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Mobility of the 1.5 Generation

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Age at Immigration and the Intergenerational Income Mobility of the 1.5 Generation *

Marie Connolly[†], Catherine Haeck[‡], Anne Mei Le Bourdais-Coffey[§]

October 20, 2025

Abstract

The present study exploits longitudinal tax files linked to Census data to measure the contribution of age at immigration to the intergenerational income mobility of immigrant children. We first estimate the causal effect of children's age at immigration on adulthood income using a siblings fixed effects model of years of exposure to the host country. Up to 10 years old, the relationship between age at immigration and income is weak, but starting at age 11, each additional year is associated with a decrease in adulthood income rank of close to half a percentile rank. We then find that adjusting the 1.5 generation's income ranks for age at immigration results in an intergenerational rank-rank coefficient that is lower by 0.018, or 10.2% of the (unadjusted) intergenerational income transmission estimate of the 1.5 generation. Earlier immigration has the potential to improve intergenerational economic mobility.

Keywords: intergenerational income mobility, immigrants, 1.5 generation, age at immigration, Canada

JEL Codes: J62, J61, J15

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to give greater consideration to research that outlines economic opportunities for immigrant children, and how age at immigration is an important factor that may play a role in shaping those opportunities.

Our main objective is thus to estimate the extent to which age at immigration explains the intergenerational mobility of children of the 1.5 generation. Our work is based on a recent data linkage between Canadian intergenerationally-linked tax files and Census data (Connolly et al., 2022; Haeck & Laliberté, 2025), especially well-suited to investigate this question. While previous studies have sought to estimate the effect of age at immigration on different outcomes, none has explored its role in the intergenerational income mobility of immigrant children. We therefore contribute not only to the literature on age-at-immigration effects, but also to that on equality of opportunities and the factors that can explain the intergenerational transmission processes that we observe. A better understanding of those processes can help us design public policies that foster greater socioeconomic mobility.

Our analysis is motivated by the relative paucity of studies on the intergenerational mobility of immigrants using intergenerationally-linked fiscal records, both in Canada and elsewhere. Due to data unavailability, previous studies on intergenerational mobility in Canada have typically looked at the population as a whole, rather than examining heterogeneity by immigrant status (Corak & Heisz, 1999; Chen et al., 2017; Connolly et al., 2019). Earlier studies that have focused on immigrants often had to resort to using group averages due to a lack of available microdata (e.g. Aydemir et al. (2009) and Aydemir et al. (2013) on Canada or Borjas (1992, 1993) and Card et al. (1998) on the United States).

More recently, the growing availability of administrative fiscal data and their various linkages have opened the door to studies based on immigrant status, though most concentrate on second-generation immigrants. Using Canadian data, Connolly et al. (2022) show that second-generation children with immigrant mothers have higher levels of intergenerational mobility than those with mothers born in Canada, but focus the rest of their analysis on the latter group and ignore the 1.5 generation. Adnan et al. (2023) study intergenerational income mobility among the 1.5 generation using Statistics Canada's Longitudinal Immigration

Database (IMDB), yet mainly focus on those with refugee status, whose parents are less positively selected. In comparison, our data cover all people in Canada, not just the 1.5 generation, allowing us to use second-generation immigrants as a comparison group. In the United States, [Abramitzky et al. \(2021\)](#) find that immigrants tend to have higher upward mobility than nonimmigrants, a fact true both historically and in more recent times, but they cannot distinguish between the children of immigrants born in the US and those born abroad. Their conclusion is at odds with what [Jensen & Manning \(2025\)](#) find using register data from Denmark. The contradiction prompted the larger study of [Boustan et al. \(2025\)](#), who offer a comparison of the intergenerational income mobility of second-generation immigrants born around the early 1980s in 15 countries. Our contribution to this literature is to focus on the 1.5 generation and highlight the role that age at immigration may have on their intergenerational income mobility.

We also contribute to the literature on age-at-immigration effects by providing the first estimates of such effects on adulthood income using a credible causal identification strategy applied to Canadian microdata. Studies on the effect of age at immigration on outcomes such as education and earnings show that immigrant children's age at arrival has a negative effect on said outcomes ([Bleakley & Chin, 2004](#); [Van Ours & Veenman, 2006](#); [Böhlmark, 2008](#); [Myers et al., 2009](#); [Beck et al., 2012](#); [Corak, 2012](#); [Åslund et al., 2015](#); [Clarke, 2018](#); [Lemmermann & Riphahn, 2018](#); [Basu, 2018](#); [Ansala et al., 2020](#); [Hull, 2023](#), for education and [Schaafsma & Sweetman, 2001](#); [Pendakur & Pendakur, 2016](#); [Hermansen, 2017](#); [Pekkala Kerr et al., 2024](#), for earnings). Most also agree that a critical point exists between the ages of eight and 12; after this age, there is a sharp and continued decrease in the observed measure. Only a handful of those studies exploit within-family variations, like we do. Our work employs a state-of-the-art methodology to tackle the issue of selection into migration and cleanly identify exposure effects.

We start by estimating the causal effect of age at immigration using a siblings fixed effects model, leveraging the variation in the outcomes of children from the same family, who immigrated to Canada in the same year, but at different ages. Our main outcome is child income rank at ages 27 to 31 years old. Using siblings fixed effects allows us to control for the socioeconomic status of the family and the

context in which the family immigrated, and second-generation children are used as a reference group. Our framework to assess the “exposure-to-Canada” effect is inspired by Chetty & Hendren (2018), as were others like Aloni & Avivi (2025). Our findings suggest that up until the age of 10, the relationship between age at immigration and income rank is negative but relatively weak. Then, starting at the age of 10, each year of belated arrival is associated with an average decline of 0.42 percentile rank.

We then document the intergenerational income transmission among immigrants and show that nonimmigrants have the strongest correlation between parental and child income rank (with a rank-rank slope of 0.250) compared to the 1.5 generation (0.176) and second generation (0.166), meaning the children of immigrants experience higher income mobility than those born from Canadian parents.

We reach our main conclusion through a counterfactual exercise in which we first adjust the adulthood income ranks of 1.5-generation children by subtracting from their observed ranks the age-at-immigration effects estimated with our siblings fixed effects model. We then re-estimate the intergenerational rank-rank relationship, and find that the income rank transmission coefficient falls by 0.018, which is 10.2% of the unadjusted rank-rank correlation for the 1.5 generation in our data. Our findings suggest that the older the 1.5 generation children are at the time of immigration, the more difficult it is for them to integrate into society. Policies aimed at bringing children in at a younger age and helping older children integrate, particularly through language and school support, may therefore have a positive impact on their future income and their economic mobility.

The remainder of this paper is organized as follows. Section 2 describes the data and offers descriptive statistics. Section 3 presents our methodology. Section 4 presents our estimates of age-at-immigration effects, complemented by robustness checks in Subsection 4.1 and a heterogeneity analysis in Subsection 4.2. We follow with intergenerational mobility differences across generation status and the contribution of age at immigration in Section 5 and offer a discussion in Section 6. Section 7 concludes.

2 Data

We use data from Statistics Canada called the IID+, which consists of the Intergenerational Income Database (IID) linked to de-identified Census data (Statistics Canada, 2016b,a).³ The IID provides annual tax records from 1978 to 2016 for children born between 1963 and 1985 and their parents. Tax data originate from the T1 forms submitted annually by tax filers to the Canada Revenue Agency. The Canadian Census of Population provides supplementary information such as immigration status and place of birth.

Based on Statistics Canada's definition of children with an immigrant background and Rumbaut's (1991) definition of the 1.5 generation, we classify as part of the 1.5 generation individuals who are born outside of Canada, have at least one immigrant parent, and arrive in Canada before adulthood (before 18 years of age). Second-generation children are born in Canada and have at least one immigrant parent. Third-generation children and above are born in Canada and have nonimmigrant parents. Consequently, determining children's generation status requires Census data on both children and parents. However, given that 20 to 25% of Canadians complete the long-form Census, it is not possible to find all IID individuals in the Census data. To maximize the number of successful matches, five Census waves, spanning from 1996 to 2016, have been linked to the IID. Table 1 presents the distribution of the immigrant generation status by IID birth cohort as well as the percentage of IID children that could not be linked to any of the Census waves (59% overall). Our final analysis sample includes almost three million child-parents pairs for which we have the longitudinal tax records of the parents and the child (from the IID) and the generation status of the child (from the Census).

Our main outcome is child income rank, in which total income is averaged over a period of five years, when the child is aged 27 to 31, and ranks are assigned within child birth year. Parental income rank is computed on average income of both parents when the child is aged 15 to 19. A more complete description of the variables can be found in Section A.2 of the Appendix.

³Section A.1 of the Appendix contains more information on the IID+. Connolly et al. (2022) were the first to use this linkage.

Table 1: Distribution of Immigrant Generation Status and Census Linkage Rates for IID Children, by Birth Cohort

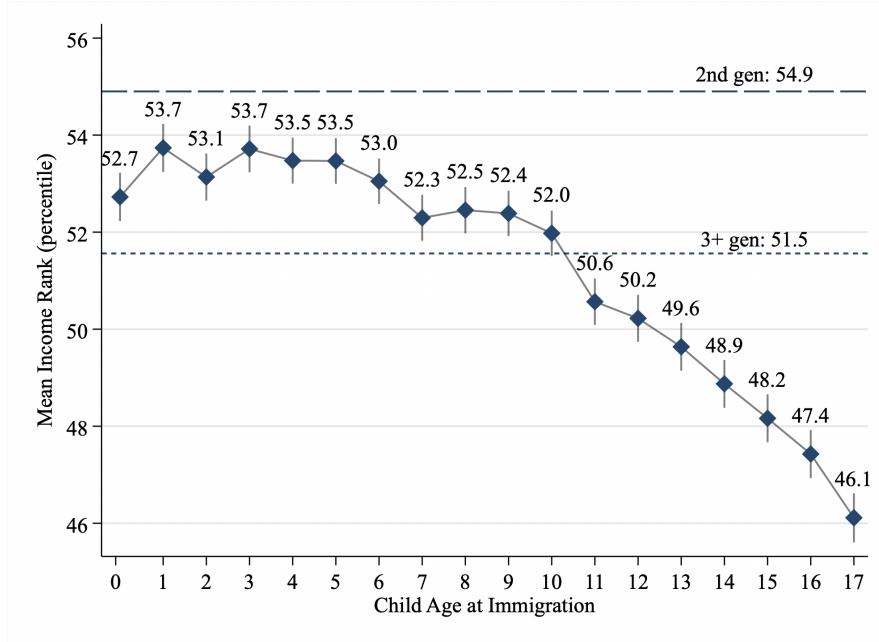
Birth cohort	Generation status			Not in sample	Unlinked	IID weighted count
	1.5	2	3+			
1963 to 1966	3.7	5.0	20.7	5.2	65.3	1,591,000
1967 to 1970	3.6	6.0	21.9	5.2	63.3	1,571,790
1972 to 1975	3.9	7.5	24.6	4.9	59.1	1,483,800
1977 to 1980	4.8	8.4	27.4	4.2	55.1	1,558,390
1982 to 1985	5.3	8.3	30.1	4.0	52.3	1,634,620
All	4.3	7.1	25.0	4.7	59.0	7,839,600

Source: Authors' calculations based on the IID+.

Note: This table presents the distribution of immigrant generation status and Census linkage rates (in percentages) by birth cohort. Generation status is determined using information from the linkage between the IID and the Census. Non-permanent residents, individuals with an age at immigration above 17 years old, individuals living in Canada's territories or with missing information are not in the analysis sample (Not in sample, column 4). Unlinked children include all IID individuals who could not be linked to any Census wave. The last column shows the weighted counts of children within the IID.

