BRIDGING THE GAP: ABSOLUTE AND RELATIVE INTERGENERATIONAL EDUCATIONAL MOBILITY*

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Abstract

Does educational expansion disrupt intergenerational transmission of advantage, or merely shift inequality upward while preserving relative positions? Using the World Bank's Global Database on Intergenerational Mobility covering 140+ countries across five birth cohorts (1940s–1980s), this paper examines how absolute and relative educational mobility relate. Three findings emerge. First, a negative association exists between the intergenerational persistence coefficient and absolute upward mobility: a 10-percentage-point increase in absolute mobility associates with 5.5-percentage-point decrease in persistence, strengthening nearly fourfold from 1940s to 1980s cohorts. Second, variance-adjusted measures reveal relationships only through cohort interactions, indicating temporal dynamics mediate these associations. Third, profound gender differences characterize recent cohorts, with daughters demonstrating different mobility patterns than sons. These findings document robust descriptive regularities indicating that absolute and relative educational mobility move together under specific conditions. While these observational patterns do not identify causal disruption mechanisms, the temporal strengthening of the absolute-relative relationship is highly consistent with theories of equity-focused institutional disruption. Crucially, this relationship survives a barrage of demanding econometric tests, including country-specific cohort trends, correlated random effects, and precision weighting, establishing a stringent set of stylized facts that modern stratification theories must accommodate.

Keywords: Social Mobility; Educational Attainment; Intergenerational Persistence; Human Capital; Educational Expansion;
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1 Introduction

Intergenerational mobility (the degree to which socioeconomic advantages or disadvantages pass from one generation to the next) lies at the heart of debates about inequality, opportunity, and social justice (Torche, 2015). Education, as both a marker of status and a vehicle for social change, occupies a central role in this process. It can perpetuate privilege when access remains unequal, yet it also holds the potential to disrupt entrenched hierarchies when broadly expanded (Hout and DiPrete, 2006). The sociology and economic literatures have long distinguished between *absolute mobility* (concrete improvements in attainment driven by structural shifts like educational expansion) and *relative mobility*, which reflects social fluidity independent of marginal distributions (Featherman et al., 1975).

A fundamental question remains unresolved: does educational expansion disrupt intergenerational transmission of advantage, or merely shift inequality upward while preserving relative positions? This question carries important theoretical and policy implications. If absolute and relative mobility systematically diverge, rising educational attainment need not translate into equal opportunity. Policymakers must explicitly target stratification mechanisms beyond enrollment expansion. If these dimensions move together conditionally, this identifies institutional arrangements that successfully convert expansion into equalization.

This paper addresses this challenge by examining how absolute and relative educational mobility interrelate using the World Bank's Global Database on Intergenerational Mobility (*GDIM*), a comprehensive resource covering over 140 countries (van der Weide et al., 2024). My analysis employs three established measures of relative mobility: 1) the intergenerational (years of schooling) regression coefficient (`IGP`, β), which gauges how strongly parental education predicts a child's educational attainment (Solon, 1999); 2) the intergenerational correlation coefficient (`IGC`, ρ), which adjusts for changes in educational inequality over time (Black and Devereux, 2011); and 3) the expected rank of descendants (`EXRANK`) born to parents in the bottom half of the education distribution (Chetty et al., 2014). For absolute mobility, I use the probability of descendants surpassing their parents' ordinal education level (conditional on parents' education not exceeding tertiary education), capturing upward movement in concrete terms (Chetty et al., 2017). This approach separates structural changes (like widespread schooling expansion) from underlying social fluidity, providing insight into education's role in mediating opportunity across diverse national contexts.

Three principal findings emerge. First, a robust negative association exists between `IGP` and absolute upward mobility: a 10-percentage-point increase in absolute mobility associates with approximately 5.5-percentage-point decrease in persistence unconditionally, strengthening to nearly 11 percentage points for the 1980s cohort. Cohort-absolute mobility interactions are positive and highly significant, indicating this relationship strengthens nearly fourfold from the 1940s to 1980s cohorts. This temporal strengthening pattern is difficult to reconcile with Maximally Maintained Inequality theory's prediction that hierarchical diffusion produces strongest effects during early expansion phases. It aligns more closely with frameworks emphasizing increasingly equity-focused policy environments.

Second, the relationship between `IGC` and absolute mobility is more nuanced. Unconditional specifications yield null relationships, which become significantly negative only when cohort interactions are included. This pattern indicates that variance evolution mediates these associations, with educational expansion producing complex, non-proportional variance dynamics that require explicit temporal modeling.

Third, profound gender differences characterize mobility patterns, particularly in recent cohorts. While sons exhibit significant cohort effects in unconditional specifications, daughters show insignificant cohort

effects that become highly significant through interaction terms. This heterogeneity suggests that daughters' mobility operates through different mechanisms than sons', with policy implementation determining whether absolute mobility gains relate to weakened intergenerational persistence.

Beyond these core findings, the analysis reveals regional variation. Latin America and the Caribbean, South Asia, and the Middle East and North Africa consistently exhibit higher persistence than North America, even after controlling for absolute mobility levels. These regional patterns underscore how institutional contexts and historical legacies shape mobility outcomes across diverse development settings.

This study offers three principal contributions to the international development and mobility literature. First, it establishes a set of robust descriptive regularities showing that absolute and relative educational mobility need not systematically diverge. Across clustered, trend-adjusted, correlated-random-effects, and precision-weighted specifications, the central association remains substantively similar. Second, by documenting that mobility relationships strengthen in recent cohorts and differ markedly by gender, it establishes that both demographic and temporal factors mediate mobility processes in ways obscured by aggregate analysis. Third, by demonstrating that different mobility measures illuminate distinct transmission dimensions (`IGP` captures expansion's compositional impact, `IGC` requires modeling variance evolution, and `EXRANK` reveals whether expansion benefits disadvantaged groups), it highlights the importance of employing multiple measures in cross-national mobility research.

The rest of the paper proceeds as follows. Section 2 discusses related literature. Section 3 develops the theoretical framework generating testable predictions. Section 4 describes the data and methodology. Section 5 reports the results. Section 6 concludes.

2 Related Literature

This paper builds on three literatures: conceptual frameworks distinguishing absolute from relative mobility, theories of educational expansion and stratification, and cross-national empirical research on intergenerational transmission.

The study of intergenerational mobility spans multiple disciplines, with economists typically focusing on income or earnings elasticities to quantify economic inheritance (Solon, 1999; Black and Devereux, 2011), while sociologists emphasize occupational prestige, social class, and education as key axes of stratification (Ganzeboom et al., 1991; Hout and DiPrete, 2006). This interdisciplinary foundation has generated rich theoretical frameworks for understanding how status persists or shifts across generations, with education emerging as an important mediator of intergenerational transmission in both developed and developing contexts.

The literature has established a fundamental distinction between absolute and relative mobility. Absolute mobility captures upward or downward shifts in educational attainment, often driven by structural changes like the expansion of schooling systems (Featherman et al., 1975). Relative mobility, by contrast, measures social fluidity (changes in an individual's position within the educational hierarchy, independent of overall growth in attainment). Early theoretical work, including the Featherman-Jones-Hauser hypothesis and research from the CASMIN project, pointed to institutional differences (such as funding models or access policies) as key drivers of cross-national variation in mobility patterns (Ganzeboom et al., 1991).

More recently, Deutscher and Mazumder (2023) provide a comprehensive framework for understanding different mobility measures and their relationships, demonstrating that while measures within broad categories tend to be highly correlated, measures across categories often show weak correlations. Bloome (2025) develops frameworks decomposing mobility into marginal shifts versus dependence changes, show-

ing mathematically how these components vary independently. These advances establish that absolute and relative mobility *can* diverge theoretically. Whether they *do* diverge empirically, and under what conditions, remains the central question this paper addresses.

Educational expansion generates competing theoretical predictions about its implications for intergenerational transmission. Maximally Maintained Inequality (Raftery and Hout, 1993) posits expansion proceeds hierarchically: advantaged groups saturate each level before expansion reaches disadvantaged groups, creating rising absolute mobility without disrupting relative mobility. Effectively Maintained Inequality (Lucas, 2001) extends this logic: even when quantitative expansion equalizes years of schooling, advantaged families may secure qualitative advantages through selective institutions and enrichment activities. These frameworks predict negative correlation between β and absolute mobility reflecting compositional shifts rather than altered association strength. If expansion operates through hierarchical diffusion, relationships should be strongest during early expansion phases, then weaken as systems saturate.

Alternative frameworks emphasize expansion's potential to disrupt advantage transmission when accompanied by equity-promoting policies. Research on advanced economies demonstrates that massive expansion in education can reduce inequalities in social mobility chances, with evidence showing that class disparities in educational attainment have generally declined in Europe (Breen et al., 2009, 2010). Similarly, van de Werfhorst (2024) finds that educational expansion enhances occupational mobility across 40 countries without systematic elite counteraction. These frameworks predict absolute and relative mobility should exhibit negative correlation when expansion targets disadvantaged groups, reflecting altered association strength rather than pure compositional shifts. If policy environments become more equity-focused over time, relationships should strengthen in recent cohorts.

Adjudicating between frameworks requires examining temporal dynamics and multiple measures. Compositional theories predict relationships strengthen early then weaken; disruption theories predict strengthening in recent cohorts. Compositional theories predict variance-adjusted measures (ρ) show weak relationships; disruption theories predict comparable or stronger relationships than β.

Nybom and Stuhler (2024) demonstrate that past expansion generates long-lasting, non-monotonic effects on contemporary patterns. This is particularly relevant for developing countries experiencing compressed expansion (achieving in 20–30 years what took advanced economies 50–80 years). Rapid enrollment growth may produce non-proportional variance changes: if children's educational variance expands faster than parents', this affects correlation coefficients independently of association changes. Policy regimes may shift during compressed expansion, beginning with hierarchical diffusion then adopting equity-focused policies as systems mature. These complexities suggest cohort interactions warrant explicit investigation.

Gender dynamics in intergenerational mobility have received increasing attention as traditional father-son models have expanded to include mother-child linkages and gender-specific patterns. Hu and Qian (2023) demonstrate that educational expansion strengthens mother-child associations, particularly in contexts where mothers outrank fathers educationally. Research has documented gendered convergence in educational transitions, showing that women and men often follow distinct intergenerational mobility paths (Bukodi et al., 2015). Nennstiel and Becker (2023) illustrate that while relative social mobility may fall, strong gender convergence can occur in certain dimensions of mobility. Hossain et al. (2025) analyze 92 countries, finding daughters systematically exceed sons in recent cohorts, particularly in historically gender-unequal regions.

In historically gender-unequal contexts, expansion may affect sons and daughters through different mechanisms. For daughters, expansion represents both compositional change (exclusion to access) and

potential disruption of patriarchal transmission patterns; for sons, expansion operates within existing class hierarchies. If so, negative absolute-relative relationships should be stronger for daughters in historically unequal regions.

Cross-national measurement and comparison of intergenerational mobility face significant methodological challenges. Kunievsky (2024) highlights difficulties in interpreting cross-country comparisons using standard mobility measures, emphasizing limitations of linear regression coefficients as summary statistics. Torche (2015) notes that different mobility measures may capture distinct reproduction mechanisms. The World Bank's GDIM addresses many of these concerns by providing standardized, education-specific data across a vast array of countries with varying development levels (van der Weide et al., 2024).

Research on developing countries has provided valuable insights into mobility patterns distinct from those observed in advanced economies. Alesina et al. (2021) investigate inequality and intergenerational mobility in education across Africa, while Neidhöfer et al. (2018) focus on educational mobility in Latin America. These studies highlight how structural barriers in developing contexts can hamper both absolute and relative mobility, with regional disparities often reflecting institutional weaknesses and economic constraints.

