029)	—	—
Mean Source Schooling: Females	—	—	—	—	—	—	-0.114^a	(0.027)
Age at Immigration								
5-9 years	−0.105	(0.120)	0.017	(0.134)	−0.024	(0.139)	−0.026	(0.127)
Difference from (1)	—	—	0.122^a	[0.031]	0.081	[0.206]	0.079	[0.218]
Difference from (2)	—	—	—	—	−0.041	[0.190]	−0.043	[0.186]
10-14 years	-0.548^a	(0.173)	-0.357^b	(0.175)	-0.438^b	(0.177)	-0.447^b	(0.174)
Difference from (1)	—	—	0.191^a	[0.003]	0.110	[0.124]	0.101	[0.155]
Difference from (2)	—	—	—	—	-0.081^b	[0.029]	-0.090^b	[0.017]
Number of Observations	76,151		76,151		76,151		76,151	
Number of Clusters	479		479		479		479	
R^2	0.320		0.324		0.324		0.326	
FEMALE CHILDREN								
Father' Schooling	—	—	0.192^a	(0.054)	0.166^a	(0.050)	0.155^a	(0.052)
Mother's Schooling	—	—	−0.082	(0.061)	−0.019	(0.058)	−0.011	(0.061)
Mean Source Schooling: Males	—	—	—	—	-0.112^a	(0.027)	—	—
Mean Source Schooling: Females	—	—	—	—	—	—	-0.105^a	(0.026)
Age at Immigration								
5-9 years	−0.173	(0.114)	−0.016	(0.124)	−0.069	(0.115)	−0.071	(0.110)
Difference from (1)	—	—	0.157^a	[0.103]	0.103^c	[0.099]	0.101	[0.108]
Difference from (2)	—	—	—	—	−0.053	[0.126]	-0.055^c	[0.096]
10-14 years	-0.578^a	(0.147)	-0.383^b	(0.151)	-0.479^a	(0.143)	-0.484^a	(0.140)
Difference from (1)	—	—	0.194^a	[0.001]	0.098	[0.170]	0.093	[0.203]
Difference from (2)	—	—	—	—	-0.096^b	[0.018]	-0.101^b	[0.023]
Number of Observations	74,898		74,898		74,898		74,898	
Number of Clusters	479		479		479		479	
R^2	0.300		0.304		0.306		0.306	

Notes: Bootstrap standard errors in parentheses, a, b, c denote statistical significance in a two-tailed test at the 1%, 5%, and 10% levels respectively. The bootstrap p-value for a 'dependent-samples' Wald test of the difference in the regression coefficients is shown in square brackets. The clustered bootstrap results are based upon 1,000 replications. The dependent variable is the age-adjusted, years of schooling for child immigrants that arrived between 0 and 14 years of age. All regressions include indicators for five year arrival cohort and indicators for seven regions of birth, and an indicator for English deficiency.

Figure A.1: Marginal Effects for Age at Immigration, 10-14 years relative to 0-4 years, Males

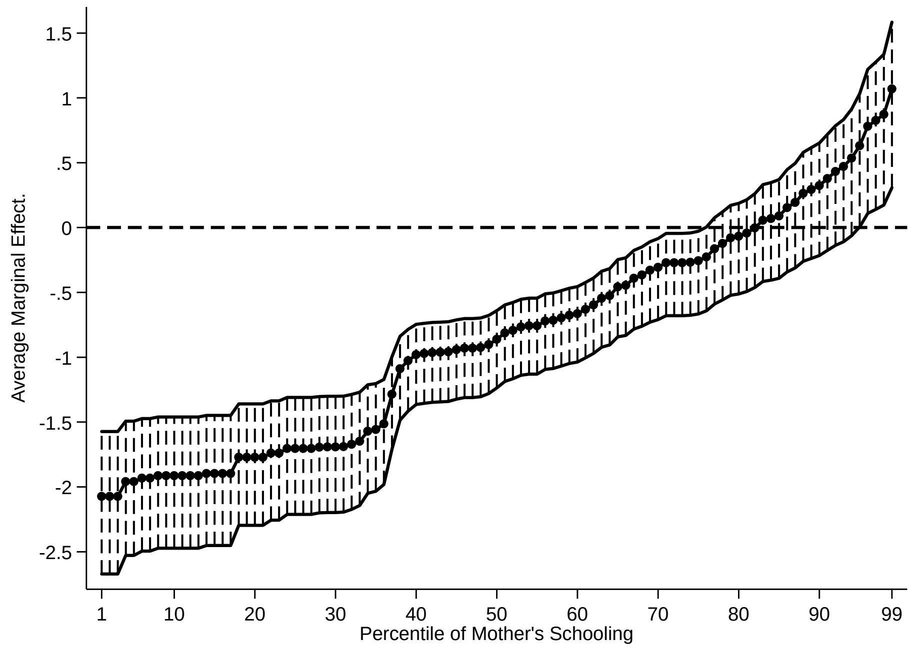
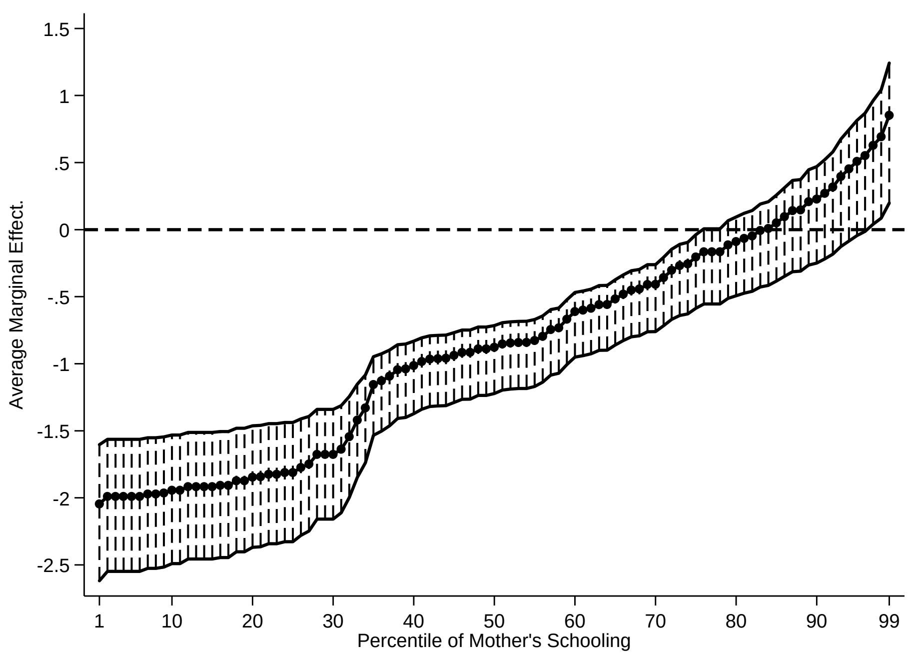


Figure A.2: Marginal Effects for Age at Immigration, 10-14 years relative to 0-4 years, Females



Notes: Dashed lines represent the 95% confidence intervals around the marginal effects. The estimated model includes controls for mean source schooling (males) and interaction terms between mother's schooling and age at immigration.