We complete our presentation of the data by showing, in Figure 1, the association between the mean income rank of 1.5-generation children and age at arrival. This figure reports a negative relationship between mean income rank and age at immigration. Children who immigrated at age 0 have an average income rank in adulthood of 52.7, a number that decreases to 46.1 for those who arrived at 17 years of age. This represents a drop of 6.6 ranks, and 8.8 ranks below the average income rank of children of the second generation (54.9, std. err. = 0.04). To put things in perspective, this 8.8 percentile rank difference equals about CAD\$6,200 (in 2025 dollars, about US\$4,500), going from CAD\$38,800 to CAD\$32,600. Figure 1 also shows that the mean income rank of children of the 1.5 generation who immigrated before the age of 11 years old is above that of nonimmigrant children (51.5, std. err. = 0.02). This is coherent with the descriptive evidence presented in Statistics Canada (2025), who also add that participation in postsecondary education is higher for those who immigrated at younger ages, particularly for the children of economic immigrants (as opposed to immigrants that were sponsored by their family or refugees). More descriptive statistics (by region of origin) are presented in Section A.3 of the Appendix.

Figure 1: Mean Child Income Ranks by Child Age at Immigration among the 1.5 Generation



Source: Authors' calculations based on the IID+.

Note: The horizontal lines represent the mean income rank for second-generation children (54.9, std. err. = 0.04, in long dashes) and nonimmigrants (51.5, std. err. = 0.02, in short dashes). Mean income percentile ranks at each child age at immigration are shown above each point. The spikes represent 95% confidence intervals.

3 Methodology

In this paper, our main objective is to estimate the impact of children's age at immigration on their income in adulthood and evaluate the contribution of age at immigration to the transmission of income between parents and children. We first describe our approach to estimate the impact of age at immigration on average income rank, and then present our income mobility framework.

We start with a simple model in which a child's income rank r_i is explained by a constant a and dummies indicating the child's age at immigration m_i :

$$r_i = a + \sum_{m=0}^{17} b_m I(m_i = m) + \nu_i \quad (1)$$

We estimate this model on a sample including children from the 1.5 generation as well as children from the second generation, such that the coefficients on age at immigration are interpreted as differences between second-generation children and those who immigrated at the corresponding age. This is a purely descriptive exercise and the b_m estimated using Equation 1 should not be interpreted causally because of potential omitted variables bias.

To capture the causal impact of age at immigration on adulthood income, we have in mind Chetty & Hendren's (2018) movers approach, in which they estimate childhood exposure effects by comparing outcomes of children who move across commuting zones or counties in the US to those of the children who grew up in the places where they move to. We adapt their movers approach to an international immigration context and compare children who immigrated to Canada at different ages (the 1.5 generation) to those that were born in Canada, but with immigrant parents (the second generation).

Chetty & Hendren (2018)'s design relies on the hypothesis that *when* people move is not correlated with the error term, an assumption that has been criticized by some (Heckman & Landersø, 2022; Eshaghnia, 2023). In the context of our study, this may be a strong assumption given that parents of young children could account for their children's age when they decide to immigrate. If parents with better unobservable characteristics immigrate to Canada when their children are younger or if parents account for their children's age when making such decisions, the assumption that selection is not a function of age at immigration would be violated. However, one could argue that in the case of international immigration, a considerable degree of uncertainty is often associated with the precise timing of the relocation. Administrative delays and processes, as well as conflicts emerging in the country of origin, can influence the timing. If children's ages do not influence the timing of immigration, then we would expect the distribution of age at arrival to be uniform. In our data, children who arrive before age two and after the age of 15 appear slightly underrepresented, but overall the distribution is fairly uniform with no clear selection pattern.⁴

⁴See Appendix Table C2 for the distribution of the age at immigration among the 1.5 generation. We also find that parental schooling attainment, the number of siblings, and sex are distributed fairly uniformly across age at arrival, while region of origin is not.

To adopt a cleaner identification strategy, we turn to a siblings fixed effects approach, in which we identify the impact of age at immigration by comparing siblings within a family arriving at different ages.⁵ This allows us to better account for socioeconomic differences and net out potential selection effects. We add siblings fixed effects a_s to the model in Equation 1 and estimate the following:

$$r_i = \sum_{m=0}^{17} b_m I(m_i = m) + a_s + e_i \quad (2)$$

In the IID, siblings are identified using the mother and father identifiers. However, siblings born outside the IID target years are not present in the IID. For example, an individual born in 1963 with one sibling born in 1960 will be considered an only child in the IID, since the earliest target birth year is 1963. This will result in the misclassification of some families with multiple children as one-child families, but not the reverse: identified siblings do share parents. However, since the misclassification criteria should be orthogonal to family characteristics (selection is purely based on birth years), we argue that our final siblings sample is representative of multi-siblings families in Canada. Our final sample of children is approximately three times smaller than our baseline analysis sample.

In the above models, the vector of b_m can be decomposed into two parts (Chetty & Hendren, 2018):

$$b_m = \beta_m + \delta_m \quad (3)$$

where β_m is the causal effect of arriving at age m in Canada and δ_m is a selection effect. The exposure effect (i.e. spending one more year in Canada) at age m is defined as follows:

$$\gamma_m = \beta_m - \beta_{m+1} \quad (4)$$

⁵Other contributions using a siblings fixed effects approach in similar analyses include Böhlmark (2008); Åslund et al. (2015); Hermansen (2017); Lemmermann & Riphahn (2018); Basu (2018); Ansala et al. (2020).

As mentioned above, the timing of the move may remain correlated with the error term in the model of Equation 1. However, our approach via siblings fixed effects means our estimation of Equation 2 should directly identify the bias-free β_m instead of b_m . The idea is that the timing of *when* parents decided to move may be correlated with age, but not within a siblings group. In other words, we recognize the existence of a selection effect ($\delta_m \neq 0$) and we allow it to be a function of age at immigration, but assume that $\delta_m = \delta$ within a siblings group. With the selection effects canceling out between siblings, the exposure-to-Canada effect can be directly extracted from the coefficients estimated with Equation 2:

$$\gamma_m = b_m - b_{m+1} \quad (5)$$

To estimate the contribution of age at immigration to intergenerational income mobility, we start by documenting rank mobility using a standard model in which child income rank (r_i) is explained by parental income rank (p_i), such that $r_i = \alpha + \beta p_i + \epsilon_i$ (Chetty et al., 2014). We estimate this model using ordinary least squares. The slope coefficient of that model (β) gives the correlation between parental income rank and child income rank. The higher the coefficient is, the stronger the intergenerational income transmission, and the lower the intergenerational mobility is.⁶

Next, to approximate the extent to which rank mobility for the 1.5 generation is explained by age at arrival, we adjust the observed child income rank (r_{im}) by subtracting the age-at-immigration coefficient from the siblings fixed effects estimation (\hat{b}_m) that corresponds to their age at immigration m . The adjusted ranks ($r_{im}^{adj} = r_{im} - \hat{b}_m$) thus corresponds to what we would expect each child's income rank to be had they been born in Canada from immigrant parents (i.e., the reference group in Equation 2). We then re-estimate the rank mobility, but using the adjusted income ranks as the explained variable, yielding a slope estimate of β^{adj} . The difference between β^{adj} and β gives us the contribution of age at immigration, or how different the intergenerational income mobility would be if

⁶For more on equality of opportunities, see Corak (2013).

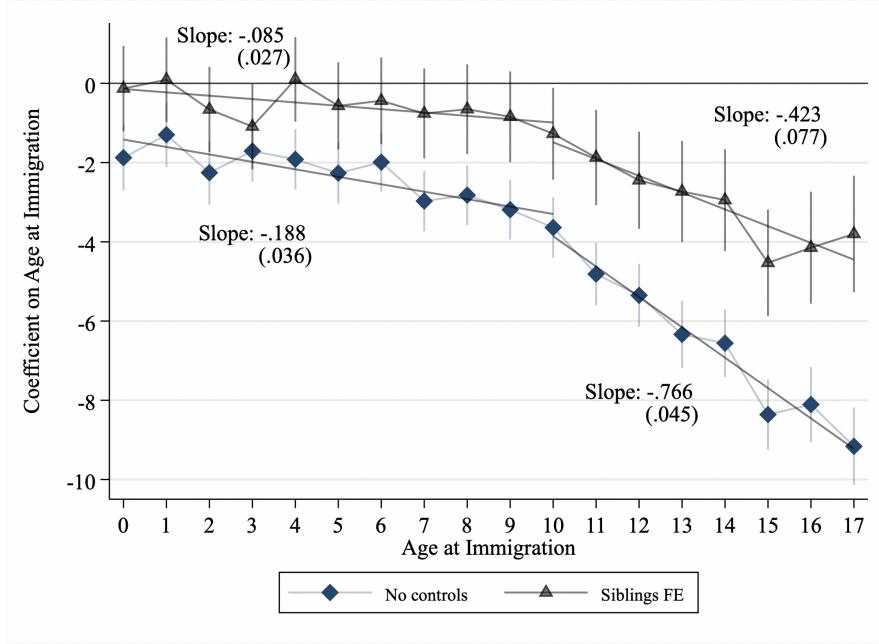
children of immigrants were all born in Canada. This exercise is informative when trying to evaluate the contribution of age at immigration in a straightforward way. Note that we always compute robust standard errors and do not cluster them at any level since the sampling process is not clustered, and the assignment mechanism is not clustered either (Abadie et al., 2022).

4 Estimates of Age-at-Immigration Effects

We start this section with our main estimates of the age-at-immigration effects, and follow with robustness checks and heterogeneity analyses. Figure 2 plots our baseline estimates of b_m for two specifications. In the first series (the blue diamonds), the child's income rank is only explained by age at arrival dummies without any other control variables, corresponding to Equation 1. This first series is a regression equivalent to the descriptive pattern displayed in Figure 1, except that the latter was computed using the full sample, whereas the one in Figure 2 is estimated on the siblings sample. The patterns are almost identical. In the second series (the gray triangles), we add siblings fixed effects; this is the model of Equation 2.

The estimates presented in Figure 2 display two key patterns. First, $\hat{b}_m < 0$ for all m in the base model without controls. This implies that children of the 1.5 generation have, on average, lower income ranks in adulthood relative to comparable children of the second generation. Second, this correlation is partially explained by differences across families, since estimates accounting for siblings fixed effects are smaller and not statistically different from zero before 10 years of age. It thus appears that age at arrival only impacts adulthood income starting at 10 years old, and each additional year of delay in arrival further increases the impact on income.

Figure 2: Age-at-Immigration Effects on Income Ranks in Adulthood



Source: Authors' calculations based on the IID+.

Note: This figure plots the coefficients b_m for each age at immigration m , with each series coming from a separate regression: first, only the age-at-immigration dummies (Equation 1), then adding siblings fixed effects (Equation 2). The sample includes all children of the 1.5 generation whose age at immigration is between 0 and 17, inclusively, as well as children from the second generation (the reference group) who have at least one sibling (to allow for the siblings fixed effects estimation). Child income is measured at ages 27 to 31. Parental income is measured when the child is 15 to 19 years old. The spikes represent 95% confidence intervals, based on robust standard errors. For each series, we add fitted lines through the \hat{b}_m coefficients using a piecewise (spline) regression with a kink at age 10. Estimated slopes for each segment are shown next to the fitted lines, with their standard errors in parentheses.

Following Equation 5, we can see the exposure effects $\hat{\gamma}_m$ by comparing the point estimates of \hat{b}_m from each age with \hat{b}_{m+1} , the point estimate for the next age at arrival. We observe on Figure 2 that $\hat{\gamma} \approx 0$ for $m \leq 10$ for the siblings fixed effects only specification. Then, as children arrive at older ages, each year of delayed immigration to Canada brings lower income ranks on average. To estimate γ ,⁷ we fit a piecewise linear model with a kink at age 10. This provides

⁷Technically, the linear model gives the $-\gamma$, since γ_m is defined as $b_m - b_{m+1}$.

the average effect of each additional year not spent in Canada between 0 and 10 years old (first segment), and between 10 and 17 years old (second segment). The resulting fitted lines can be seen in Figure 2. We observe that, although the 1.5 generation is at a disadvantage relative to second-generation children (model without controls), age at immigration has a fairly small effect on adulthood income if immigration happens no later than 10 years old when siblings fixed effects are accounted for: the slope for that segment is -0.085 (std. err. = 0.027). Starting at age 10, however, each additional year not spent in Canada is associated with a decrease in income rank of about 0.8 percentile rank when using the model without controls. Adding siblings fixed effects reduces that (non)exposure effect to -0.423 (std. err. = 0.077), meaning that immigrating at older ages has a negative effect on adulthood income. In dollar terms, this difference of 0.423 ranks represents about CAD\$300 (in 2025 dollars, about US\$225). The income rank of a child arriving at 17 years old will be, on average, 4 percentile ranks lower relative to a child who arrived between 0 and 10 years old (equivalent to about CAD\$2,600⁸ or US\$1,900). These results align with previous literature on the effects of age at immigration on outcomes such as earnings, education, and language proficiency (Arellano-Bover et al., 2025). As mentioned above in the introduction, past research showed that a turning point exists between the ages of eight and 12. Age at arrival has a pronounced detrimental impact on these outcomes when arrival occurs beyond this critical period. However, before reaching this pivotal stage, age at arrival demonstrates minimal influence on the observed outcome.