Three gaps remain in the literature. First, no study systematically tests whether absolute and relative measures empirically diverge across diverse contexts. Existing work focuses on single countries (Deutscher and Mazumder, 2023) or documents descriptive patterns without examining relationships between measures (van der Weide et al., 2024). Second, temporal dynamics (particularly cohort interactions that could discriminate between competing theories) remain underexplored despite theoretical predictions. Third, gender-differentiated mechanisms warrant investigation given recent evidence of daughters' accelerating mobility and theoretical predictions about disruption versus composition.

This paper addresses these gaps using comprehensive global data spanning 140+ countries and five birth cohorts to test: (1) Do absolute and relative mobility measures show consistent relationships across institutional contexts? (2) How do relationships evolve across cohorts as systems expand? (3) Do patterns differ systematically by gender and region? Answering these questions adjudicates between competing theories and identifies conditions under which expansion translates into genuine equalization versus preserved stratification through new mechanisms.

3 Theoretical Framework and Empirical Predictions

Educational mobility scholarship distinguishes conceptually between absolute and relative dimensions, yet lacks frameworks predicting how these empirically relate across diverse institutional contexts. This gap matters because competing stratification theories generate distinct predictions about whether these dimensions move together or diverge, with important implications for understanding educational expansion's role in social equalization.

To keep the main text intuitive, the discussion below remains verbal. Appendix A presents a compact statistical framework that maps these verbal mechanisms into three objects: an upward shift in average schooling, a persistence parameter linking parental and child attainment, and a variance component shaping correlation-based measures. This mapping is useful because it clarifies why absolute mobility can rise even when relative mobility changes little, and why `IGP` and `IGC` need not move one-for-one.

3.1 Competing Theoretical Frameworks

Maximally Maintained Inequality (Raftery and Hout, 1993) posits that educational expansion proceeds hierarchically: advantaged groups saturate each level before additional capacity reaches disadvantaged groups. Effectively Maintained Inequality (Lucas, 2001) extends this logic by arguing that even when quantitative expansion equalizes years of schooling, advantaged families may preserve advantages through selective institutions, fields of study, and other quality margins.

In the language of Appendix A, these accounts primarily emphasize a broad upward shift in schooling levels rather than a large decline in the dependence of child outcomes on parental background. Educational expansion therefore raises absolute mobility because more children cross thresholds their parents did not reach, but relative mobility may change only modestly if the persistence parameter remains high or if new quality margins reproduce advantage. Under this view, a negative association between absolute mobility and IGP can still arise, but much of it reflects changing marginal distributions rather than a fundamental weakening of intergenerational transmission.

These compositional accounts also help explain why variance-adjusted measures may behave differently from persistence coefficients. If expansion alters dispersion across generations or shifts competition toward unmeasured qualitative margins, then IGC may remain only weakly related to absolute mobility even when IGP declines. Moreover, if hierarchical diffusion is strongest in earlier stages of expansion and weakens as systems mature, the negative absolute-IGP relationship should be steeper for earlier cohorts than for later cohorts.

Alternative frameworks emphasize expansion's potential to weaken intergenerational persistence when it is accompanied by equity-promoting policies. Breen et al. (2009) document declining class inequalities in European educational attainment, while van de Werfhorst (2024) finds that expansion can enhance mobility without systematic elite counteraction. These perspectives imply not only an upward shift in schooling levels, but also a decline in the extent to which parental schooling predicts child schooling.

In the appendix framework, disruption-oriented accounts therefore combine a positive level effect with a reduction in the persistence parameter. Under that configuration, absolute mobility and relative mobility improve together: more children surpass their parents, and parental background becomes less predictive of their eventual attainment. If policy environments become more equity-oriented over time, the negative relationship between absolute mobility and IGP should be stronger for later cohorts than for earlier cohorts. Because the empirical specification omits the 1980s cohort, that prediction implies positive coefficients on the earlier-cohort interaction terms: those coefficients measure how much weaker the slope is relative to the more negative 1980s benchmark.

3.2 Gender as Discriminating Test Case

Gender provides a particularly revealing empirical test of competing frameworks. In historically gender-unequal contexts, expansion affects sons and daughters through different mechanisms. For daughters in regions with historical female educational exclusion, expansion represents both compositional gains (access where none existed) and potential disruption of patriarchal transmission patterns as girls' education campaigns and normative shifts specifically target disadvantaged backgrounds. For sons in the same regions, expansion operates primarily within existing class hierarchies (compositional change without fundamental disruption of transmission mechanisms).

This generates distinct predictions. In historically gender-unequal regions (South Asia, Middle East and North Africa, Sub-Saharan Africa), daughters should show stronger absolute mobility gains than sons

in recent cohorts. The absolute-relative mobility relationship should differ by gender: for daughters, negative relationships may reflect genuine disruption of traditional barriers; for sons, they more likely reflect compositional expansion. Cohort-absolute mobility interactions should be stronger for daughters if targeted policies successfully disrupt traditional transmission patterns.

3.3 Institutional Context and Regional Variation

Institutional contexts moderate how expansion affects mobility dimensions. Educational systems with universal access, minimal tracking, equalized funding, and strong redistribution may convert expansion into equalization more effectively than stratified systems with sharp public-private quality gaps and limited redistribution. This predicts regional variation: regions with more comprehensive systems should show strong negative relationships for both β and ρ; regions with stratified systems should show strong negative β but weak ρ relationships if quality differentiation preserves advantages.

3.4 Temporal Dynamics and Cohort Effects

Nybom and Stuhler (2024) demonstrate that past expansion generates long-lasting, non-monotonic effects on contemporary patterns. For developing countries experiencing compressed expansion, temporal dynamics warrant particular attention because the relevant components of mobility may change at different speeds across cohorts.

In the appendix framework, cohort variation can arise through three channels. First, later cohorts may experience larger upward shifts in average schooling, mechanically increasing absolute mobility. Second, the persistence parameter may weaken if later reforms target disadvantaged groups more directly. Third, the variance of child schooling may change non-proportionally relative to parental schooling, affecting correlation-based measures even when persistence itself moves in the same direction.

These distinctions generate testable predictions about cohort interactions. If expansion increasingly reaches disadvantaged groups and weakens persistence over time, the negative relationship between absolute mobility and relative persistence should be stronger in later cohorts. In the empirical specification, where the 1980s cohort is the omitted category, this appears as positive interaction coefficients on earlier cohorts because those coefficients indicate that the slope was less negative before the 1980s. By contrast, if hierarchical diffusion dominates early and then fades as systems saturate, earlier cohorts should exhibit the steeper negative relationship, implying negative interaction coefficients relative to the 1980s benchmark. The pattern of cohort interactions therefore provides a useful descriptive way to distinguish between compositional and disruption-oriented interpretations.

3.5 Summary of Testable Predictions

Table 1 summarizes how different empirical patterns support or contradict competing theoretical frameworks.

The key discriminating test involves cohort interactions. Maximally Maintained Inequality predicts positive cohort × absolute mobility interactions (relationship weakens as systems saturate and hierarchical diffusion completes). Disruption theories predict negative interactions (relationship strengthens as policies become more equity-focused). Because the intergenerational persistence coefficient enters as the dependent variable (higher values indicate lower mobility), positive interaction coefficients in regressions mean the negative absolute-relative relationship strengthens in recent cohorts, supporting disruption theories and contradicting hierarchical diffusion predictions.

Table 1: Theoretical Predictions and Discriminating Empirical Tests

Empirical Pattern	Supports	Contradicts
Strong negative IGP-absolute correlation	Both theories	Neither
Positive cohort × absolute interactions (relationship strengthens over time)	Disruption theories	MMI/EMI hierarchical diffusion
Strong negative IGC-absolute correlation (when controlling for cohort interactions)	Disruption theories	MMI/EMI quality stratification
Stronger relationships for daughters (historically unequal regions)	Targeted disruption	Universal composition
Regional variation in IGC relationships	Institutional	Universal

This framework guides the empirical analysis. I test whether: (1) cohort interactions support hierarchical diffusion or disruption mechanisms; (2) variance-adjusted measures show patterns consistent with quality stratification or genuine association changes; (3) gender heterogeneity reveals targeted disruption or universal compositional effects; and (4) regional variation reflects institutional differences in converting expansion into equalization.

4 Data and Methodology

4.1 Data

This study utilizes the World Bank's Global Database on Intergenerational Mobility (GDIM) by van der Weide et al. (2024), which provides comprehensive data on intergenerational mobility (IGM) in education across over 140 countries. The GDIM covers individuals born between 1940 and 1989, grouped into five 10-year birth cohorts (1940-1949, 1950-1959, 1960-1969, 1970-1979, and 1980-1989).

The GDIM focuses on educational mobility for several compelling reasons. First, human capital is fundamental to economic well-being, and education data across generations is more readily available compared to income data. Second, the estimation of educational mobility presents fewer methodological challenges than income mobility: education levels, once attained, remain constant throughout an individual's life, providing a stable measure for comparison.[1] Third, people can typically report their parents' education levels with high accuracy, enabling mobility studies without the need for extensive panel data, thereby reducing concerns regarding measurement errors. These factors collectively make education an ideal metric for studying intergenerational mobility across diverse global contexts.

The database provides estimates for different subpopulations:

- For parental educational attainment: 1) Mothers, 2) Fathers, 3) Average, and 4) Maximum
- For children's educational attainment: 1) Sons, 2) Daughters, and 3) All respondents

This allows for exploring various mobility relationships, such as mother-daughter or father-son mobility combinations. For each economy and cohort, the GDIM provides 12 estimates covering these different subpopulations. In this study, I investigate each of these 12 combinations in detail by using country and (10-year) cohort-specific estimates for these 12 parent and descendant combinations as the unit of analysis in regressions.[2] The final dataset consists of observations from over 140 countries spanning five birth

[1] See Nybom and Stuhler (2017) and Nybom and Stuhler (2016) for general concerns about estimating intergenerational income mobility consistently. See also Demirtaş and Torul (2024) for issues regarding developing economies.

[2] I append the cases of maximum parental education to the Online Appendix for brevity. The results when maximum parental education is used are often quantitatively similar to the case of father's education.

cohorts.[3]

4.2 Methodology

In this study, I employ multiple measures of intergenerational mobility to capture both absolute and relative aspects of mobility:

4.2.1 Absolute Mobility (ABSCAT)

I use the probability of descendants surpassing their parents' ordinal education level as a measure of absolute mobility. This metric captures the extent to which individuals achieve higher educational attainment than their parents conditional on their parental education being less than tertiary education.[4] Formally, this can be expressed as:

$$\text{ABSCAT} = P(E_c > E_p | E_p < \text{Tertiary})$$

where E_c and E_p represent the educational categories of children and parents, respectively.

4.2.2 Relative Mobility Measures

I employ three measures of relative mobility to test whether relationships with absolute mobility reflect compositional change versus genuine disruption of stratification:

Intergenerational Persistence (`IGP`, β). This measure captures the persistence of educational attainment across generations.[5] It is estimated from the regression equation:

$$Y_c = \alpha + \beta Y_p + \epsilon$$

where Y_c and Y_p are years of schooling for children and parents, respectively. This captures persistence including both compositional effects from marginal distribution changes and association strength. Higher β indicates stronger intergenerational persistence (lower mobility).

Intergenerational Correlation Coefficient (`IGC`, ρ). This metric accounts for changes in the variance of educational attainment across generations:

$$\rho = \beta \frac{\sigma_p}{\sigma_c}$$

where σ_p and σ_c denote the years of schooling standard deviations of parents and children, respectively (Black and Devereux, 2011). By adjusting for variance changes across generations, this isolates association

[3] Not all countries have data for all five cohorts due to survey availability and data quality constraints. For example, Table 1 reports 549 observations covering 142 countries when using average parental education and both sons and daughters. Observation counts vary slightly across specifications (ranging from 538 to 559) depending on which parent-child combination is analyzed, as some countries lack data for specific subpopulations or cohorts.