Our results also echo the findings of Chetty & Hendren (2018), who find that when families relocate to more favorable commuting zones or counties, the outcomes of the children improve progressively and consistently. This linear improvement is proportional to the duration of time the children spend growing up in the new area, with an approximate rate of 4% per year of exposure. Our results are also in line with what Deutscher (2020) estimates when replicating Chetty & Hendren (2018) in the Australian context. Notably, Deutscher (2020) is able to consider moves at earlier ages than Chetty & Hendren (2018), and finds that the

⁸Around the median of the income distribution.

expected ranks of children who relocate converge to those of permanent residents in their destination neighborhood at a rate of approximately 1.1% per year for each year spent in the neighborhood prior to the age of 11. However, after the age of 11, this rate of convergence increases substantially to approximately 4.2% per year. These findings, based on internal migrations and with their piecewise linear relationship, are coherent with ours but our convergence rates relative to second-generation children are much smaller, at 0.15% before age 10 and 0.77% after. Canada's lower convergence rates, combined with the fact that on average 1.5 generation children do better than natives of the 3+ generation before age 10, may imply that Canada is a relatively better place for immigrant children.

4.1 Robustness checks

We perform robustness checks, the results of which can be found in the Appendix. We start our robustness checks by estimating Equations 1 and 2 using alternative specifications, samples, and definitions (see Appendix Figure B2). First, we classify children as immigrants when both parents are immigrants, whereas our baseline definition only requires one immigrant parent. Second, we use child income ranks measured at ages 30 to 34 instead of 27 to 31. Note that parental income does not directly enter in the estimations of Figure 2, so a robustness check using a different definition of parental income would give the exact same estimates. By and large, the estimates of \hat{b}_m do not differ much from the baseline seen in Figure 2.

In another set of robustness checks, we estimate a model with origin, destination, cohort, and parental income decile groups instead of the siblings fixed effects (see Appendix Figure B3). We first note that controlling only for origin, destination, and birth cohort does not introduce statistically significant differences with the no-controls estimates. However, adding parental income closes much of the gap with second-generation immigrants, meaning that parental income accounts for a large share of the income gaps, a finding that echoes the cross-country work of Boustan et al. (2025). We further observe that these estimates (with the full set of controls) are very similar to our main results based on siblings fixed effects, shown in the series of gray triangles in Figure 2. Parental income deciles must

not vary much within families, and we argue that our siblings fixed effects strategy provides a cleaner identification of the age-at-immigration effects. That being said, the general pattern in the exposure effects that we observe in Figure 2 is the same as that observed here, with stronger effects the later a child immigrates to Canada.

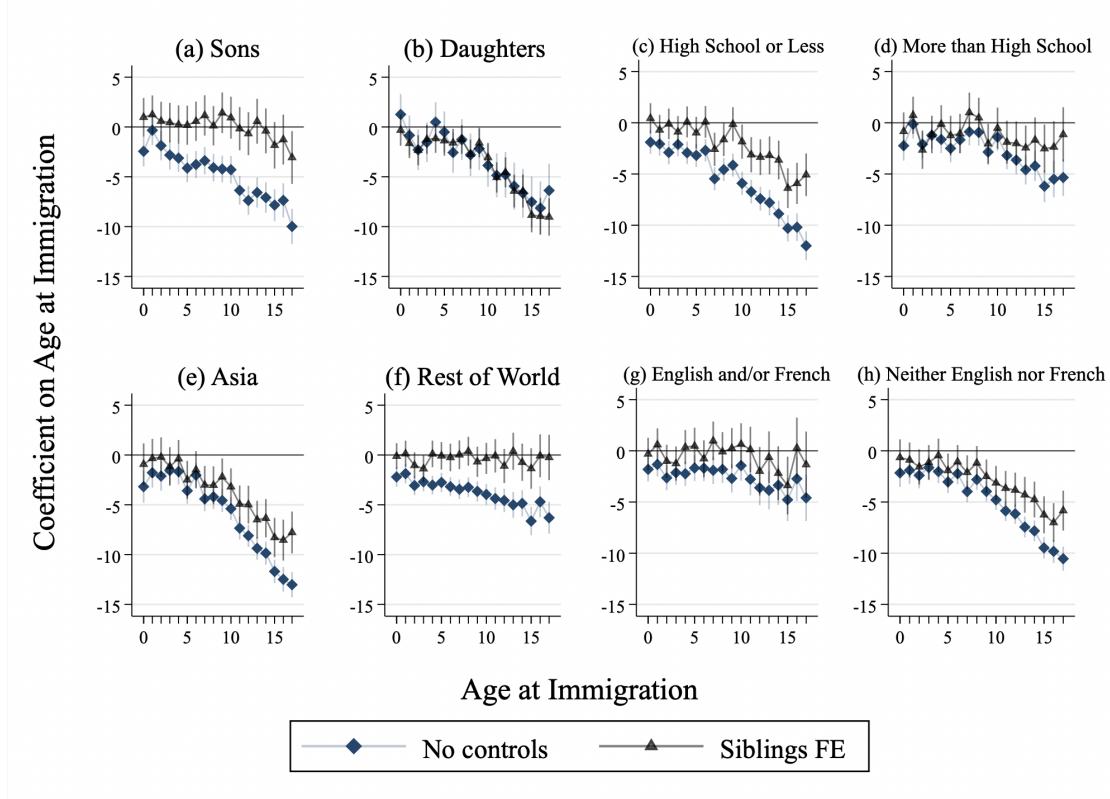
One point to note is that the siblings fixed effects model needs to be estimated on a sample that only includes children with identified siblings in our dataset. Appendix Figure B4 plots on the same graph the estimates of the model without controls (Equation 1) whether they come from the siblings sample (the blue diamonds, coming from Figure 2) or from the full sample (the yellow circles, coming from Figure B3). The two series are very close to one another. This tells us that using our siblings subsample does not introduce a bias when estimating age-at-immigration coefficients without controls, and provides support for our claim about the randomness of the unidentified siblings in the IID.

4.2 Heterogeneity

In this subsection, we estimate the impact of age at immigration for different subgroups.⁹ Figure 3 shows the age-at-immigration effects, both without controls and with siblings fixed effects, for those subgroups of our sample. We start with Subfigures (a) and (b), showing the difference between sons and daughters. We see that the unconditional effect of age at immigration (the blue diamonds) is similar for both groups. However, once we include siblings fixed effects (the gray triangles), age at immigration only influences the income of daughters who arrive after the age of 9. We also notice that adding siblings fixed does not close the gap with the second generation for daughters, which is a finding that calls for further investigation in a future study.

⁹We present in the Appendix additional figures using different cuts of the data not discussed here, including by birth cohort (B7), by immigrant category (economic, family or refugee, B8), and by type of family (two parents or monoparental, B9).

Figure 3: Heterogeneity in Age-at-Immigration Effects



Source: Authors' calculations based on the IID+.

Note: This figure plots the coefficients b_m for each age at immigration m from Equations 1 and 2 estimated separately for different subgroups (as identified by the header), with each series coming from a separate regression (either without controls or with siblings fixed effects). The sample includes all children of the 1.5 generation whose age at immigration is between 0 and 17, inclusively, as well as children from the second generation (the reference group). Child income is measured at ages 27 to 31. Parental income is measured when the child is 15 to 19 years old. The spikes represent 95% confidence intervals, based on robust standard errors.

We then consider parental schooling attainment in Subfigures (c) and (d). Two findings stand out. First, the association between the age at arrival and income is larger for children whose mothers have at most a high school diploma, with the unconditional effects more than twice as large by age 17 for children of low educated mothers (coef. = -12.0 , std. err. = 0.72) compared to those of more highly educated mothers (coef. = -5.3 , std. err. = 0.91). Second, once we control

for siblings fixed effects, the estimated b_m are not statistically different from 0 for children of more educated mothers, but they do remain negative and significant at ages at immigration of 10 and above for children whose mothers have at most a high school diploma.¹⁰

Combining child sex and maternal schooling attainment (shown in Appendix Figure B6), we observe that within a maternal education category, the unconditional effects are similar for sons and daughters. Once we include siblings fixed effects, we find that daughters of mothers with a lower education level are the only ones affected by age at immigration: by age 17, the estimated b_{17} for sons is -3.0 (std. err. = 1.93), while it is -9.63 for daughters (std. err. = 1.92).

Next we estimate the impact of age at immigration for children coming from Asia only, compared to the rest of the world (Subfigures (e) and (f)). We see that children whose parents came from Asia have a much steeper negative profile compared to children from the rest of the world, a finding that echoes Hermansen (2017). Once siblings fixed effects are accounted for, children from the rest of the world do not have lower average income ranks when arriving at older ages, while children coming from Asia continue to have significantly negative coefficients ($b_{17} = -7.8$, std. err. = 1.1). This pattern is only present for those coming from Asia: see Appendix Figure B5 for the estimated coefficients separately for each region of origin.

Finally, we investigate the heterogeneity by mother tongue. Language skills are an important factor for successful integration in a host society, especially for school-age children who might face instruction in a language they do not understand well (Bleakley & Chin, 2004, 2008, 2010; Guven & Islam, 2015). Language also bounds social networks: a better understanding of the local language can help navigate the systems of the host country, and information circulates more easily within a language group (Bertrand et al., 2000). Ideally, we would have information on proficiency in one of the official languages in Canada (English or French) upon arrival, but we only have either mother tongue or knowledge of official languages at the time of the Census (potentially many years after immigration). We

¹⁰The patterns are extremely similar for fathers' schooling attainment given the high correlation between maternal and paternal schooling. Those estimates can be provided by the authors upon request.

present the age-at-immigration effects estimated separately by whether the child's mother tongue is English and/or French (Subfigure (g)), or neither one of the official languages (Subfigure (h)). While we do see a negative unconditional effect of age at immigration for children whose mother tongue is English and/or French, the effects become indistinguishable from zero when adding siblings fixed effects. This is different from what we find for those whose mother tongue was neither English nor French: for them, arriving at age 17, for example, translated into a 5.8 percentile rank loss in adulthood income rank (std. err. = 0.99) relative to those born in Canada or who immigrated before 10 years old. Mastery of one of the official languages is therefore a critical step in the successful inclusion of children immigrating to Canada, with potential ripple effects to the next generation, as shown by Bleakley & Chin (2008) in the American context.

In summary, while age-at-immigration effects disappear for certain subgroups once we include siblings fixed effects, the baseline findings of Figure 2 remain for others. Even when siblings fixed effects are accounted for, daughters whose mothers have no more than a high school diploma, children coming from Asia, as well as those whose mother tongue was neither English nor French all have lower adulthood income if they immigrated after the age of 10, and the impact increases with each additional year spent outside of Canada.¹¹

5 Intergenerational Income Mobility

So far we have focused on the impact of age at immigration on adulthood income rank, either unconditionally or including siblings fixed effects. In this section, we turn our attention to intergenerational income mobility and the role of age at immigration.

Figure 4 presents a visualization of the relationship between child income rank

¹¹Exposure effects (γ) estimated using a piecewise linear function with a kink at age 10 for the different subsamples presented above can be found in Appendix Table C3. While our baseline $-\hat{\gamma}$ for arrival at age 10 and above in the model with full controls was -0.423 (std. err. = 0.077), we find larger (in absolute value) values for the groups identified above, though the differences do not appear to be statistically significant: -0.511 (std. err. = 0.158) for daughters, -0.549 (std. err. = 0.12) for those with low educated mothers, -0.707 (std. err. = 0.206) for daughters with low educated mothers, -0.713 (std. err. = 0.13) for those coming from Asia, and -0.517 (std. err. = 0.103) for those whose mother tongue was neither English nor French.

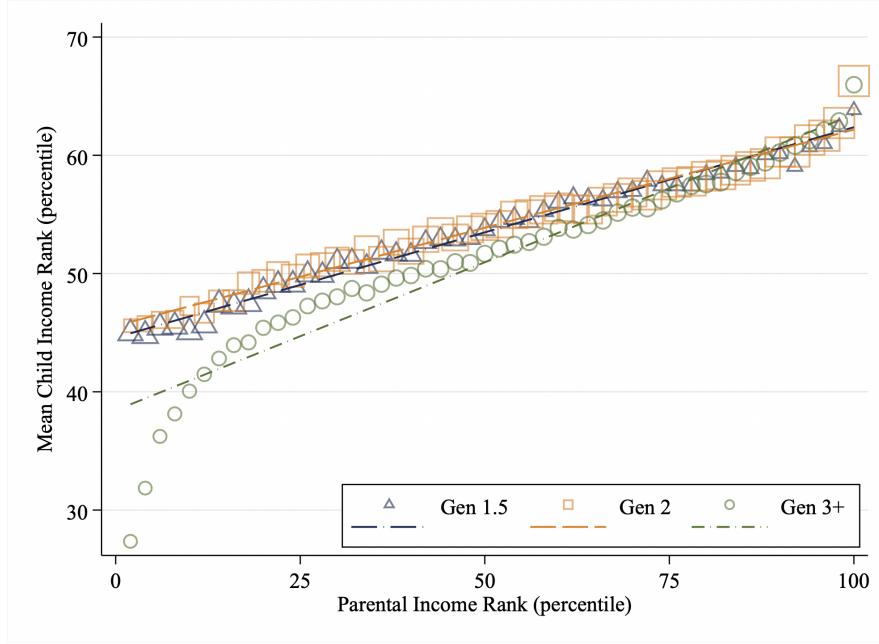
and parental income rank by generation status.¹² The figure is a binned scatter plot: each point represents a given parental income percentile rank (on the X axis), and plots the mean child income rank on the Y axis, with different series by generation status. The size of the markers are relative to the within-series share of children in each bin. For ease of interpretation, parental income percentiles are grouped into pairs, hence there are 50 points per series. Intergenerational income rank mobility estimates are given by the estimated slope coefficients of a model in which child income rank is explained by parental income rank (equivalent to a fitted line through the points on the binned scatter plot). A high rank-rank slope means that parental income rank is highly correlated with child income rank, and thus that intergenerational mobility is low.

Figure 4 shows that second-generation children have a greater intergenerational mobility than children of the 1.5 generation: their slopes are 0.166 (95% CI: [0.162; 0.170]) and 0.176 (95% CI: [0.172; 0.180]), respectively. Nonetheless, they display greater intergenerational mobility than nonimmigrant children, whose slope is estimated at 0.250 (95% CI: [0.248; 0.252]). Tests of equality of slopes across generation groups reject equality for all comparisons (p -values < 0.001). These estimates are in line with previous findings by Connolly et al. (2022) showing that, for every birth year in the IID, children of immigrant mothers had on average a higher income mobility relative to nonimmigrant children. But this figure further deepens our understanding of mobility between different groups of children.

First, it shows that the average ranks of children from the 1.5 generation are very similar to those of the second generation, but that the distribution of parental ranks within each generation differs. Parents of the 1.5 generation are more concentrated in the lower income ranks, while those of the second generation can be found more in the upper ranks. This difference in the distribution of parental ranks explains why the estimated slopes differ: for the 1.5 generation, lower parental ranks are overrepresented and help tilt the fitted regression line to have a slightly steeper slope.

¹²Appendix Figure B10 focuses on the 1.5 generation, splitting the sample by bands of age at immigration. The figure shows that the relationship between parental income rank and child income rank becomes stronger with age at immigration, increasing from 0.154 when children are not yet in primary school (ages zero to five) to 0.180 when they are six to 11 years old and 0.187 once in their teenage years.

Figure 4: Intergenerational Income Rank Mobility by Generation Status



Source: Authors' calculations based on the IID+.

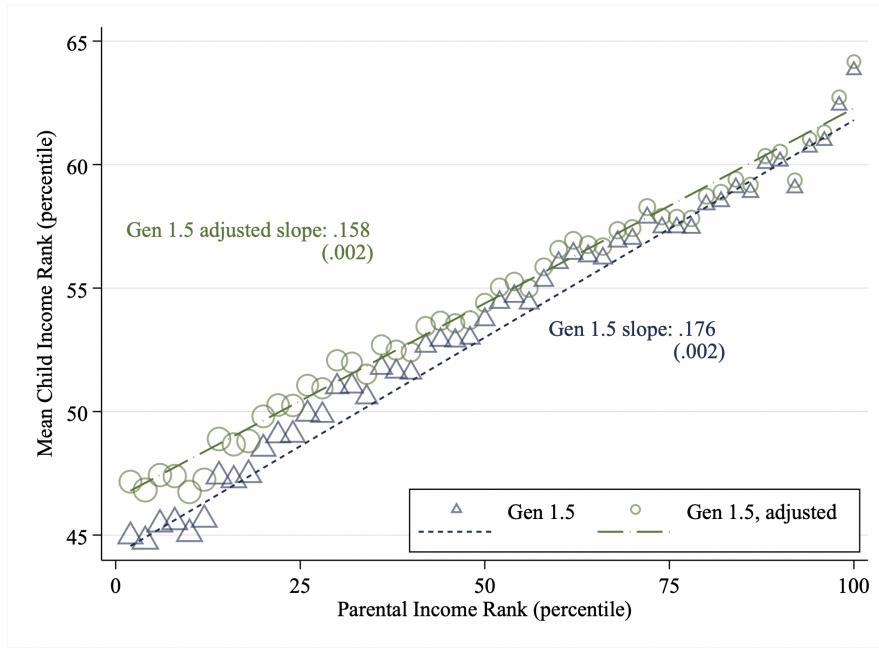
Note: Child income is measured at ages 27 to 31. Parental income is measured when the child is 15 to 19 years old. Ranks are assigned within child birth year irrespective of immigration status. Parental income ranks are grouped in pairs; each point represents two percentiles. Size of markers is relative to the number of children in each bin within a given series.

Second, when both groups of immigrant children are compared to nonimmigrant children, we observe that immigrant children have much higher mobility rates relative to nonimmigrant children. While the difference in average child income ranks is particularly striking below the 25th percentile, it is also visible roughly up until the 75th percentile, at which point the three series practically overlap. This suggests that children of the 1.5 generation from low-income families have a harder time integrating than second-generation immigrant children from comparable economic backgrounds, but immigrant children from the low end of the income distribution outdo by far nonimmigrants from similar backgrounds. Both findings merit further attention in subsequent studies.

Now that we have estimated the effects of age at immigration on adulthood income for the 1.5 generation, we can perform an accounting exercise to gauge the

contribution of age at immigration to intergenerational mobility. As explained in Section 3, we adjust the observed income ranks of 1.5-generation children in our entire sample ($N = 334,730$) by subtracting the estimated coefficient that corresponds to their age at arrival, in effect computing a counterfactual of their income rank were they born in the country from immigrant parents.

Figure 5: Intergenerational Income Rank Mobility of the 1.5 Generation, Adjusting for Age at Immigration



Source: Authors' calculations based on the IID+.

Note: Child income is measured at ages 27 to 31. Parental income is measured when the child is 15 to 19 years old. Ranks are assigned within child birth year irrespective of immigration status. In the adjusted series, child income ranks are adjusted by subtracting the coefficient estimated in the baseline siblings fixed effects model (shown in Figure 2) corresponding to their age at immigration. Parental income ranks are grouped in pairs; each point represents two percentiles. Size of markers is relative to the number of children in each bin within a given series.

In Figure 5, we plot the average child income rank by parental income for the observed 1.5 generation (the blue triangles), identical to that from Figure 4, and the adjusted 1.5 generation (the green circles). We can see that the adjustment raises the income ranks, especially in the bottom half of the parental income distribution: the circles lie above the triangles. The resulting adjusted slope is

0.158 (95% CI: [0.154; 0.162]), significantly lower than the non-adjusted one at 0.176 (95% CI: [0.172; 0.180]). The difference is 0.018, in favor of the adjusted series, meaning removing age-at-immigration effects results in increased intergenerational mobility. In this calculation, age at immigration accounts for 10.2% of the intergenerational mobility estimate of the 1.5 generation in Canada.¹³ ¹⁴

6 Discussion

We start this section by offering a number of policy recommendations that have the potential to enhance the economic integration of immigrant children. First, since age at arrival matters for later economic success, immigration requests of families with children should be treated as quickly as possible to ensure that selected families arrive when their children are younger. Another promising approach may be to implement incentives for families to immigrate when their children are younger.

Next, we should increase our support to children arriving during adolescence since they experience the largest negative impacts. Governments could invest in programs that target the integration of children who arrive during the high school years, especially those that belong to one of the groups that we identified as experiencing stronger age-at-immigration effects, such as immigrants from Asia or whose mother tongue is neither French nor English. Language barriers appear to be a limitation to economic success that could be alleviated with better language programs for both parents and children.

Finally, the economic integration of parents into society has repercussions on their children. The degree of influence exerted by parental integration on children's outcomes is partly linked to the strength of intergenerational mobility. When there

¹³Note that this accounting could vary by region of origin, with a larger share (almost 20%) being accounted by age-at-immigration effects for those coming from Asia, and a much smaller share for those from other regions of the world. This is a direct result of the large and negative coefficients for children from Asia and the near-zero effects for others seen in Subfigures 3(e) and 3(f) and in Appendix Figure B5. When the effects are close to zero, the adjusted ranks are almost identical to the observed ranks, but when the coefficients are larger, the adjustment results in a larger explained share.

¹⁴Appendix Table C4 presents additional results illustrating the heterogeneity in intergenerational mobility by region of origin.

is a strong correlation between the income rank of parents and their children, the ripple effects of parental integration becomes more pronounced. Consequently, improving parental outcomes, particularly for those facing low intergenerational mobility, may enhance their children's future. By focusing on boosting the outcomes of parents, particularly immigrants from Asia who arrive with older children, the overall well-being and prospects of their children can also be positively affected.

The present study encountered a number of limitations, primarily due to constraints in the available data. First, the IID-Census linkage does not recover the immigration status of every IID individual, resulting in a smaller analysis sample. We do use five different Census waves to reduce the sample loss, and the attrition should not be selective, but it still results in a final sample that is around 60% smaller than the original IID. Linking the IID with more complete data on immigrants, such as administrative landing files, would help identify the generational status for a larger share of the IID.

Additionally, our main identification strategy relies on siblings fixed effects, but as mentioned earlier, we can only identify siblings in the administrative tax data when their birth year is part of the IID target years. The result is again a loss of sample, but while we recognize that we are losing some kids who have siblings, we argue that the selection process should be reasonably random, allowing us to maintain a representative—albeit smaller—sample.

A more subtle point relates to the identification of our main explanatory variable, age at immigration. As previously stated, the information available in the Census pertains to the age at which an individual became a landed immigrant (permanent resident), which may not correspond to the actual year of arrival in Canada. As a result, if a child entered Canada as a refugee and later obtained their immigrant status, their age at arrival might be overestimated, which could bias our estimates. However, according to Corak (2012), it is not clear how individuals answer the Census question: respondents could report the year of their arrival in Canada as opposed to the year in which they obtained their landed immigrant status. The present data set does not allow for the validation of this hypothesis and again, linking the fiscal data with landing records might help bet-

ter identify age at immigration. We also note that illegal immigration to Canada was—and to some extent still is, especially compared to the United States—very low: one estimate put the figure at around 1% of the total number of immigrants (Robinson, 1984).