[4] For the ordinal education categories, I rely on the same 5-level categorization by GDIM: 1-Less than primary school graduate (ISCED 0), 2-Primary school graduate (ISCED 1), 3- Lower secondary school graduate (ISCED 2), 4- Upper secondary or post-secondary (non-tertiary) school graduate (ISCED 3-4), and 5-Tertiary school graduate (ISCED 5-8). As a robustness check, I also examine the case when the frequency of descendants surpassing their parents' years of education (ABSYOS) is proxied for absolute mobility. These results are available upon request.

[5] Several studies use `IGP` as an inverse measure of mobility as it reflects the association between parental and child educational attainment. An `IGP` value of 1 signifies a perfect correlation, implying that children's years of schooling, on average, precisely mirror that of their parents. Conversely, an `IGP` of 0 indicates no statistically significant association between parental and child educational levels. In this scenario, all children have an equal chance of educational attainment, regardless of their socioeconomic background. For a comprehensive recent discussion on intergenerational mobility, see Mogstad and Torsvik (2023).

strength from compositional effects. Comparing IGP and IGC patterns tests whether observed relationships reflect variance expansion (supporting compositional hypotheses) versus genuine changes in association strength (supporting disruption hypotheses).

Expected Rank (EXRANK). This measure indicates the expected educational rank of descendants when their parents are in the bottom half of the education distribution. It ranges from 0 to 100, with values above 50 indicating upward mobility (Chetty et al., 2014). This measure is particularly sensitive to whether expansion benefits the bottom versus middle of the distribution.

4.3 Empirical Specifications

To analyze the relationship between absolute and relative mobility, I employ five progressively richer regression models designed to test the theoretical predictions developed in Section 3:

Model 1 (Country fixed effects):

$$\text{RelMob}_{i,c,t} = \alpha + \beta \text{AbsMob}_{i,c,t} + \gamma_c + \epsilon_{i,c,t}$$

This baseline specification controls for time-invariant country characteristics, estimating the within-country correlation between absolute and relative mobility across cohorts. It establishes whether the basic negative association between IGP and absolute mobility holds when exploiting within-country temporal variation.

Model 2 (Random effects GLS):

$$\text{RelMob}_{i,c,t} = \alpha + \beta \text{AbsMob}_{i,c,t} + u_c + \epsilon_{i,c,t}$$

This specification exploits both within and between-country variation, testing whether relationships hold across institutional contexts.[6] Comparing Models 1 and 2 provides evidence on whether the absolute-relative mobility relationship reflects universal patterns or country-specific factors.

Model 3 (Country and cohort fixed effects):

$$\text{RelMob}_{i,c,t} = \alpha + \beta \text{AbsMob}_{i,c,t} + \gamma_c + \delta_t + \epsilon_{i,c,t}$$

This specification adds cohort fixed effects, testing whether relative mobility exhibits systematic temporal trends beyond changes in absolute mobility. Significant cohort fixed effects would indicate that mobility patterns evolved across birth cohorts through channels beyond educational expansion captured by absolute mobility.

Model 4 (Country FE, cohort FE, and cohort interactions):

$$\text{RelMob}_{i,c,t} = \alpha + \beta \text{AbsMob}_{i,c,t} + \gamma_c + \delta_t + \sum_{t' \neq 1980s} \theta_{t'} \left(\mathbb{1}[t = t'] \times \text{AbsMob}_{i,c,t} \right) + \epsilon_{i,c,t}$$

This is the key specification for discriminating between competing theoretical frameworks. The cohort interaction terms $\{\theta_{t'}\}$ test whether the absolute-relative mobility relationship evolved across birth cohorts. Because IGP enters as the dependent variable (higher values indicate lower mobility), positive θ coefficients indicate the negative relationship between absolute mobility and IGP strengthens over time.

[6] Random effects treat countries as random draws from a broader population. Similarity between Models 1 and 2 serves as an informal specification test, providing evidence that country-specific factors do not bias the estimated relationship.

As established in Section 3, Maximally Maintained Inequality predicts positive θ values (relationship weakens as hierarchical diffusion completes and systems saturate), while disruption theories predict negative θ values (relationship strengthens as policies become more equity-focused). The pattern of these interactions thus provides evidence about whether expansion operates through hierarchical diffusion or increasingly targets disadvantaged groups.

Model 5 (Region FE, cohort FE, and cohort interactions):

$$\text{RelMob}_{i,c,t} = \alpha + \beta \text{AbsMob}_{i,c,t} + \lambda_r + \delta_t + \sum_{t' \neq 1980s} \theta_{t'} \left(\mathbb{1}[t = t'] \times \text{AbsMob}_{i,c,t} \right) + u_c + \epsilon_{i,c,t}$$

This specification examines regional patterns while maintaining the cohort interaction structure. By replacing country fixed effects with region fixed effects (following the World Bank's seven-region classification system), it tests whether the temporal dynamics identified in Model 4 remain robust and examines regional heterogeneity in mobility levels. The random country effect u_c accounts for within-region country variation.

where $\text{RelMob}_{i,c,t}$ represents relative mobility (`IGP`, `IGC`, or `EXRANK`) for subpopulation i in country c and cohort t; $\text{AbsMob}_{i,c,t}$ is absolute mobility (ABSCAT); γ_c are country fixed effects; δ_t are cohort fixed effects (base: 1980s); λ_r are region fixed effects following World Bank's seven-region classification (base: North America); $\{\theta_{t'}\}$ are cohort-specific interaction coefficients capturing how the absolute-relative mobility relationship varies across birth cohorts; u_c is country-specific random effect; and $\epsilon_{i,c,t}$ is the error term.

Together, these five baseline models test the core theoretical predictions. However, cross-national panel regressions of this nature are susceptible to severe omitted variable bias and serial correlation. To ensure these descriptive regularities are genuinely robust, I implement a dedicated robustness package (detailed in Section 5.1). I cluster all standard errors at the country level to correct for autocorrelation. I employ Correlated Random Effects (the Mundlak device) to explicitly test within- versus between-country variance. I introduce precision weighting using the inverse variance of the mobility estimates. Most critically, I introduce country-specific linear cohort trends to absorb slow-moving unobservables (e.g., democratization or welfare state expansion) that might spuriously drive both educational expansion and social fluidity.

I conduct separate analyses for different subpopulations (e.g., sons only, daughters only) to explore potential heterogeneity in mobility patterns across genders, as predicted by the theoretical framework. Additionally, I examine how the relationship between absolute and relative mobility varies across geographical regions, providing insights into how institutional contexts moderate these relationships. [7]

5 Results

This section presents empirical findings on the relationship between absolute and relative intergenerational educational mobility. I organize results around the theoretical framework developed in Section 3, focusing on how empirical patterns support or contradict competing theories of educational expansion and stratification.

[7] Unless otherwise noted, reported standard errors are heteroskedasticity-robust; Section 5.1 additionally reports country-clustered standard errors as part of the robustness analysis.

5.1 Robustness and Institutional Heterogeneity

Table 2 reports a set of robustness checks for the baseline average-parent/all-children specification. Across specifications that cluster standard errors by country, add country means of absolute mobility, absorb country-specific linear cohort trends, and reweight observations by estimate precision, the negative association between absolute mobility and intergenerational persistence remains negative and substantively similar.

Model 1 clusters standard errors at the country level, confirming that the negative association remains statistically significant when accounting for within-country serial correlation. Model 2 introduces the Mundlak device, showing that the association remains similar when controlling for country-level historical averages of absolute mobility. Model 3 adds country-specific linear cohort trends, a demanding specification that absorbs smoothly evolving country-specific unobservables, and the estimated association remains negative and statistically significant. Model 4 shows that the same pattern holds when observations are weighted by the inverse variance of the IGP estimate.

Model 5 explores whether the association varies with parental mean education. The estimated interaction between absolute mobility and parental mean education is negative, but it is imprecisely estimated in this specification, so these results should be interpreted as suggestive rather than conclusive. Taken together, Table 2 indicates that the core negative association is not driven solely by one particular choice of standard errors, panel estimator, weighting scheme, or trend adjustment.

Table 2: Headline Robustness Package: Average Parents & All Children

	(1)	(2)	(3)	(4)	(5)
VARIABLES	Clustered FE	CRE/Mundlak	Country Trends	Weighted FE	Parental Ed. Het.
Absolute Mobility	-0.498***	-0.499***	-0.360***	-0.657***	-0.661***
	(0.092)	(0.092)	(0.137)	(0.076)	(0.077)
Parental Mean Education (centered)					-0.033
					(0.021)
Absolute Mobility x Parental Mean Education					-0.023
					(0.020)
1940 Cohort	0.086***	0.080***	-0.053**	0.150***	-0.083
	(0.026)	(0.026)	(0.021)	(0.023)	(0.075)
1950 Cohort	0.075***	0.068***	-0.031	0.146***	-0.053
	(0.020)	(0.020)	(0.035)	(0.018)	(0.053)
1960 Cohort	0.041***	0.034**	-0.033	0.097***	-0.053
	(0.014)	(0.015)	(0.027)	(0.016)	(0.036)
1970 Cohort	0.012	0.005	-0.022	0.047***	-0.038*
	(0.013)	(0.013)	(0.016)	(0.012)	(0.021)
Country Mean Absolute Mobility		-0.120			
		(0.104)			
Constant	0.771***	0.848***	7.371***	0.868***	0.948***
	(0.043)	(0.054)	(1.389)	(0.045)	(0.053)
Observations	549	549	549	549	549
R-squared	0.213	0.250	0.695	0.898	0.275
Number of countrycode	142	142	142	142	142
Country FE	Yes	Mundlak	Yes	Yes	Yes
Cohort FE	Yes	Yes	Yes	Yes	Yes
Country Trends	No	No	Yes	No	No
Weighting	No	No	No	Yes	No

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

† *Notes*: Asterisks denote statistical significance (*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$). Heteroskedasticity-robust standard errors clustered at the country level are reported in parentheses. The dependent variable is the intergenerational regression coefficient (`IGP`). Absolute mobility is `ABSCAT`. Model 2 includes country-level means of `ABSCAT`. Model 3 includes country-specific linear cohort trends. Model 4 is weighted by the inverse variance of the `IGP` estimate. Model 5 includes an interaction with centered parental mean years of schooling (`c_MEANp`).

5.2 Intergenerational Persistence: Evidence for Disruption Theories

Table 3 presents estimation results using intergenerational persistence (`IGP`) as the measure of relative mobility and the probability of descendants surpassing their parents' ordinal education (ABSCAT) as the

measure of absolute mobility. For IGP estimation, I use the arithmetic mean of both parents' years of schooling as the independent variable, while descendants' years of schooling (including both sons and daughters) serves as the dependent variable.

Table 3: Estimation Results for Average Parental Education & Both Sons and Daughters

VARIABLES	(1) Model 1	(2) Model 2	(3) Model 3	(4) Model 4	(5) Model 5
Absolute Mobility	-0.548***	-0.573***	-0.498***	-1.066***	-0.888***
	(0.106)	(0.084)	(0.092)	(0.087)	(0.082)
1940 Cohort			0.086***	-0.338***	-0.259**
			(0.026)	(0.105)	(0.105)
1950 Cohort			0.075***	-0.200***	-0.132*
			(0.020)	(0.075)	(0.078)
1960 Cohort			0.041***	-0.138***	-0.089*
			(0.014)	(0.048)	(0.050)
1970 Cohort			0.012	-0.185***	-0.148***
			(0.013)	(0.043)	(0.044)
Absolute Mobility x 1940 Cohort				0.808***	0.666***
				(0.170)	(0.170)
Absolute Mobility x 1950 Cohort				0.507***	0.381***
				(0.118)	(0.123)
Absolute Mobility x 1960 Cohort				0.335***	0.243***
				(0.080)	(0.081)
Absolute Mobility x 1970 Cohort				0.359***	0.292***
				(0.069)	(0.070)
East Asia and Pacific					0.141***
					(0.044)
Europe and Central Asia					0.042*
					(0.024)
Latin America and Caribbean					0.375***
					(0.027)
Middle East and North Africa					0.262***
					(0.062)
South Asia					0.362***
					(0.068)
Sub-Saharan Africa					0.160***
					(0.033)
Constant	0.838***	0.853***	0.771***	1.080***	0.843***
	(0.056)	(0.055)	(0.043)	(0.049)	(0.049)
Observations	549	549	549	549	549
R-squared	0.156		0.213	0.334	
Number of countrycode	142	142	142	142	142
Country FE	Yes	No	Yes	Yes	No

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

† *Notes*: Asterisks denote statistical significance (*** p < 0.01, ** p < 0.05, * p < 0.1). Heteroskedasticity-robust standard errors are reported in parentheses. The dependent variable, proxying for relative mobility, is the intergenerational regression coefficient (IGP), also known as β. Independent variables are listed in rows. Absolute mobility is proxied by ABSCAT, a threshold variable that equals 1 if the child's ordinal education category surpasses the parent's and 0 otherwise. The base birth cohort is 1980 when cohort variables are introduced. North America serves as the base region when regional variables are included. For IGP estimation, parental education is the average of both parents' years of schooling. IGP calculations include both sons and daughters. Refer to the Data section for further details. The sample comprises 549 observations rather than $5 \times 142 = 710$ due to unbalanced panel structure. Not all countries have mobility estimates for all five cohorts, primarily because earlier cohorts lack sufficient sample sizes in GDIM data for some countries, particularly those with delayed survey implementation. This yields approximately 77% of potential country-cohort combinations.

Models 1 and 2 establish the baseline relationship between absolute and relative mobility. Model 1 employs fixed-effects regression with country fixed effects, while Model 2 uses random-effects generalized least squares (GLS) regression. Both specifications reveal a strong, negative correlation between absolute mobility and intergenerational persistence: $\hat{\beta} = -0.548$ (Model 1) and $\hat{\beta} = -0.573$ (Model 2), both significant at $p < 0.01$. The similarity between specifications indicates this relationship reflects genuine patterns rather than artifacts of country-specific factors. Figure AB.2 illustrates this negative relationship visually, with the fitted line's slope closely approximating the absolute mobility coefficients in Models 1 and 2.

This finding alone cannot adjudicate between compositional (MMI/EMI) and disruption theories, as both predict negative correlations. A 10-percentage-point increase in absolute mobility associates with approximately 5.5-percentage-point decline in IGP. However, this correlation could reflect either shifting marginal distributions through hierarchical diffusion or genuine disruption of stratification mechanisms.

Model 3 adds cohort fixed effects to isolate temporal trends in relative mobility beyond changes captured by absolute mobility. The results reveal significant cohort fixed effects: compared to the 1980s cohort (base category), IGP was higher by 0.086 (1940s), 0.075 (1950s), and 0.041 (1960s) percentage points, all significant at $p < 0.01$. This indicates relative mobility exhibits a declining trend across birth cohorts even when controlling for absolute mobility levels and country-specific factors. Importantly, the absolute mobility coefficient remains negative and significant ($\hat{\beta} = -0.498$, $p < 0.01$), confirming the robustness of the core relationship.

Model 4 provides the key test for discriminating between competing theoretical frameworks by introducing cohort-absolute mobility interactions. As developed in Section 3, Maximally Maintained Inequality predicts positive interactions (relationship weakens as systems saturate), while disruption theories predict negative interactions (relationship strengthens as policies become equity-focused). Because IGP enters as the dependent variable (higher values indicate lower mobility), positive θ coefficients indicate the negative relationship strengthens over time.

The results strongly support disruption theories. All four interaction coefficients are positive and highly significant: $\hat{\theta}_{1940} = 0.808$ ($p < 0.01$), $\hat{\theta}_{1950} = 0.507$ ($p < 0.01$), $\hat{\theta}_{1960} = 0.335$ ($p < 0.01$), $\hat{\theta}_{1970} = 0.359$ ($p < 0.01$). The absolute mobility coefficient for the 1980s cohort (base category) is -1.066, nearly double the unconditional effect. For the 1940s cohort, the net effect is $-1.066 + 0.808 = -0.258$, considerably weaker. This represents more than fourfold strengthening from 1940s to 1980s cohorts.

This temporal pattern is difficult to reconcile with MMI's prediction that hierarchical diffusion produces strongest effects during early expansion phases. Instead, the strengthening relationship suggests contemporary expansion increasingly associates with weakened persistence in ways consistent with equity-focused policy environments. Figure AB.3 illustrates these temporal dynamics visually: the negative relationship clearly strengthens from 1940s through 1980s cohorts, with fitted lines becoming progressively steeper.

Model 5 replaces country fixed effects with region fixed effects, confirming robustness while revealing regional heterogeneity. The interaction pattern persists with similar magnitudes and significance levels. Regional fixed effects show that, compared to North America (base), several regions exhibit higher persistence: Latin America and Caribbean (+0.375, $p < 0.01$), South Asia (+0.362, $p < 0.01$), Middle East and North Africa (+0.262, $p < 0.01$), Sub-Saharan Africa (+0.160, $p < 0.01$), and East Asia and Pacific (+0.141, $p < 0.01$). Europe and Central Asia shows a modest positive coefficient (+0.042, $p < 0.1$).

Figure AB.4 illustrates regional heterogeneity visually. East Asia and Pacific demonstrates particularly steep negative slopes, consistent with patterns combining absolute gains and declining persistence. Latin America and Caribbean exhibits flatter slopes despite moderate absolute mobility, consistent with stratified expansion. Nordic countries (within Europe and Central Asia) cluster in the high-absolute-mobility, low-persistence quadrant, while South Asian countries show lower absolute mobility with higher persistence.

The online appendix verifies robustness to parental education specification using father's and mother's education separately. Both reproduce core results: negative absolute mobility coefficients, significant cohort fixed effects, and positive cohort interactions. Estimates using father's education are moderately smaller in Model 1 ($\hat{\beta} = -0.352$, $p < 0.01$) than average parental education, while mother's education estimates are comparable ($\hat{\beta} = -0.442$, $p < 0.01$). Cohort interaction patterns remain qualitatively identical across specifications, indicating findings reflect robust phenomena rather than measurement-specific artifacts.

5.3 Gender Heterogeneity: Differential Mechanisms

Table 4: Estimation Results for Average Parental Education & Daughters Only

VARIABLES	(1) Model 1	(2) Model 2	(3) Model 3	(4) Model 4	(5) Model 5
Absolute Mobility	-0.403***	-0.450***	-0.387***	-0.981***	-0.801***
	(0.103)	(0.070)	(0.087)	(0.119)	(0.088)
1940 Cohort			0.014	-0.428***	-0.352***
			(0.029)	(0.097)	(0.091)
1950 Cohort			0.030	-0.265***	-0.201***
			(0.022)	(0.081)	(0.076)
1960 Cohort			0.019	-0.168***	-0.120**
			(0.015)	(0.055)	(0.052)
1970 Cohort			0.002	-0.172***	-0.135***
			(0.014)	(0.046)	(0.045)
Absolute Mobility x 1940 Cohort				0.876***	0.745***
				(0.157)	(0.148)
Absolute Mobility x 1950 Cohort				0.532***	0.418***
				(0.127)	(0.118)
Absolute Mobility x 1960 Cohort				0.330***	0.240***
				(0.089)	(0.082)
Absolute Mobility x 1970 Cohort				0.299***	0.233***
				(0.071)	(0.069)
East Asia and Pacific					0.126***
					(0.044)
Europe and Central Asia					0.053**
					(0.025)
Latin America and Caribbean					0.368***
					(0.031)
Middle East and North Africa					0.261***
					(0.069)
South Asia					0.259***
					(0.058)
Sub-Saharan Africa					0.137***
					(0.041)
Constant	0.750***	0.779***	0.730***	1.056***	0.829***
	(0.052)	(0.044)	(0.042)	(0.069)	(0.059)
Observations	544	544	544	544	544
R-squared	0.089		0.094	0.272	
Number of countrycode	141	141	141	141	141
Country FE	Yes	No	Yes	Yes	No

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

† *Notes*: Asterisks denote statistical significance (*** p < 0.01, ** p < 0.05, * p < 0.1). Heteroskedasticity-robust standard errors are reported in parentheses. The dependent variable proxying for relative mobility is the intergenerational regression coefficient (`IGP`), also known as β. Independent variables are listed in rows. Absolute mobility is proxied by `ABSCAT`, a threshold variable that equals 1 if the child's ordinal education category surpasses the parent's and 0 otherwise. The base birth cohort is 1980 when cohort variables are introduced. North America serves as the base region when regional variables are included. For `IGP` estimation, parental education is the average of both parents' years of schooling. `IGP` calculations include daughters only. Refer to the Data section for further details.

Table 4 examines whether absolute-relative mobility relationships differ by gender, as predicted by the theoretical framework, presenting results for daughters only (results for sons only are available in the online appendix).

For daughters, the patterns differ markedly from the pooled analysis. While Models 1 and 2 show significant negative relationships ($\hat{\beta} = -0.403$ and -0.450, both $p < 0.01$), Model 3 reveals a striking finding: cohort fixed effects are all statistically insignificant. This contrasts sharply with the secular trends observed for sons (see appendix) and suggests daughters' mobility patterns do not exhibit the same unconditional temporal trends.

However, Model 4 reveals this null result stems from omitted interaction terms. Once cohort-absolute mobility interactions are included, they are highly significant: $\hat{\theta}_{1940} = 0.876$ ($p < 0.01$), $\hat{\theta}_{1950} = 0.532$ ($p < 0.01$), $\hat{\theta}_{1960} = 0.330$ ($p < 0.01$), $\hat{\theta}_{1970} = 0.299$ ($p < 0.01$). The absolute mobility coefficient for 1980s daughters is -0.981, comparable to sons. The interaction magnitudes actually exceed those for sons, indicating even stronger temporal strengthening of the absolute-relative relationship for daughters.

Model 5 confirms regional variation for daughters, though with different patterns than sons. While

Latin America and Caribbean (+0.368, $p < 0.01$) and Middle East and North Africa (+0.261, $p < 0.01$) show high persistence comparable to sons, South Asia (+0.259, $p < 0.01$) shows lower persistence for daughters than sons, consistent with targeted policies benefiting daughters in historically unequal regions.

The gender-specific variations align with theoretical predictions developed in Section 3. The insignificant unconditional cohort effects for daughters (Model 3) but highly significant interactions (Model 4) suggest daughters' mobility operates through different mechanisms than sons'. For daughters in historically gender-unequal contexts, the strong cohort interactions may reflect targeted disruption of traditional barriers through girls' education campaigns and affirmative policies. For sons, mobility patterns reflect both secular trends (significant cohort fixed effects) and compositional expansion within existing hierarchies.