7 Conclusion

The successful integration of immigrants is a crucial concern, particularly in Canada, where immigrants make up a large proportion of the population—more so than in many other countries. Evaluating successful integration can be done through various measures, such as language acquisition, educational outcomes, or adulthood income. This study has focused on adulthood income and intergenerational income mobility, a measure that captures an individual’s ability to move up or down the income distribution relative to their parents. While a society may deem it important that children from disadvantaged and privileged backgrounds are provided with equal opportunities for success regardless of their immigrant status, we directed our attention to the 1.5 generation. Those are immigrants who arrived in the country, along with their parents, before the age of 18. Our findings highlight that they face greater challenges, particularly the older they are at the time of immigration.

The main objective of this paper was to evaluate the contribution of age at immigration to the intergenerational income mobility of children from the 1.5 generation. We achieved this by exploiting a novel data linkage: intergenerational tax files linked to Census data.

Our results show that children who immigrate during their childhood have lower income ranks at ages 27 to 31 relative to children of immigrants born in the country, but a similar average income relative to nonimmigrant children. However, children who arrive to Canada after age 10 face more challenges. More specifically, the average unconditional income in adulthood of someone who arrived at the age of 17 is 8.8 percentile ranks lower than that of a second-generation immigrant, a difference of about CAD\$6,200 in 2025 dollars (US\$4,500).

To identify the causal effect of age at immigration, we have leveraged variation

within families using a siblings fixed effects model, comparing children who immigrated to Canada in the same year, but at different ages. We find that the negative influence of age at immigration on income is dampened, but by no means erased. Furthermore, coherent with previous findings, we show that the exposure to host country effects are not linear: exposure effects are relatively flat until age 10, but then become larger in absolute value. Each year of delayed arrival to Canada after age 10 is estimated to cost close to half a percentile rank of income. Finally, we also find differences across subgroups. The interaction between sex and maternal schooling level reveals that daughters of mothers with a lower education level are particularly affected by age at immigration. Additionally, children from Asia and children whose mother tongue is neither English nor French also suffer more from delayed immigration.

Our results on the intergenerational income mobility of Canadians by immigrant generation status using a rank-rank model show that second-generation immigrants do particularly well in Canada. Their rank-rank correlation is 0.166 (95% CI: [0.162; 0.170]), while the correlation for children from the 1.5 generation is 0.176 (95% CI: [0.172; 0.180]) and that of nonimmigrants is 0.250 (95% CI: [0.248; 0.252]). Nonimmigrants have the highest rank-rank correlation, indicating the lowest level of mobility. While 0.25 indicates more mobility than what is typically seen in the United States (Chetty et al., 2014; Connolly et al., 2019), the low levels of opportunities of nonimmigrant children from the bottom part of the parental income distribution is something that deserves public attention and further research.

Finally, we reached our main conclusion: we estimated the contribution of age at immigration to intergenerational income mobility, or said differently, we evaluated how much would income mobility improve if the 1.5 generation were born in Canada. Using an adjustment to income ranks, in which we subtracted the estimated age-at-immigration coefficient from the observed income rank, we performed a counterfactual exercise that showed that the intergenerational correlation between parental and child income would be lower by 0.018 or 10.2%.

Based on these findings, we outlined a number of policy recommendations including targeted interventions to facilitate language acquisition and integration

into the schooling system for immigrants who arrive during adolescence, as well as incentives for families to immigrate while their children are young or prioritizing families with younger children. Better support for children whose parents have low income levels, irrespective of their immigration status, should also be a priority. Finally, future research could focus on improving our understanding of the dynamics behind the very low incomes of nonimmigrant children from the bottom of the parental income distribution. Should more data linkages become available, we may also be able to deepen our comprehension of the factors that could improve the economic integration of children of immigrants.

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A Data Appendix

A.1 IID and Census

Our analysis is based on data from Statistics Canada’s Intergenerational Income Database (IID) linked to de-identified Census data (Statistics Canada, 2016b,a). This subsection contains additional information on these data. Each year, Canadians file their taxes with the Canada Revenue Agency using the T1 form. The Canada Revenue Agency then share these data with Statistics Canada, who compile them in the T1 Family File (T1FF). In the T1FF, tax filers from the same family are identified through common links such as spousal social insurance number, surname, and address. The IID identifies parents of children born between 1963 and 1985 in the T1FF, excluding birth years 1971, 1976, and 1981 due to the definition of the target population. For more information on the child-parents linkage in the IID, see Corak & Heisz (1999) and Corak (2020).

The IID comes with weights that we are required to use to compute counts and all statistics, as well as to estimate models. The weights inflate the total counts, but by a small factor, given that the IID is an administrative database covering all of the children present in Canada during the target years (at ages 16 to 19). For example, the ratio of total weighted count to unweighted count is 1.21 for the cohort born from 1982 to 1985 (Connolly et al., 2019).

While administrative tax data contain precise information on income, sociodemographic information is limited. Statistics Canada’s Social Data Linkage Environment allows linkages between existing data using keys generated from record IDs and stored in a key registry. In the IID+, linkages are made with the 1996, 2001, 2006 and 2016 Census and the 2011 National Household Survey, which was a temporary replacement for the long-form Census.

A.2 Variables definitions

In this subsection, we define our key variables and some of the alternatives used in the robustness checks.

1. *Child Income Rank.* Our income measure is based on Canada Revenue Agency’s definition of total income and includes earnings, interest and investment income, self-employment net income, taxable capital gains, losses and dividends, and benefits, which we express in 2017 dollars (using the Consumer Price Index from Statistics Canada, 2021). We average child income in adulthood over five

years from ages 27 to 31 to reduce the errors-in-variables bias that occurs when using a single year of data (Solon, 1992). Income rank is determined using the distribution of total income for all children within the same birth year, irrespective of Census link status, restricted to total incomes above \$500. We also compute child income rank using total income from ages 30 to 34 as a robustness check.

2. Parental Income Rank. Parental income is the sum of both mother's and father's total income. It is also averaged over a five-year period, when the child is 15 to 19 years old, in order to reflect parental economic resources during adolescence. Rank is relative to all other parents with a child born during the same birth year, and once again computed only for those with a total income over \$500. We also compute parental income rank using total income when the parents are aged 40 to 49 as a robustness check.

3. Age at immigration. The age at immigration variable available in Census data is derived from the date of birth and year of immigration (Statistics Canada, 2018). The Census does not ask directly about age at immigration. For example, in the 2016 Census, Question 3 was “What are this person’s date of birth and age?” and Question 15 was “In what year did this person first become a landed immigrant?” Question 15 was only asked to people who answered “Yes” to Question 14, which reads “Is this person now, or has this person ever been, a landed immigrant? A “landed immigrant” (permanent resident) is a person who has been granted the right to live in Canada permanently by immigration authorities.” (Statistics Canada, 2015). We note, as Corak (2012) did, that this could lead to an overestimation of age at immigration, particularly among refugees, for whom the year of arrival in Canada and the year when they received landed immigrant status may differ.

4. Generation status. Children are considered 1.5 generation immigrants if they are born outside of Canada, have at least one immigrant parent, and arrived before the age of 18. If immigration happened during adulthood, they are considered first-generation immigrants and are thus excluded from our final sample. As a robustness check, we also use a definition requiring both parents to be immigrants to classify a child in the 1.5 generation. As shown in Table 1, among the close to three million children in our analysis sample, 334,730 belong to the 1.5 generation, 553,090 are second-generation, and 1,959,000 are nonimmigrant children. The distribution of children by generation status is relatively stable across birth cohorts: depending on the child’s birth cohort, the 1.5 generation represents between 10.9% and 12.5% of the children whose generation status could be deter-

mined, while the second generation represents 17.1% to 20.8% of those children. See Appendix Figure B1 for the distribution of the generation status by birth cohort.

5. *Region of Origin.* From the Census, we obtain information on country of birth. The child's country of birth is used as the country of origin for children of the 1.5 generation and parental country of birth is used for second-generation children. If the mother's and father's country of birth are not identical, the mother's country of birth is kept. Following Aydemir et al. (2009, 2013), countries are then grouped into five regions of origin: (1) North America and Northern and Western Europe, (2) Caribbean, Central and South America, and Oceania, (3) Southern and Eastern Europe, (4) Africa, and (5) Asia. Children with missing information are excluded from the analysis sample.

6. *Region of Destination.* We use IID data to determine the child's province of residence during adolescence.¹⁵ Provinces are grouped into five regions of residence: (1) British Columbia, (2) Prairies (Alberta and Manitoba), (3) Ontario, (4) Quebec, and (5) Atlantic provinces (New Brunswick, Nova Scotia, Prince Edward Island and Newfoundland and Labrador). Individuals residing in Canada's territories are excluded due to low observation counts.

7. *Birth Cohort.* IID children are divided into five birth cohorts according to their birth year: (1) 1963 to 1966, (2) 1967 to 1970, (3) 1972 to 1975, (4) 1977 to 1980, and (5) 1982 to 1985 (see 1).

A.3 Descriptive statistics by region of origin

The averages shown in Figure 1 potentially hide important differences by country of origin. Figure A1 reports the mean income rank by region of origin, generation status, and age-at-immigration categories for parents (Subfigure A1a) and children (Subfigure A1b).¹⁶ Most immigrant children have higher average income ranks than nonimmigrants (the dashed line). The exceptions are children from the Caribbean, Central and South America, and Oceania, whose average income rank is below that of nonimmigrant children regardless of generation status and age at immigration, as well as children who arrived between 12 and 17 years old from Asia. Subfigure A1b also shows that children's mean rank decreases with age at immigration, in line with Figure 1.

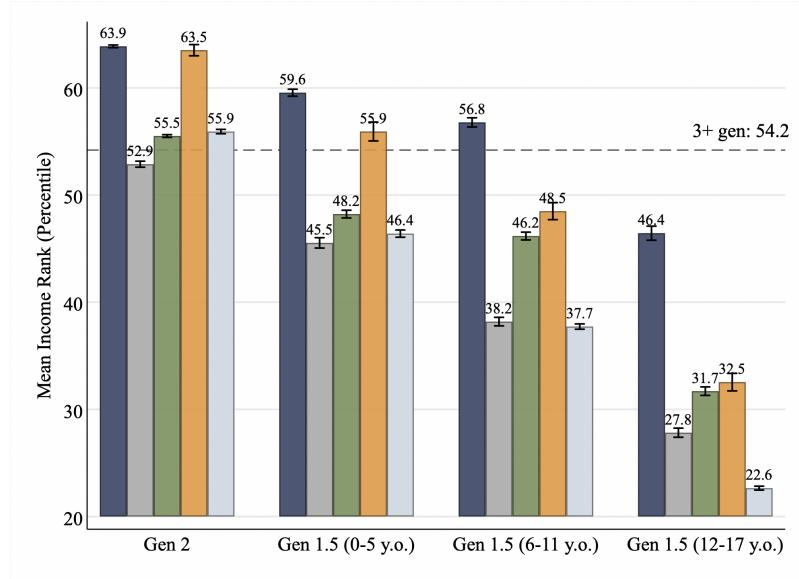
¹⁵Technically, at the time of the parents-child linkage in the IID, so at age 16 for the majority of children.

¹⁶See Appendix Table C1 for the distribution of region of origin by generation status.