5.4 Intergenerational Correlation: Variance Dynamics Matter

Table 5: Estimation Results for Average Parental Education & Both Sons and Daughters

	(1)	(2)	(3)	(4)	(5)
VARIABLES	Model 1	Model 2	Model 3	Model 4	Model 5
Absolute Mobility	-0.034	-0.009	-0.028	-0.331***	-0.257***
	(0.049)	(0.041)	(0.045)	(0.060)	(0.049)
1940 Cohort			-0.021	-0.227***	-0.210***
			(0.015)	(0.049)	(0.047)
1950 Cohort			-0.038***	-0.212***	-0.204***
			(0.013)	(0.039)	(0.039)
1960 Cohort			-0.037***	-0.174***	-0.174***
			(0.010)	(0.029)	(0.028)
1970 Cohort			-0.013	-0.118***	-0.119***
			(0.008)	(0.023)	(0.022)
Absolute Mobility x 1940 Cohort				0.388***	0.357***
				(0.081)	(0.077)
Absolute Mobility x 1950 Cohort				0.320***	0.298***
				(0.063)	(0.062)
Absolute Mobility x 1960 Cohort				0.253***	0.246***
				(0.048)	(0.047)
Absolute Mobility x 1970 Cohort				0.190***	0.187***
				(0.039)	(0.038)
East Asia and Pacific					-0.011
					(0.057)
Europe and Central Asia					0.018
					(0.055)
Latin America and Caribbean					0.179***
					(0.056)
Middle East and North Africa					-0.013
					(0.064)
South Asia					0.043
					(0.066)
Sub-Saharan Africa					-0.003
					(0.057)
Constant	0.449***	0.441***	0.466***	0.630***	0.574***
	(0.026)	(0.027)	(0.023)	(0.034)	(0.060)
Observations	549	549	549	549	549
R-squared	0.002		0.046	0.156	
Number of countrycode	142	142	142	142	142
Country FE	Yes	No	Yes	Yes	No

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

† *Notes*: Asterisks denote statistical significance (*** p < 0.01, ** p < 0.05, * p < 0.1). Heteroskedasticity-robust standard errors are reported in parentheses. The dependent variable proxying for relative mobility is the intergenerational correlation coefficient (`IGC`), also known as ρ. Independent variables are listed in rows. Absolute mobility is proxied by `ABSCAT`, a threshold variable that equals 1 if the child's ordinal education category surpasses the parent's and 0 otherwise. The base birth cohort is 1980 when cohort variables are introduced. North America serves as the base region when regional variables are included. For `IGC` estimation, parental education is the average of both parents' years of schooling. `IGC` calculations include both sons and daughters. Refer to the Data section for further details.

Table 5 presents results using the intergenerational correlation coefficient (`IGC`, ρ) as the measure of relative mobility. By adjusting for variance changes across generations, `IGC` isolates association strength from compositional effects. Comparing `IGC` patterns with `IGP` results tests whether observed relationships re-

flect variance expansion (supporting compositional hypotheses) versus genuine changes in association strength (supporting disruption hypotheses).

Models 1 and 2 reveal a striking contrast with IGP results: the intergenerational correlation coefficient shows no significant association with absolute mobility. The coefficients are small and statistically insignificant ($\hat{\beta} = -0.034$, $p > 0.1$ in Model 1; $\hat{\beta} = -0.009$, $p > 0.1$ in Model 2). Model 3 shows significant cohort fixed effects, all negative relative to the 1980s cohort, but the absolute mobility coefficient remains insignificant ($\hat{\beta} = -0.028$, $p > 0.1$).

However, Model 4 negates this null result by including cohort-absolute mobility interactions. The absolute mobility coefficient for the 1980s cohort becomes large and highly significant ($\hat{\beta} = -0.331$, $p < 0.01$), and all four interaction terms are positive and significant: $\hat{\theta}_{1940} = 0.388$ ($p < 0.01$), $\hat{\theta}_{1950} = 0.320$ ($p < 0.01$), $\hat{\theta}_{1960} = 0.253$ ($p < 0.01$), $\hat{\theta}_{1970} = 0.190$ ($p < 0.01$). This indicates the relationship between IGC and absolute mobility strengthens over time, similar to IGP patterns.

Model 5 reveals limited region-specific variation in IGC levels. Only Latin America and Caribbean shows higher IGC than North America (+0.179, $p < 0.01$), while other regions show no significant differences. This contrasts with IGP results where multiple regions exhibited heterogeneity. Yet the cohort interaction pattern remains robust with similar magnitudes and significance levels.

These findings have important implications for interpreting mobility dynamics. The null IGC-absolute mobility relationship in Models 1-3 does not indicate absence of association strength changes but rather reflects complex variance dynamics during educational expansion. When children's educational variance expands faster than parents' (common during rapid system expansion), this affects ρ independently of β changes. The significant negative relationship emerging in Model 4 indicates that, once temporal variance dynamics are properly modeled through cohort interactions, IGC shows patterns consistent with genuine association changes rather than pure compositional effects.

This supports disruption theories over compositional theories. If stratification operated primarily through unmeasured quality dimensions while quantity expansion proceeded hierarchically, IGC should show weak relationships even with cohort controls. Instead, the strong negative relationship (when properly specified) indicates expansion affects both composition and association strength.

5.5 Expected Rank: Evidence for Daughters

While no robust findings emerge for descendants of both genders or sons only when using expected rank (EXRANK) as the relative mobility measure, the analysis of daughters yields consistent and significant results. Table 6 reports expected ranks of daughters against average parental education.

Table 6: Estimation Results for Average Parental Education & Daughters Only

VARIABLES	(1) Model 1	(2) Model 2	(3) Model 3	(4) Model 4	(5) Model 5
Absolute Mobility	3.788*	3.498**	1.069	12.639***	10.632***
	(2.008)	(1.466)	(1.942)	(2.776)	(2.235)
1940 Cohort			-1.974***	6.558***	5.720***
			(0.628)	(1.347)	(1.217)
1950 Cohort			-1.598***	4.721***	4.098***
			(0.533)	(1.339)	(1.221)
1960 Cohort			-0.456	3.732***	3.295***
			(0.430)	(1.050)	(0.963)
1970 Cohort			-0.198	1.473*	1.099
			(0.347)	(0.806)	(0.766)
Absolute Mobility x 1940 Cohort				-16.852***	-15.580***
				(2.450)	(2.275)
Absolute Mobility x 1950 Cohort				-11.522***	-10.588***
				(2.414)	(2.234)
Absolute Mobility x 1960 Cohort				-7.449***	-6.815***
				(2.110)	(1.964)
Absolute Mobility x 1970 Cohort				-2.520*	-2.033
				(1.496)	(1.422)
East Asia and Pacific					0.048
					(1.439)
Europe and Central Asia					1.223
					(1.264)
Latin America and Caribbean					-6.662***
					(1.508)
Middle East and North Africa					-1.204
					(1.633)
South Asia					-1.398
					(2.498)
Sub-Saharan Africa					0.290
					(1.548)
Constant	37.369***	37.658***	39.505***	33.164***	34.496***
	(1.010)	(0.944)	(1.030)	(1.493)	(1.811)
Observations	544	544	544	544	544
R-squared	0.020		0.081	0.276	
Number of countrycode	141	141	141	141	141
Country FE	Yes	No	Yes	Yes	No

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

† *Notes*: Asterisks denote statistical significance (*** p < 0.01, ** p < 0.05, * p < 0.1). Heteroskedasticity-robust standard errors are reported in parentheses. The dependent variable proxying for relative mobility is the expected rank (`EXRANK`). Independent variables are listed in rows. Absolute mobility is proxied by `ABSCAT`, a threshold variable that equals 1 if the child's ordinal education category surpasses the parent's and 0 otherwise. The base birth cohort is 1980 when cohort variables are introduced. North America serves as the base region when regional variables are included. For `EXRANK` estimation, parental education is the average of both parents' years of schooling. `EXRANK` calculations include daughters only. Refer to the Data section for further details.

Models 1 and 2 show positive coefficients ($\hat{\beta} = 3.788$, $p < 0.1$ in Model 1; $\hat{\beta} = 3.498$, $p < 0.05$ in Model 2), indicating daughters' expected ranks are higher in environments with higher average ordinal upward mobility. Model 3 shows this relationship disappears when time fixed effects are included without interactions ($\hat{\beta} = 1.069$, $p > 0.1$).

Model 4 reveals this null result stems from omitted interaction terms. The absolute mobility coefficient for 1980s daughters is large and positive ($\hat{\beta} = 12.639$, $p < 0.01$), while all interaction terms are negative and highly significant: $\hat{\theta}_{1940} = -16.852$ ($p < 0.01$), $\hat{\theta}_{1950} = -11.522$ ($p < 0.01$), $\hat{\theta}_{1960} = -7.449$ ($p < 0.01$), $\hat{\theta}_{1970} = -2.520$ ($p < 0.1$). Because `EXRANK` is scaled such that higher values indicate better mobility (opposite of `IGP`), negative interactions indicate the positive relationship strengthens over time, consistent with `IGP` and `IGC` patterns.

Model 5 confirms robustness with regional controls. The pattern persists, with Latin America and Caribbean showing lower expected ranks (-6.662, $p < 0.01$), consistent with high persistence in that region. The cohort interactions remain large and significant.

The `EXRANK` results for daughters provide additional evidence for targeted disruption mechanisms. Expected rank specifically captures whether expansion benefits children from disadvantaged backgrounds (bottom half of parental distribution). The strong positive relationship for 1980s daughters indicates recent expansion benefits daughters from lower-educated families, with this benefit weakening in earlier

cohorts. This pattern is difficult to reconcile with hierarchical diffusion and instead suggests targeted policies increasingly reach disadvantaged populations in recent cohorts.

5.6 Synthesis of Findings

Analysis across multiple mobility measures reveals four insights. First, measurement choice matters. IGP exhibits robust unconditional negative correlation with absolute mobility; IGC requires explicit temporal modeling through cohort interactions; EXRANK reveals whether expansion benefits disadvantaged populations. Single-measure analyses risk missing important dimensions.

Second, temporal dynamics are essential and provide the key evidence discriminating between competing theories. Cohort-absolute mobility interactions emerge as consistently significant across measures and subpopulations. The positive interactions for IGP (higher IGP values indicate lower mobility, so positive interactions mean strengthening negative relationships) are difficult to reconcile with Maximally Maintained Inequality's prediction that hierarchical diffusion produces strongest effects in early phases. Instead, the nearly fourfold strengthening from 1940s to 1980s cohorts aligns with disruption theories emphasizing increasingly equity-focused policy environments.

Third, gender-differentiated patterns provide additional evidence for disruption mechanisms. Daughters in recent cohorts show patterns distinct from sons: insignificant unconditional cohort effects but highly significant interactions, indicating daughters' mobility operates through different mechanisms. The strong EXRANK results for 1980s daughters suggest expansion increasingly benefits daughters from disadvantaged backgrounds, consistent with targeted policies disrupting traditional barriers.

Fourth, regional institutional variation moderates relationships as predicted. Latin America and Caribbean consistently exhibits high persistence alongside moderate absolute mobility; East Asia and Pacific shows strong absolute mobility with declining persistence; South Asia demonstrates different patterns for sons versus daughters, suggesting targeted gender policies. These regional patterns underscore how institutional contexts shape whether expansion translates into equalization.

Together, these findings demonstrate that absolute and relative educational mobility can move together under certain conditions. The conditional nature of this relationship (strengthening over cohorts, differing by gender, varying across regions) reveals that expansion's effects depend on policy design, institutional arrangements, and whether expansion targets previously excluded populations versus expanding access within existing hierarchies.

6 Conclusions

This paper examines whether educational expansion is associated with weaker intergenerational persistence or whether absolute gains in schooling can coexist with stable relative stratification. Using data spanning more than 140 countries and five birth cohorts, I analyze how absolute and relative educational mobility covary across a wide range of institutional settings.

Three broad findings emerge. First, there is a robust negative association between IGP and absolute upward mobility in the baseline specifications, and this association is stronger for later cohorts than for earlier ones. This temporal pattern is difficult to reconcile with a simple version of hierarchical diffusion in which the strongest relationship should appear early and then weaken as systems saturate. At the same time, the analysis remains observational, so these results should be interpreted as descriptive regularities rather than as evidence that educational expansion itself causes weaker persistence.

Second, the IGC results suggest that temporal structure matters for inference. In simpler specifications, IGC is only weakly related to absolute mobility, but once cohort interactions are introduced the association becomes more clearly negative. This pattern is consistent with the idea that changes in educational variance across generations can obscure relationships unless temporal heterogeneity is modeled explicitly.

Third, the results reveal meaningful heterogeneity across subpopulations and measures. Daughters show a different pattern from sons, especially in the interaction specifications, and the EXRANK results suggest that more recent expansion may be more strongly associated with gains for daughters from less educated family backgrounds. Regional differences further indicate that similar levels of expansion can be associated with different levels of persistence across institutional settings.

Taken together, the findings establish a set of robust descriptive regularities. The negative association between absolute mobility and intergenerational persistence remains visible across clustered, trend-adjusted, correlated-random-effects, and precision-weighted specifications. These patterns are more consistent with disruption-oriented interpretations than with accounts that treat absolute and relative mobility as largely disconnected, but they do not by themselves identify the mechanisms linking the two.

This analysis contributes to mobility research in three ways. First, it shows empirically that absolute and relative educational mobility need not systematically diverge. Second, it highlights the importance of cohort, gender, and regional heterogeneity in shaping how these dimensions relate. Third, it shows that conclusions depend in part on the mobility measure used, underscoring the value of examining persistence, correlation, and rank-based outcomes together.

From a methodological perspective, the results suggest that cohort interactions and robustness checks are useful tools for describing how mobility relationships evolve over time. For policymakers, the findings suggest caution against treating educational expansion as either automatically equalizing or purely compositional. Whether expansion is associated with weaker persistence appears to depend on when it occurs, who benefits, and the institutional context in which it unfolds.

Several limitations warrant acknowledgment. While GDIM provides unprecedented global coverage, it relies on retrospective survey data with potential measurement error and recall bias. The analysis remains correlational, limiting causal inference about mechanisms. Cross-national correlations could reflect multiple mechanisms operating simultaneously, and distinguishing compositional from disruption effects fully requires assumptions the data cannot directly validate. The analysis cannot fully account for quality differentiation within educational levels or other unmeasured dimensions through which advantages may transmit despite quantity equalization.

Future research should exploit policy reforms generating exogenous variation in educational access to establish causal mechanisms. Natural experiments from major disruptions could provide additional evidence about malleability of intergenerational transmission patterns. Developing strategies for measuring quality differentiation directly would strengthen tests of competing theories. Extending the analysis to other dimensions of intergenerational transmission (occupational mobility, income mobility, wealth transmission) would establish whether educational patterns generalize to broader socioeconomic outcomes.

This study's central contribution demonstrates that absolute and relative educational mobility need not systematically diverge: empirical patterns suggest they can move together under certain conditions, though relationships are conditional on temporal dynamics, gender, institutional arrangements, and measurement choices. As developing countries continue experiencing compressed educational expansion, understanding conditions under which quantity expansion translates into genuine equalization grows increasingly important. The findings provide grounds for cautious optimism: educational expansion can potentially serve as an equalizing force, but temporal patterns suggest this occurs primarily when designed to

do so through explicit targeting of disadvantaged populations rather than universal hierarchical diffusion.

Competing Interests

The author has no conflicts of interest to declare. There is no financial interest to report. The author certifies that the submission is original work and is not under review at any other publication.

Ethics

Ethics approval was not required for this study/Not Applicable.

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Author Contributions

Dr. Torul is the sole contributor of the manuscript.

Data Statement

The Global Database on Intergenerational Mobility (GDIM) dataset used in this analysis is provided freely by the World Bank. The author will offer the Stata codes that generate the results in the manuscript, appendix, and the online appendix.

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Appendix

A A Simple Statistical Model Linking Absolute and Relative Mobility

This subsection presents a compact statistical framework linking the mobility measures used in the main text. The purpose is not to estimate a structural model or identify causal mechanisms, but to clarify how educational expansion can increase absolute mobility while leaving relative mobility largely unchanged, or alternatively cause the two to move together.

Let a child's latent years of schooling in country c and cohort t satisfy

$$Y_{ict}^{child} = a_{ct} + b_{ct} Y_{ict}^{parent} + \varepsilon_{ict}, \qquad \varepsilon_{ict} \sim (0, \sigma_{\varepsilon,ct}^2),$$

where a_{ct} captures cohort-country shifts in educational attainment common to all children, b_{ct} captures intergenerational persistence, and ε_{ict} captures idiosyncratic variation. In this formulation, the intergenerational persistence coefficient used in the main text is simply

$$\texttt{IGP}_{ct} = b_{ct}.$$

A decline in b_{ct} therefore corresponds directly to weaker intergenerational persistence.

Let $\sigma_{p,ct}$ and $\sigma_{c,ct}$ denote the standard deviations of parental and child schooling, respectively. The intergenerational correlation coefficient can then be written as

$$\texttt{IGC}_{ct} = \rho_{ct} = b_{ct} \frac{\sigma_{p,ct}}{\sigma_{c,ct}},$$

with

$$\sigma_{c,ct}^2 = b_{ct}^2 \sigma_{p,ct}^2 + \sigma_{\varepsilon,ct}^2.$$

This expression clarifies why IGC can behave differently from IGP. Even when persistence b_{ct} declines, the correlation coefficient may change little if children's educational variance expands relative to parents' variance.

To connect the model to absolute mobility, define a years-based counterpart to upward mobility as

$$\texttt{Abs}_{ct}^{yos} = P\left(Y_{ict}^{child} > Y_{ict}^{parent}\right).$$

Under the linear-normal setup, this probability is

$$\texttt{Abs}_{ct}^{yos} = \Phi\left(\frac{a_{ct} + (b_{ct} - 1)\mu_{p,ct}}{\sqrt{(b_{ct} - 1)^2 \sigma_{p,ct}^2 + \sigma_{\varepsilon,ct}^2}}\right),$$

where $\mu_{p,ct}$ is mean parental schooling and $\Phi(\cdot)$ denotes the standard normal cumulative distribution function. This equation shows that absolute mobility can rise for two distinct reasons: because the intercept a_{ct} increases, shifting the entire schooling distribution upward, or because persistence b_{ct} falls, weakening the dependence of children's attainment on parental attainment.

The ordinal measure used in the paper, ABSCAT, can be represented with schooling thresholds $\tau_1 < \tau_2 < \tau_3 < \tau_4$ defining the five educational categories. Let parental and child educational categories be generated

by these thresholds. Then

$$\texttt{ABSCAT}_{ct} = P(E^{child} > E^{parent} \mid E^{parent} < 5).$$

Equivalently,

$$\texttt{ABSCAT}_{ct} = \sum_{j=1}^{4} \frac{P\left(\tau_{j-1} < Y^{parent} \leq \tau_j,\, Y^{child} > \tau_j\right)}{P\left(Y^{parent} \leq \tau_4\right)}.$$

This makes clear that the paper's empirical measure is a threshold-based version of the same underlying latent process.

The model also helps organize the interpretation of the main results. Suppose educational expansion has two components: a quantitative component X_{ct} that shifts average attainment upward, and an equity-oriented component Q_{ct} that weakens the dependence of child outcomes on parental background. A simple parameterization is

$$a_{ct} = \alpha_0 + \alpha_1 X_{ct},$$

$$b_{ct} = \beta_0 - \beta_1 X_{ct} - \beta_2 Q_{ct},$$

$$\sigma^2_{\varepsilon,ct} = s_0 + s_1 X_{ct}.$$

Three cases follow. First, under purely compositional expansion, $\alpha_1 > 0$ while $\beta_1 \approx \beta_2 \approx 0$; absolute mobility rises mainly because schooling levels increase, but relative mobility changes little. Second, under disruption-oriented expansion, $\alpha_1 > 0$ and $\beta_2 > 0$; absolute mobility rises while persistence falls, so absolute and relative mobility move together. Third, when variance changes are large, $\sigma^2_{\varepsilon,ct}$ can rise enough that IGC appears weakly related to absolute mobility even when IGP declines.

This simple framework is consistent with the empirical patterns in the main text. The strong negative relationship between IGP and absolute mobility is consistent with cases in which upward shifts in schooling are accompanied by weaker persistence. The weaker unconditional relationship for IGC, followed by a clearer relationship once cohort interactions are introduced, is consistent with variance dynamics that differ across cohorts. Finally, the stronger patterns for daughters and later cohorts can be interpreted as settings in which the equity-related component of expansion is more pronounced. The model does not identify which mechanisms generated these changes, but it provides a transparent language for interpreting why absolute and relative mobility sometimes move together and sometimes do not.

B Educational and Income Mobility Across Countries

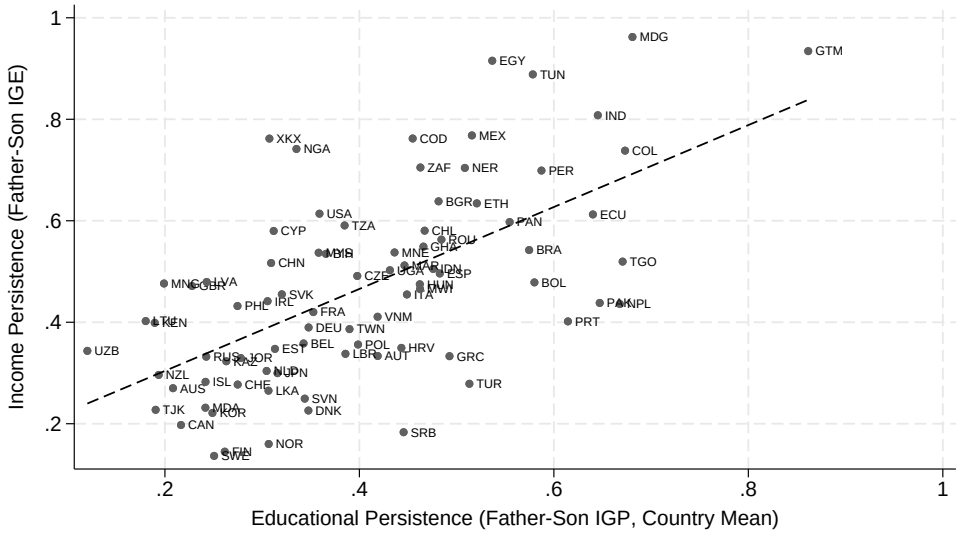
To assess the broader socioeconomic relevance of the educational mobility metrics used in the main text, this appendix section presents a descriptive external-validity exercise comparing educational persistence with intergenerational income persistence. For this cross-national comparison, I utilize the World Bank's Global Database on Intergenerational Income Mobility.

Because the income mobility database provides cross-sectional estimates restricted exclusively to father-son pairings, it cannot replicate the cohort-by-gender panel architecture that forms the core of this paper's main analysis. To ensure demographic comparability, I restrict the educational GDIM comparator to the father's education and sons only sample. Furthermore, to match the cross-sectional structure of the income data, the primary specification constructs a country-level cross-section by averaging the educational mobility measures across all available cohorts. A secondary sensitivity check restricting the educational data to the most recent (1980s) birth cohort yields substantively similar descriptive patterns.

The purpose of this supplementary exercise is to examine whether countries displaying higher educa-

tional persistence also tend to display higher intergenerational income persistence within the overlapping sample. Table AB.1 reports the results of a simple bivariate regression of the income persistence measure on the educational persistence coefficient (`IGP`), and Figure AB.1 plots the cross-country association. This comparison is strictly descriptive; it explores the cross-dataset consistency of stratification measures rather than identifying causal links between educational expansion and labor market returns.