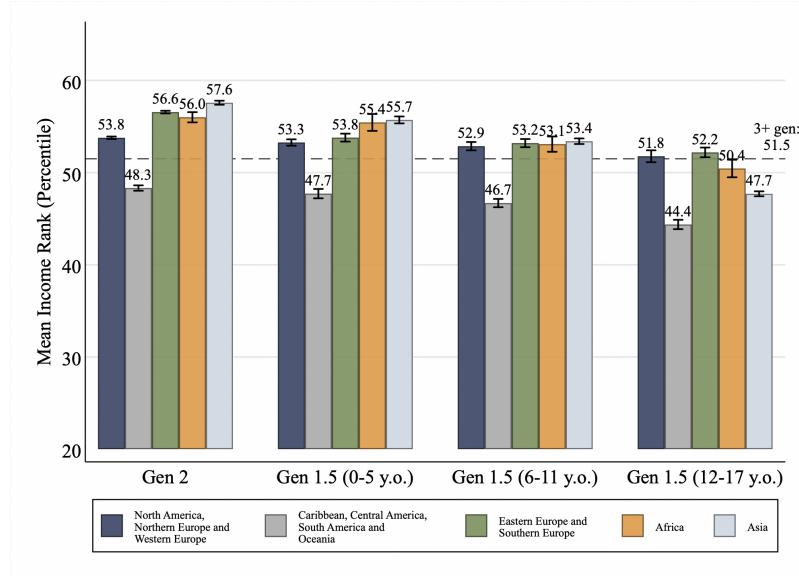
Subfigure A1a shows that second-generation parents have higher incomes than nonimmigrant parents on average, except for those from the Caribbean, Central and Southern America, and Oceania. In contrast, parents of the 1.5 generation generally do not surpass nonimmigrant parents in terms of average income ranks. There are two exceptions: parents from North America and Northern and Western Europe arriving to Canada with children under the age of 12, and those from Africa when arriving with children under six. Subfigure A1a also shows that parental average rank decreases with children's age at immigration for all regions of origin, going as low as 22.6 for Asian parents with children arriving during adolescence.

Coming back to Subfigure A1b and the relationship between age at immigration and adulthood income, we observe that the strongest association appears to be for children from Asia, with an average rank spanning from 55.7 for those who immigrated between the ages of zero and five to 47.7 for those that arrived at later ages, a 8.0 percentile income rank gap. A similar but steeper income rank gap is observed at the parental level with the gap ranging from 13.1 percentile rank for the North America and Northern and Western Europe region to 23.8 for Asia. This suggests that age at immigration has a greater cost for certain groups (Schaafsma & Sweetman, 2001; Pendakur & Pendakur, 2016). Several factors can contribute to the lower income ranks experienced by these parents in the early years after their arrival, including language barriers, challenges in recognizing foreign credentials, and potential discrimination. Additionally, the income ranks can vary depending on the type of immigrant (such as economic immigrants, family-sponsored immigrants, or refugees), which in turn can vary with year at arrival and origin. In the late 1970s, a significant influx of Asian refugees arrived following the Vietnam War, which may explain the lower income ranks of parents from Asia in our data. Adnan et al. (2023) compare the intergenerational income mobility of refugees and non-refugees for immigrant children who arrived in Canada after 1980, and find that refugees have lower income ranks (conditional on parental income), especially at the bottom of the parental income distribution. The authors also offer an analysis of the gap by refugee status and the selection effect of pre-immigration variables.

Figure A1: Mean Income Rank by Region of Origin, Generational Status and Age at Immigration



(a) Parents



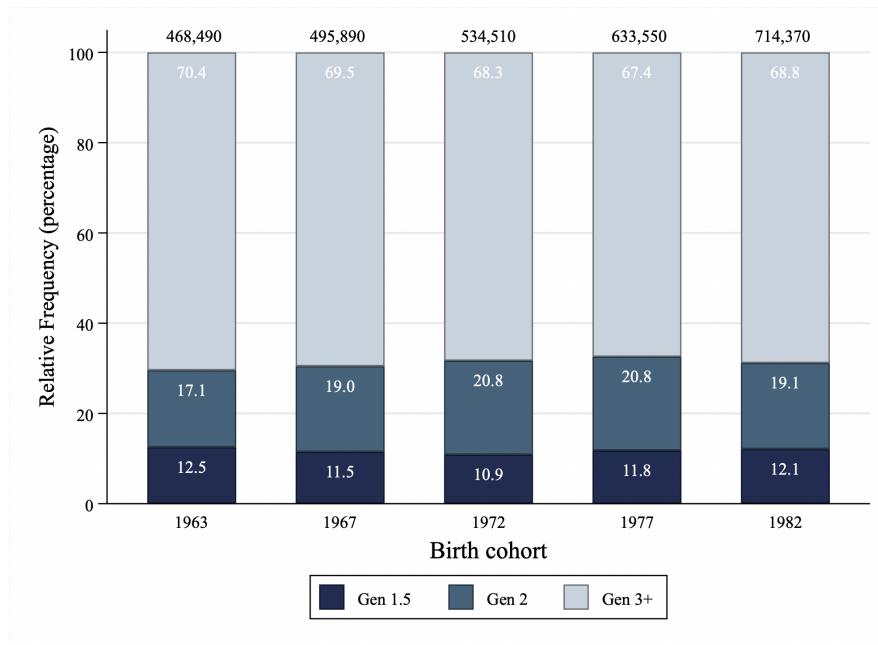
(b) Children

Source: Authors' calculations based on the IID+.

Note: The horizontal lines labeled "3+ gen" represent the mean income rank for nonimmigrant parents (54.2, Subfigure a) and children (51.5, Subfigure b). Mean percentile ranks are shown above each bar. The spikes represent 95% confidence intervals.

B Appendix Figures

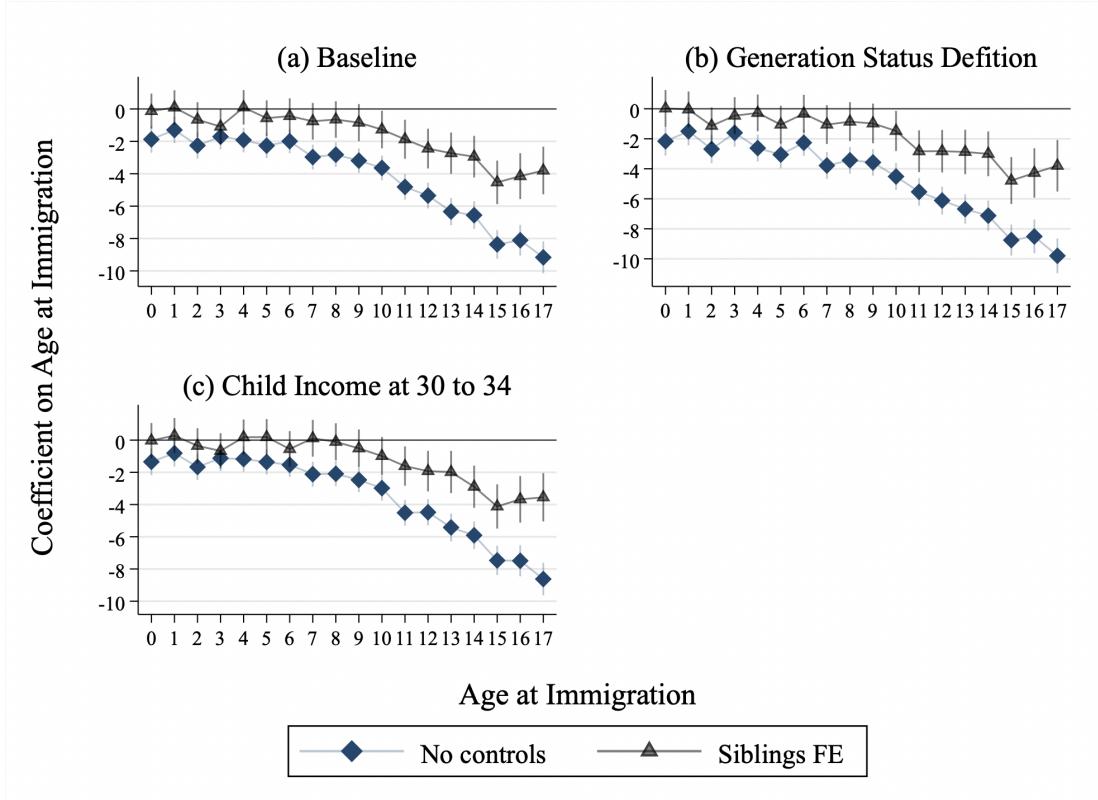
Figure B1: Distribution of Generation Status by Birth Cohort



Source: Authors' calculations based on the IID+.

Note: Percentages by category are in white on the bars; total (weighted) number of children per cohort are above the bars.

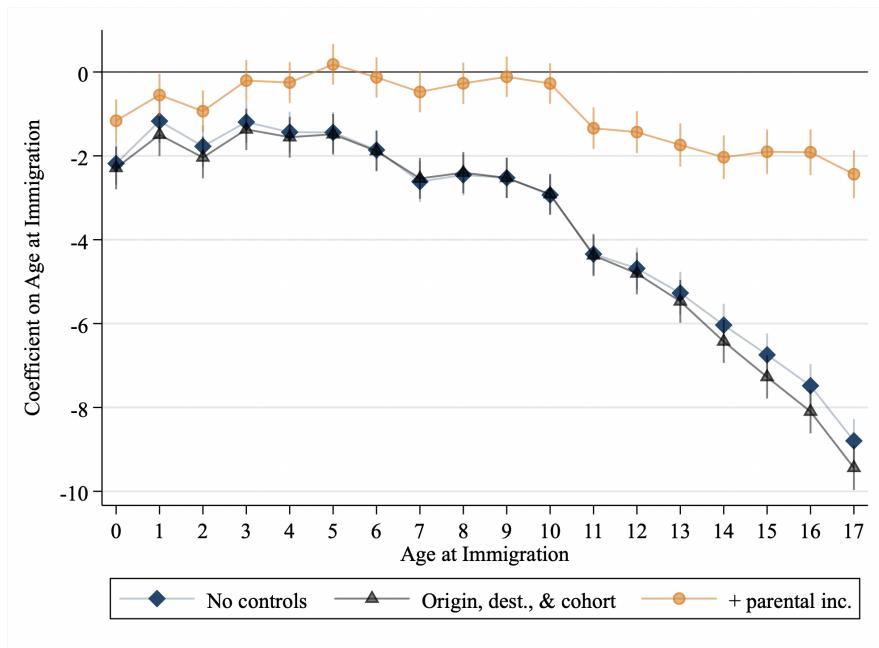
Figure B2: Age-at-Immigration Effects by Specification and Sample Definition



Source: Authors' calculations based on the IID+.

Note: As in Figure 2, this figure plots the estimated coefficients b_m for each age at immigration m from Equations 1 and 2, with each series coming from a separate regression (either without controls or with siblings fixed effects). Subfigure (a) reports the baseline from Figure 2. In Subfigure (b), immigrant children are defined as having two immigrant parents. Subfigure (c) shows estimates when child income is measured at ages 30 to 34. The spikes represent 95% confidence intervals, based on robust standard errors.

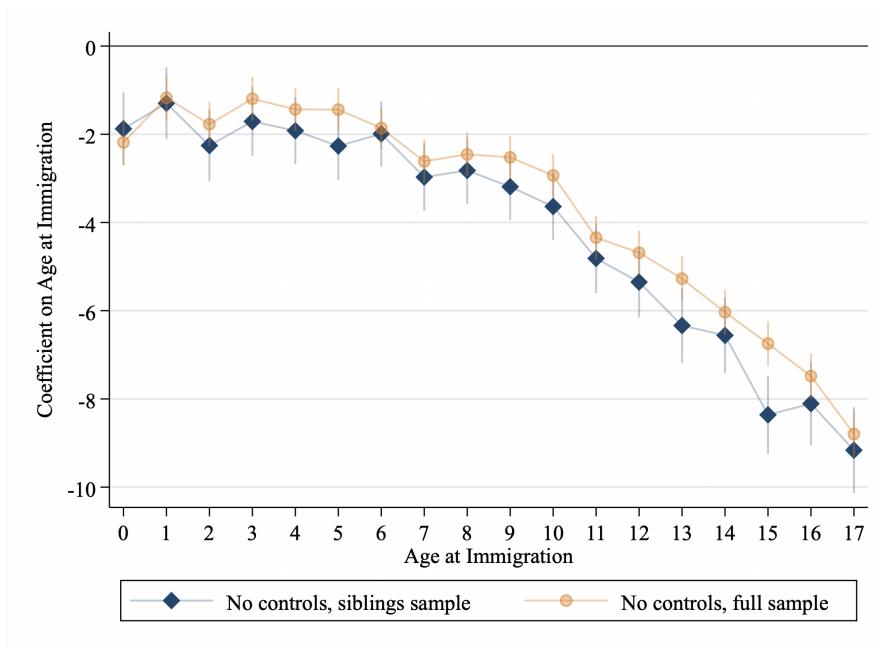
Figure B3: Age-at-Immigration Effects on Income Ranks in Adulthood without Siblings Fixed Effects



Source: Authors' calculations based on the IID+.