Figure AB.1: Cross-Country Association: Educational vs. Income Persistence (Father-Son)



† *Notes*: This figure plots the descriptive overlap between educational and income persistence for the 82 matching countries. The horizontal axis denotes the father-son intergenerational educational persistence coefficient (`IGP`), averaged across all available cohorts in the GDIM dataset. The vertical axis denotes the father-son intergenerational income elasticity (`IGE`) from the World Bank Global Database on Intergenerational Income Mobility. Dots represent country ISO3 codes and the solid line represents the fitted linear association.

Table AB.1: Descriptive Regression: Income Persistence on Educational Persistence

VARIABLES	(1) OLS
Educational Persistence (Father-Son IGP, Country Mean)	0.807***
	(0.116)
Constant	0.143***
	(0.046)
Observations	82
R-squared	0.387
Sample	Father-Son Only
Aggregation	Country Means

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

† *Notes*: Asterisks denote statistical significance (*** p < 0.01, ** p < 0.05, * p < 0.1). Heteroskedasticity-robust standard errors are reported in parentheses. The dependent variable is the father-son intergenerational income elasticity (`IGE`). The independent variable is the father-son intergenerational educational persistence coefficient (`IGP`), averaged across all available cohorts per country.

Table AB.2: Detailed Descriptive Statistics

	Cohort (which decade individuals are born in, by first year of decade)					
	1940	1950	1960	1970	1980	Total
N	1,186 (17.6%)	1,259 (18.7%)	1,275 (19.0%)	1,271 (18.9%)	1,731 (25.8%)	6,722 (100.0%)
Intergenerational Persistence Coefficient	0.516 (0.295)	0.480 (0.266)	0.448 (0.220)	0.440 (0.182)	0.444 (0.184)	0.463 (0.231)
Intergenerational Correlation Coefficient	0.415 (0.171)	0.393 (0.147)	0.391 (0.117)	0.413 (0.111)	0.427 (0.114)	0.409 (0.133)
Expected Rank Born to Bottom-Half Parent	42.969 (7.050)	42.831 (6.911)	42.632 (6.295)	42.003 (5.743)	40.975 (5.306)	42.183 (6.269)
Pr(C>P)	0.479 (0.262)	0.542 (0.238)	0.565 (0.213)	0.545 (0.198)	0.543 (0.188)	0.536 (0.220)
Pr(Years(C)>Years(P))	0.578 (0.258)	0.633 (0.223)	0.650 (0.196)	0.626 (0.183)	0.618 (0.175)	0.621 (0.208)

† *Notes*: This table reports the descriptive statistics of the GDIM dataset. Columns denote birth cohorts by decade. `IGP` and `IGC` stand for intergenerational regression and correlation coefficients, respectively. `EXRANK` stands for the expected rank of a child born to a parent from the bottom half of the education distribution. `ABSCAT` stands for the dummy variable that takes the value 1 if child's ordinal education category surpasses that of parent, and 0 otherwise; conditional on parent not being on the highest ordinal (*tertiary*) education category. The numbers in parentheses in the number of observation rows refer to their respective frequencies. The remaining numbers in parentheses next to the mean values denote standard deviations.

Table AB.3: Estimation Results for Father's Parental Education & Both Sons and Daughters

	(1)	(2)	(3)	(4)	(5)
VARIABLES	Model 1	Model 2	Model 3	Model 4	Model 5
Absolute Mobility	-0.352***	-0.357***	-0.293***	-0.712***	-0.599***
	(0.082)	(0.060)	(0.073)	(0.073)	(0.059)
1940 Cohort			0.054***	-0.248***	-0.201***
			(0.016)	(0.061)	(0.061)
1950 Cohort			0.035**	-0.172***	-0.134***
			(0.014)	(0.047)	(0.048)
1960 Cohort			0.021**	-0.099***	-0.071**
			(0.010)	(0.031)	(0.032)
1970 Cohort			0.012	-0.120***	-0.098***
			(0.009)	(0.031)	(0.031)
Absolute Mobility x 1940 Cohort				0.561***	0.482***
				(0.101)	(0.100)
Absolute Mobility x 1950 Cohort				0.371***	0.301***
				(0.075)	(0.075)
Absolute Mobility x 1960 Cohort				0.216***	0.163***
				(0.052)	(0.052)
Absolute Mobility x 1970 Cohort				0.232***	0.192***
				(0.049)	(0.050)
East Asia and Pacific					0.084***
					(0.031)
Europe and Central Asia					0.046**
					(0.022)
Latin America and Caribbean					0.303***
					(0.025)
Middle East and North Africa					0.123***
					(0.034)
South Asia					0.160***
					(0.040)
Sub-Saharan Africa					0.085***
					(0.028)
Constant	0.608***	0.612***	0.554***	0.787***	0.634***
	(0.044)	(0.038)	(0.037)	(0.040)	(0.040)
Observations	559	559	559	559	559
R-squared	0.123		0.159	0.304	
Number of countrycode	144	144	144	144	144
Country FE	Yes	No	Yes	Yes	No

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

† *Notes*: Asterisks denote statistical significance (*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$). Heteroskedasticity-robust standard errors are reported in parentheses. The dependent variable proxying for relative mobility is the intergenerational regression coefficient (`IGP`), also known as β. Independent variables are listed in rows. Absolute mobility is proxied by `ABSCAT`, a threshold variable that equals 1 if the child's ordinal education category surpasses the parent's and 0 otherwise. The base birth cohort is 1980 when cohort variables are introduced. North America serves as the base region when regional variables are included. For `IGP` estimation, parental education is the father's years of schooling. `IGP` calculations include both sons and daughters. Refer to the Data section for further details.

Table AB.4: Estimation Results for Mother's Parental Education & Both Sons and Daughters

VARIABLES	(1) Model 1	(2) Model 2	(3) Model 3	(4) Model 4	(5) Model 5
Absolute Mobility	-0.442***	-0.454***	-0.408***	-1.074***	-0.881***
	(0.093)	(0.079)	(0.080)	(0.095)	(0.081)
1940 Cohort			0.064**	-0.459***	-0.379***
			(0.026)	(0.100)	(0.097)
1950 Cohort			0.066***	-0.272***	-0.208**
			(0.021)	(0.084)	(0.088)
1960 Cohort			0.036**	-0.181***	-0.141**
			(0.014)	(0.058)	(0.058)
1970 Cohort			0.013	-0.193***	-0.169***
			(0.011)	(0.045)	(0.045)
Absolute Mobility x 1940 Cohort				0.903***	0.771***
				(0.147)	(0.141)
Absolute Mobility x 1950 Cohort				0.568***	0.457***
				(0.121)	(0.127)
Absolute Mobility x 1960 Cohort				0.371***	0.297***
				(0.087)	(0.086)
Absolute Mobility x 1970 Cohort				0.343***	0.299***
				(0.067)	(0.067)
East Asia and Pacific					0.137***
					(0.043)
Europe and Central Asia					0.058**
					(0.025)
Latin America and Caribbean					0.374***
					(0.028)
Middle East and North Africa					0.230***
					(0.067)
South Asia					0.378***
					(0.074)
Sub-Saharan Africa					0.106***
					(0.037)
Constant	0.712***	0.721***	0.659***	1.054***	0.811***
	(0.054)	(0.057)	(0.042)	(0.058)	(0.054)
Observations	549	549	549	549	549
R-squared	0.110		0.146	0.312	
Number of countrycode	142	142	142	142	142
Country FE	Yes	No	Yes	Yes	No

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

† *Notes*: Asterisks denote statistical significance (*** p < 0.01, ** p < 0.05, * p < 0.1). Heteroskedasticity-robust standard errors are reported in parentheses. The dependent variable proxying for relative mobility is the intergenerational regression coefficient (`IGP`), also known as β. Independent variables are listed in rows. Absolute mobility is proxied by `ABSCAT`, a threshold variable that equals 1 if the child's ordinal education category surpasses the parent's and 0 otherwise. The base birth cohort is 1980 when cohort variables are introduced. North America serves as the base region when regional variables are included. For `IGP` estimation, parental education is the mother's years of schooling. `IGP` calculations include both sons and daughters. Refer to the Data section for further details.

Table AB.5: Estimation Results for Average Parental Education & Sons Only

VARIABLES	(1) Model 1	(2) Model 2	(3) Model 3	(4) Model 4	(5) Model 5
Absolute Mobility	-0.560***	-0.584***	-0.593***	-1.093***	-0.888***
	(0.104)	(0.082)	(0.095)	(0.093)	(0.084)
1940 Cohort			0.151***	-0.296***	-0.193*
			(0.030)	(0.096)	(0.101)
1950 Cohort			0.125***	-0.136*	-0.052
			(0.024)	(0.080)	(0.089)
1960 Cohort			0.072***	-0.061	-0.013
			(0.017)	(0.063)	(0.065)
1970 Cohort			0.026*	-0.150***	-0.120**
			(0.015)	(0.049)	(0.049)
Absolute Mobility x 1940 Cohort				0.824***	0.637***
				(0.148)	(0.154)
Absolute Mobility x 1950 Cohort				0.483***	0.322**
				(0.122)	(0.133)
Absolute Mobility x 1960 Cohort				0.265**	0.167
				(0.102)	(0.102)
Absolute Mobility x 1970 Cohort				0.327***	0.271***
				(0.079)	(0.077)
East Asia and Pacific					0.175***
					(0.056)
Europe and Central Asia					0.039
					(0.031)
Latin America and Caribbean					0.401***
					(0.035)
Middle East and North Africa					0.280***
					(0.067)
South Asia					0.449***
					(0.082)
Sub-Saharan Africa					0.212***
					(0.036)
Constant	0.853***	0.866***	0.803***	1.074***	0.798***
	(0.058)	(0.058)	(0.046)	(0.047)	(0.050)
Observations	538	538	538	538	538
R-squared	0.154		0.286	0.377	
Number of countrycode	141	141	141	141	141
Country FE	Yes	No	Yes	Yes	No

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

† *Notes*: Asterisks denote statistical significance (*** p < 0.01, ** p < 0.05, * p < 0.1). Heteroskedasticity-robust standard errors are reported in parentheses. The dependent variable proxying for relative mobility is the intergenerational regression coefficient (`IGP`), also known as β. Independent variables are listed in rows. Absolute mobility is proxied by `ABSCAT`, a threshold variable that equals 1 if the child's ordinal education category surpasses the parent's and 0 otherwise. The base birth cohort is 1980 when cohort variables are introduced. North America serves as the base region when regional variables are included. For `IGP` estimation, parental education is the average of both parents' years of schooling. `IGP` calculations include sons only. Refer to the Data section for further details.

Table AB.6: Estimation Results for Father's Parental Education & Sons Only

VARIABLES	(1) Model 1	(2) Model 2	(3) Model 3	(4) Model 4	(5) Model 5
Absolute Mobility	-0.451***	-0.448***	-0.431***	-0.716***	-0.559***
	(0.108)	(0.080)	(0.096)	(0.067)	(0.061)
1940 Cohort			0.105***	-0.117	-0.038
			(0.022)	(0.108)	(0.115)
1950 Cohort			0.074***	-0.092*	-0.026
			(0.015)	(0.052)	(0.058)
1960 Cohort			0.050***	-0.015	0.024
			(0.012)	(0.052)	(0.051)
1970 Cohort			0.026**	-0.102**	-0.081*
			(0.012)	(0.044)	(0.043)
Absolute Mobility x 1940 Cohort				0.403**	0.262
				(0.167)	(0.179)
Absolute Mobility x 1950 Cohort				0.295***	0.172*
				(0.081)	(0.090)
Absolute Mobility x 1960 Cohort				0.125	0.047
				(0.081)	(0.079)
Absolute Mobility x 1970 Cohort				0.229***	0.188***
				(0.068)	(0.067)
East Asia and Pacific					0.095**
					(0.038)
Europe and Central Asia					0.027
					(0.029)
Latin America and Caribbean					0.300***
					(0.034)
Middle East and North Africa					0.105***
					(0.039)
South Asia					0.192***
					(0.051)
Sub-Saharan Africa					0.106***
					(0.032)
Constant	0.677***	0.672***	0.619***	0.777***	0.601***
	(0.061)	(0.051)	(0.049)	(0.036)	(0.043)
Observations	549	549	549	549	549
R-squared	0.139		0.228	0.268	
Number of countrycode	143	143	143	143	143
Country FE	Yes	No	Yes	Yes	No

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

† *Notes*: Asterisks denote statistical significance (*** p < 0.01, ** p < 0.05, * p < 0.1). Heteroskedasticity-robust standard errors are reported in parentheses. The dependent variable proxying for relative mobility is the intergenerational regression coefficient (`IGP`), also known as β. Independent variables are listed in rows. Absolute mobility is proxied by `ABSCAT`, a threshold variable that equals 1 if the child's ordinal education category surpasses the parent's and 0 otherwise. The base birth cohort is 1980 when cohort variables are introduced. North America serves as the base region when regional variables are included. For `IGP` estimation, parental education is father's years of schooling. `IGP` calculations include sons only. Refer to the Data section for further details.