Note: This figure plots the coefficients b_m for each age at immigration m estimated using the full sample (not restricting to having siblings), with each series coming from a separate regression with progressively more controls. The series with the blue diamonds has no controls and corresponds to Equation 1. The series represented by the gray triangles adds region of origin, region of destination, and birth cohort fixed effects to the baseline model, and the one with the yellow circles adds parental income decile. The spikes represent 95% confidence intervals, based on robust standard errors.

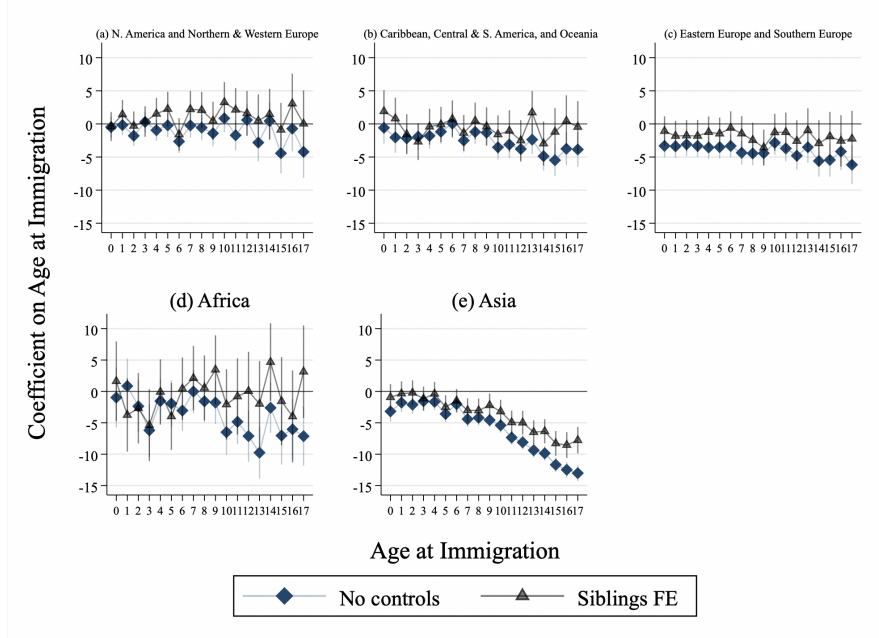
Figure B4: Age-at-Immigration Effects on Income Ranks in Adulthood, Siblings Sample and Full Sample



Source: Authors' calculations based on the IID+.

Note: This figure plots the coefficients b_m for each age at immigration m from Equation 1 (with no other controls). The full sample (the yellow circles) includes all children of the 1.5 generation whose age at immigration is between 0 and 17, inclusively, as well as children from the second generation (the reference group). The siblings sample (the blue diamonds) corresponds to the one used in the siblings fixed effects estimations. Child income ranks are measured at ages 27 to 31. The spikes represent 95% confidence intervals, based on robust standard errors.

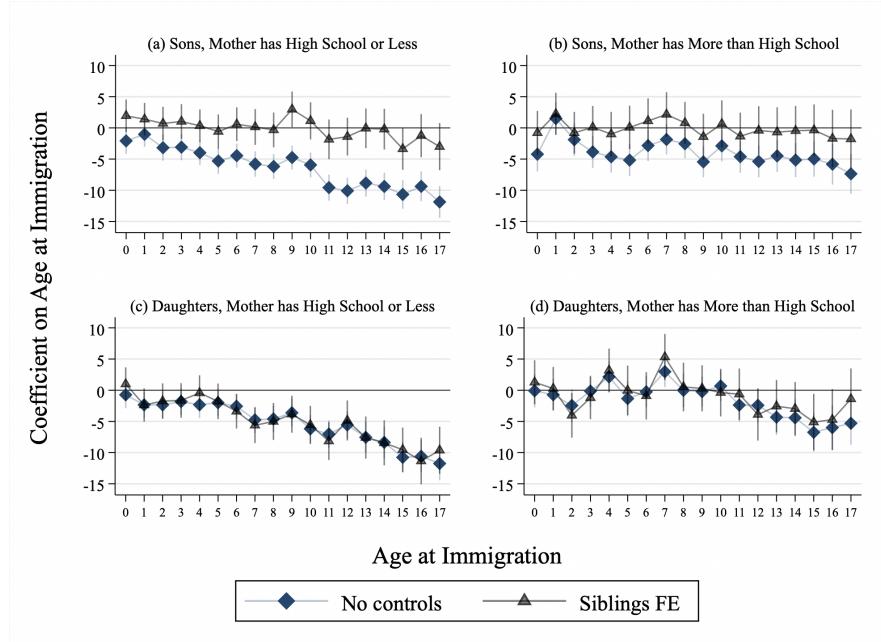
Figure B5: Age-at-Immigration Effects, by Region of Origin



Source: Authors' calculations based on the IID+.

Note: This figure plots the coefficients b_m for each age at immigration m from Equations 1 and 2 estimated separately by region of origin, with each series coming from a separate regression with progressively more controls. Regions are: North America and Northern and Western Europe (Subfigure a); Caribbean, Central and South America, and Oceania (Subfigure b); Eastern Europe and Southern Europe (Subfigure c); Africa (Subfigure d); and Asia (Subfigure e). The sample includes all children of the 1.5 generation whose age at immigration is between 0 and 17, inclusively, as well as children from the second generation (the reference group). Child income ranks are measured at ages 27 to 31. Parental income ranks are measured when the child is 15 to 19 years old. The spikes represent 95% confidence intervals, based on robust standard errors.

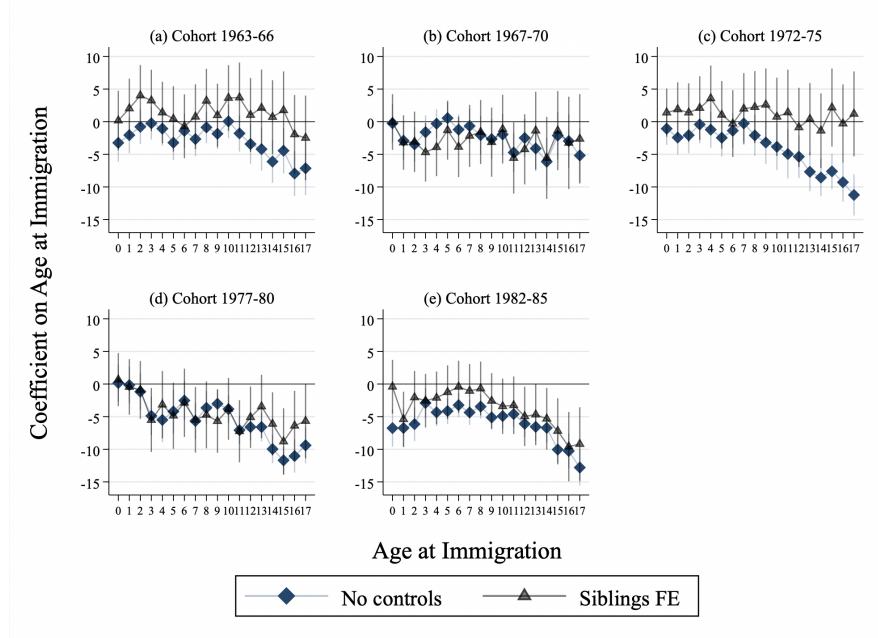
Figure B6: Age-at-Immigration Effects, by Child Sex and Mother's Schooling Attainment



Source: Authors' calculations based on the IID+.

Note: This figure plots the coefficients b_m for each age at immigration m from Equations 1 and 2 estimated separately by child sex and by mother's schooling attainment, with each series coming from a separate regression (either without controls or with siblings fixed effects). The sample includes all children of the 1.5 generation whose age at immigration is between 0 and 17, inclusively, as well as children from the second generation (the reference group). Child income is measured at ages 27 to 31. Parental income is measured when the child is 15 to 19 years old. The spikes represent 95% confidence intervals, based on robust standard errors.

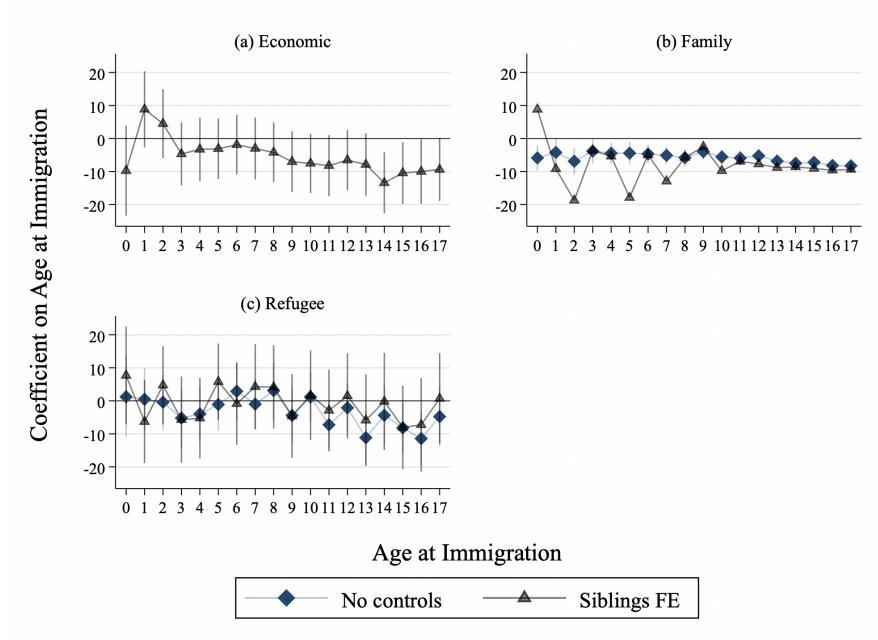
Figure B7: Age-at-Immigration Effects, by Birth Cohort



Source: Authors' calculations based on the IID+.

Note: This figure plots the coefficients b_m for each age at immigration m from Equations 1 and 2 estimated separately by child birth cohort, with each series coming from a separate regression (either without controls or with siblings fixed effects). The sample includes all children of the 1.5 generation whose age at immigration is between 0 and 17, inclusively, as well as children from the second generation (the reference group). Child income is measured at ages 27 to 31. Parental income is measured when the child is 15 to 19 years old. The spikes represent 95% confidence intervals, based on robust standard errors.

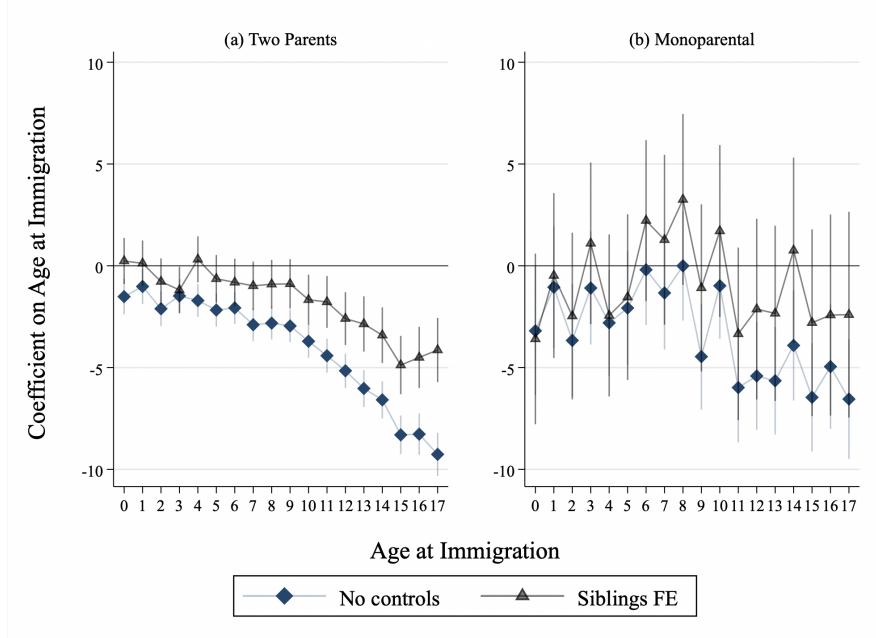
Figure B8: Age-at-Immigration Effects, by Immigrant Category



Source: Authors' calculations based on the IID+.

Note: This figure plots the coefficients b_m for each age at immigration m from Equations 1 and 2 estimated separately by the parents' immigrant category, with each series coming from a separate regression (either without controls or with siblings fixed effects). The sample includes all children of the 1.5 generation whose age at immigration is between 0 and 17, inclusively, as well as children from the second generation (the reference group). Child income is measured at ages 27 to 31. Parental income is measured when the child is 15 to 19 years old. The spikes represent 95% confidence intervals, based on robust standard errors.

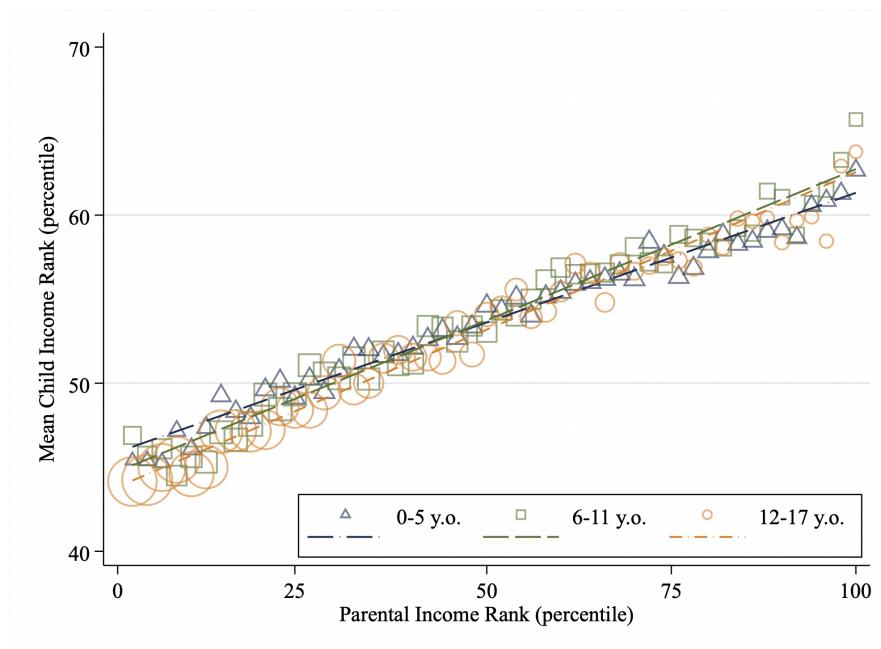
Figure B9: Age-at-Immigration Effects, by Type of Family



Source: Authors' calculations based on the IID+.

Note: This figure plots the coefficients b_m for each age at immigration m from Equations 1 and 2 estimated separately by type of family (two parents or monoparental), with each series coming from a separate regression (either without controls or with siblings fixed effects). The sample includes all children of the 1.5 generation whose age at immigration is between 0 and 17, inclusively, as well as children from the second generation (the reference group). Child income is measured at ages 27 to 31. Parental income is measured when the child is 15 to 19 years old. The spikes represent 95% confidence intervals, based on robust standard errors.

Figure B10: Intergenerational Income Rank Mobility of the 1.5 Generation by Age at Immigration



Source: Authors' calculations based on the IID+.

Note: Child income is measured at ages 27 to 31. Parental income is measured when the child is 15 to 19 years old. Ranks are assigned within child birth year irrespective of immigration status. Parental income ranks are grouped in pairs; each point represents two percentiles. Size of markers is relative to the number of children in each bin within a given series.

C Appendix Tables

Table C1: Relative Distribution by Region of Origin

Region of origin	Gen. 2		Gen. 1.5	
	0-17 y.o.	0-5 y.o.	6-11 y.o.	12-17 y.o.
North America, Northern and Western Europe	40.3	20.7	33.4	18.6
Caribbean, Central and South America, and Oceania	8.4	14.8	14.3	16.6
Eastern and Southern Europe	33.1	16.8	18.8	17.3
Africa	2.5	5.2	5.0	5.6
Asia	15.7	42.6	28.5	41.9
<i>N</i>	553,090	334,730	109,450	116,190
				109,090

Source: Authors' calculations based on the IID+.

Note: This table shows the relative distribution of immigrant children by region of origin, generation status, and age at immigration. Numbers are percentages, with each column summing to 100. The bottom row gives the number of (weighted) observations.

Table C2: Descriptive Statistics by Age at Immigration, Siblings Fixed Effects Sample

Age	Relative frequency	Average family size	Child sex		Maternal schooling attainment		Paternal schooling attainment		Region of origin	
			Female	Male	HS or less	Above HS	HS or less	Above HS	Asia	Rest of world
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	
0	5.3	2.2	49.6	50.4	59.2	40.8	44.1	55.9	28.3	71.7
1	5.3	2.2	50.7	49.3	58.6	41.4	44.4	55.6	31.7	68.3
2	5.6	2.2	50.5	49.5	58.2	41.8	42.1	57.9	31.6	68.4
3	5.8	2.3	50.8	49.2	56.7	43.3	42.3	57.7	33.5	66.5
4	6.0	2.3	50.0	50.0	57.2	42.8	43.1	56.9	36.4	63.6
5	6.1	2.3	48.8	51.2	59.2	40.8	44.9	55.1	38.1	61.9
6	6.4	2.3	49.6	50.4	58.6	41.4	44.6	55.4	40.5	59.5
7	6.2	2.3	50.3	49.7	57.9	42.1	44.5	55.5	41.2	58.8
8	6.2	2.3	49.9	50.1	58.2	41.8	45.4	54.6	44.0	56.0
9	6.4	2.3	48.5	51.5	55.7	44.3	44.4	55.6	45.9	54.1
10	6.3	2.3	50.1	49.9	56.9	43.1	44.1	55.9	48.9	51.1
11	5.9	2.3	48.9	51.1	58.0	42.0	44.9	55.1	52.6	47.4
12	5.7	2.3	48.6	51.4	56.9	43.1	43.4	56.6	54.2	45.8
13	5.2	2.3	49.2	50.8	56.2	43.8	45.8	54.2	58.1	41.9
14	5.1	2.3	48.2	51.8	55.4	44.6	44.0	56.0	58.9	41.1
15	4.7	2.3	48.8	51.2	57.4	42.6	45.2	54.8	61.2	38.8
16	4.0	2.3	48.9	51.1	60.0	40.0	47.4	52.6	61.8	38.2
17	3.6	2.3	50.1	49.9	60.1	39.9	45.5	54.5	63.7	36.3

Source: Authors' calculations based on the IID+.

Note: Sample corresponds to the siblings fixed effects estimations. Total number of (weighted) observations = 118,540. Column (1) shows the relative frequency of each age at immigration in the sample in percentage (column sums to 100). Column (2) gives the average family size by age at immigration, i.e. the number of children in the family, including the respondent. Columns (3) and (4) give the distribution of child sex by age at immigration in percentage (rows sum to 100). Columns (5) and (6) give the distribution of maternal schooling attainment by age at immigration in percentage (rows sum to 100; HS = high school). Columns (7) and (8) give the distribution of paternal schooling attainment by age at immigration in percentage (rows sum to 100). Columns (9) and (10) give the distribution of region of origin by age at immigration in percentage (rows sum to 100).

Table C3: Exposure Effects by Subsample

Specification	No controls		With siblings fixed effects	
	Up to age 10	Age 10 and above	Up to age 10	Age 10 and above
A: Baseline	-0.188 (.036)	-0.766 (.045)	0.085 (.027)	-0.423 (.077)
B: Child Sex				
Sons	-0.302 (.080)	-0.558 (.159)	0.008 (.045)	-0.451 (.098)
Daughters	-0.138 (.071)	-0.884 (.100)	-0.318 (.100)	-0.511 (.158)
C: Maternal Schooling Attainment				
HS or less	-0.362 (.059)	-0.831 (.062)	-0.156 (.071)	-0.549 (.120)
Above HS	-0.024 (.089)	-0.553 (.131)	0.03 (.113)	-0.088 (.141)
D: Child Sex and Maternal Schooling Attainment				
Sons with Mother HS or less	-0.446 (.081)	-0.511 (.230)	-0.021 (.112)	-0.381 (.181)
Sons with Mother above HS	-0.157 (.212)	-0.439 (.119)	0.033 (.132)	-0.219 (.110)
Daughters with Mother HS or less	-0.411 (.086)	-0.871 (.120)	-0.540 (.110)	-0.707 (.206)
Daughters with Mother above HS	0.154 (.098)	-0.869 (.227)	0.132 (.194)	-0.376 (.310)
E: Region of Origin				
Asia	-0.314 (.094)	-1.075 (.070)	-0.291 (.054)	-0.713 (.130)
Rest of World	-0.164 (.028)	-0.287 (.091)	0.022 (.048)	-0.015 (.064)
F: Mother Tongue				
English or French	-0.001 (.052)	-0.296 (.140)	0.086 (.056)	-0.226 (.197)
Neither English nor French	-0.250 (.058)	-0.838 (.031)	-0.184 (.050)	-0.517 (.103)

Source: Authors' calculations based on the IID+.

Note: This table reports the slope coefficients for the two segments of a piecewise (spline) regression with a kink at age 10 estimated on the \hat{b}_m coefficients coming from the models without controls (Equation 1) and with the siblings fixed effects (Equation 2), by subsample. For the baseline, this corresponds to what is depicted in Figure 2. Robust standard errors are in parentheses.

Table C4: Rank Mobility by Generation Status, Age at Immigration and Region of Origin

Generation status Age at immigration	3+	2	1.5			1.5	
	(1)	(2)	0-5yo (3)	6-11yo (4)	12-17yo (5)	0-17yo (all) (6)	(adjusted) (7)
All	0.250 (0.001)	0.166 (0.002)	0.154 (0.004)	0.180 (0.004)	0.187 (0.004)	0.176 (0.002)	0.158 (0.002)
<i>N</i>	1,959,000	553,090	109,450	116,190	109,090	334,730	334,730
<i>Region of origin</i>							
North America, Northern and Western Europe	0.165 (0.003)	0.141 (0.007)	0.162 (0.008)	0.133 (0.011)	0.145 (0.005)	0.138 (0.005)	
Caribbean, Central and Southern America, and Oceania	0.203 (0.005)	0.175 (0.010)	0.199 (0.010)	0.170 (0.012)	0.182 (0.006)	0.167 (0.006)	
Southern and Eastern Europe	0.186 (0.003)	0.170 (0.010)	0.182 (0.010)	0.157 (0.012)	0.164 (0.006)	0.147 (0.006)	
Africa	0.192 (0.010)	0.205 (0.016)	0.215 (0.015)	0.173 (0.018)	0.201 (0.009)	0.186 (0.009)	
Asia	0.152 (0.004)	0.166 (0.007)	0.196 (0.006)	0.221 (0.007)	0.216 (0.004)	0.193 (0.004)	

Source: Authors' calculations based on the IID+.

Note: Each column within a row contains the result of a separate regression of the rank-rank mobility model. For each subgroup defined by the row and column titles, the table presents the rank-rank mobility coefficient ($\hat{\beta}$) and its robust standard error in parentheses. In the “adjusted” column (7), child income ranks are adjusted for the age-at-immigration effects before the rank-rank model is estimated. Robust standard errors are in parentheses. The third row (*N*) gives the number of (weighted) observations.