Table AB.7: Estimation Results for Mother's Parental Education & Sons Only

VARIABLES	(1) Model 1	(2) Model 2	(3) Model 3	(4) Model 4	(5) Model 5
Absolute Mobility	-0.433***	-0.458***	-0.445***	-1.048***	-0.844***
	(0.114)	(0.096)	(0.108)	(0.101)	(0.088)
1940 Cohort			0.098***	-0.414***	-0.324**
			(0.035)	(0.153)	(0.155)
1950 Cohort			0.083***	-0.284***	-0.218**
			(0.025)	(0.087)	(0.097)
1960 Cohort			0.050***	-0.110*	-0.087
			(0.017)	(0.061)	(0.062)
1970 Cohort			0.020	-0.155***	-0.152***
			(0.013)	(0.057)	(0.058)
Absolute Mobility x 1940 Cohort				0.869***	0.712***
				(0.222)	(0.222)
Absolute Mobility x 1950 Cohort				0.617***	0.492***
				(0.121)	(0.134)
Absolute Mobility x 1960 Cohort				0.293***	0.236***
				(0.090)	(0.090)
Absolute Mobility x 1970 Cohort				0.300***	0.287***
				(0.083)	(0.086)
East Asia and Pacific					0.159***
					(0.057)
Europe and Central Asia					0.060*
					(0.036)
Latin America and Caribbean					0.403***
					(0.041)
Middle East and North Africa					0.225***
					(0.066)
South Asia					0.424***
					(0.092)
Sub-Saharan Africa					0.163***
					(0.044)
Constant	0.700***	0.717***	0.661***	1.018***	0.746***
	(0.069)	(0.070)	(0.057)	(0.059)	(0.058)
Observations	539	539	539	539	539
R-squared	0.090		0.143	0.252	
Number of countrycode	141	141	141	141	141
Country FE	Yes	No	Yes	Yes	No

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

† *Notes*: Asterisks denote statistical significance (*** p < 0.01, ** p < 0.05, * p < 0.1). Heteroskedasticity-robust standard errors are reported in parentheses. The dependent variable proxying for relative mobility is the intergenerational regression coefficient (`IGP`), also known as β. Independent variables are listed in rows. Absolute mobility is proxied by `ABSCAT`, a threshold variable that equals 1 if the child's ordinal education category surpasses the parent's and 0 otherwise. The base birth cohort is 1980 when cohort variables are introduced. North America serves as the base region when regional variables are included. For `IGP` estimation, parental education is mother's years of schooling. `IGP` calculations include sons only. Refer to the Data section for further details.

Table AB.8: Estimation Results for Father's Parental Education & Daughters Only

VARIABLES	(1) Model 1	(2) Model 2	(3) Model 3	(4) Model 4	(5) Model 5
Absolute Mobility	-0.217**	-0.244***	-0.212***	-0.671***	-0.539***
	(0.087)	(0.054)	(0.078)	(0.095)	(0.063)
1940 Cohort			0.003	-0.333***	-0.282***
			(0.020)	(0.057)	(0.053)
1950 Cohort			0.005	-0.224***	-0.183***
			(0.017)	(0.054)	(0.051)
1960 Cohort			0.001	-0.143***	-0.112***
			(0.012)	(0.037)	(0.036)
1970 Cohort			0.002	-0.112***	-0.086***
			(0.011)	(0.028)	(0.028)
Absolute Mobility x 1940 Cohort				0.649***	0.568***
				(0.091)	(0.086)
Absolute Mobility x 1950 Cohort				0.401***	0.331***
				(0.083)	(0.078)
Absolute Mobility x 1960 Cohort				0.243***	0.186***
				(0.059)	(0.056)
Absolute Mobility x 1970 Cohort				0.185***	0.141***
				(0.043)	(0.044)
East Asia and Pacific					0.076**
					(0.031)
Europe and Central Asia					0.066***
					(0.021)
Latin America and Caribbean					0.306***
					(0.027)
Middle East and North Africa					0.142***
					(0.037)
South Asia					0.100**
					(0.039)
Sub-Saharan Africa					0.080**
					(0.032)
Constant	0.524***	0.541***	0.520***	0.776***	0.610***
	(0.044)	(0.033)	(0.040)	(0.054)	(0.046)
Observations	554	554	554	554	554
R-squared	0.045		0.045	0.266	
Number of countrycode	143	143	143	143	143
Country FE	Yes	No	Yes	Yes	No

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

† *Notes*: Asterisks denote statistical significance (*** p < 0.01, ** p < 0.05, * p < 0.1). Heteroskedasticity-robust standard errors are reported in parentheses. The dependent variable proxying for relative mobility is the intergenerational regression coefficient (`IGP`), also known as β. Independent variables are listed in rows. Absolute mobility is proxied by `ABSCAT`, a threshold variable that equals 1 if the child's ordinal education category surpasses the parent's and 0 otherwise. The base birth cohort is 1980 when cohort variables are introduced. North America serves as the base region when regional variables are included. For `IGP` estimation, parental education is father's years of schooling. `IGP` calculations include daughters only. Refer to the Data section for further details.

Table AB.9: Estimation Results for Mother's Parental Education & Daughters Only

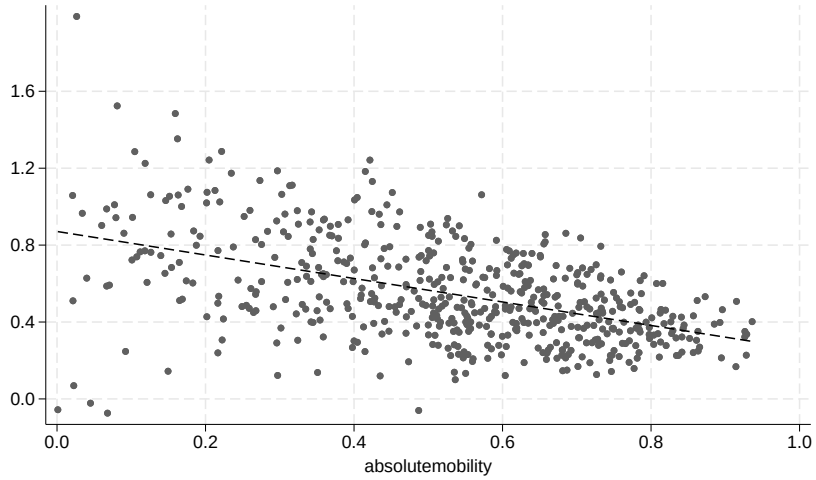
VARIABLES	(1) Model 1	(2) Model 2	(3) Model 3	(4) Model 4	(5) Model 5
Absolute Mobility	-0.478***	-0.470***	-0.438***	-1.020***	-0.788***
	(0.098)	(0.082)	(0.081)	(0.120)	(0.098)
1940 Cohort			0.034	-0.411***	-0.316***
			(0.031)	(0.120)	(0.114)
1950 Cohort			0.050**	-0.236**	-0.160
			(0.023)	(0.098)	(0.099)
1960 Cohort			0.025*	-0.194***	-0.141**
			(0.015)	(0.064)	(0.063)
1970 Cohort			0.005	-0.182***	-0.151***
			(0.011)	(0.043)	(0.046)
Absolute Mobility x 1940 Cohort				0.782***	0.634***
				(0.176)	(0.167)
Absolute Mobility x 1950 Cohort				0.467***	0.339**
				(0.143)	(0.145)
Absolute Mobility x 1960 Cohort				0.356***	0.263***
				(0.094)	(0.091)
Absolute Mobility x 1970 Cohort				0.300***	0.246***
				(0.063)	(0.068)
East Asia and Pacific					0.125***
					(0.042)
Europe and Central Asia					0.053**
					(0.023)
Latin America and Caribbean					0.344***
					(0.028)
Middle East and North Africa					0.221***
					(0.079)
South Asia					0.364***
					(0.075)
Sub-Saharan Africa					0.067
					(0.045)
Constant	0.742***	0.741***	0.699***	1.048***	0.798***
	(0.054)	(0.059)	(0.043)	(0.075)	(0.070)
Observations	544	544	544	544	544
R-squared	0.104		0.115	0.224	
Number of countrycode	141	141	141	141	141
Country FE	Yes	No	Yes	Yes	No

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

† *Notes*: Asterisks denote statistical significance (*** p < 0.01, ** p < 0.05, * p < 0.1). Heteroskedasticity-robust standard errors are reported in parentheses. The dependent variable proxying for relative mobility is the intergenerational regression coefficient (`IGP`), also known as β. Independent variables are listed in rows. Absolute mobility is proxied by `ABSCAT`, a threshold variable that equals 1 if the child's ordinal education category surpasses the parent's and 0 otherwise. The base birth cohort is 1980 when cohort variables are introduced. North America serves as the base region when regional variables are included. For `IGP` estimation, parental education is mother's years of schooling. `IGP` calculations include daughters only. Refer to the Data section for further details.

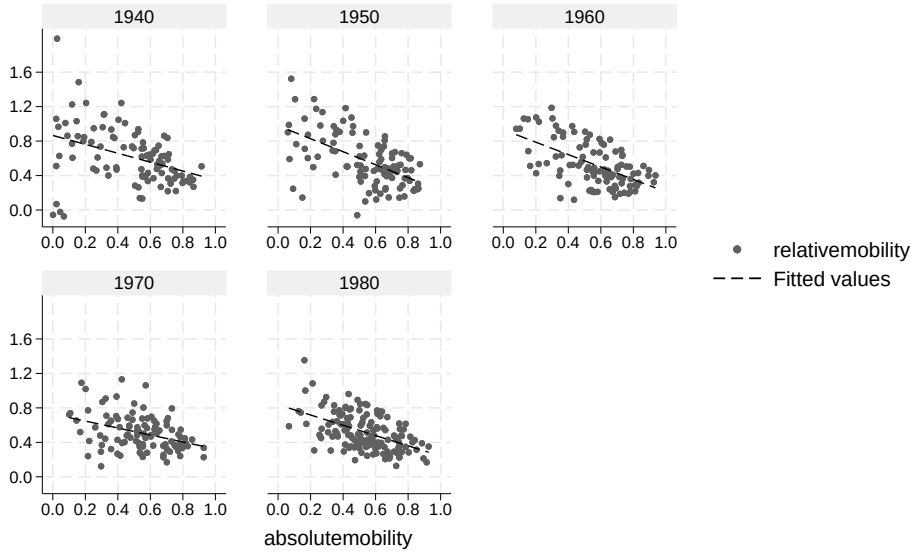
Supplementary Figures

Figure AB.2: IGP (via Average Parental Education & Both Sons and Daughters) vs. ABSCAT



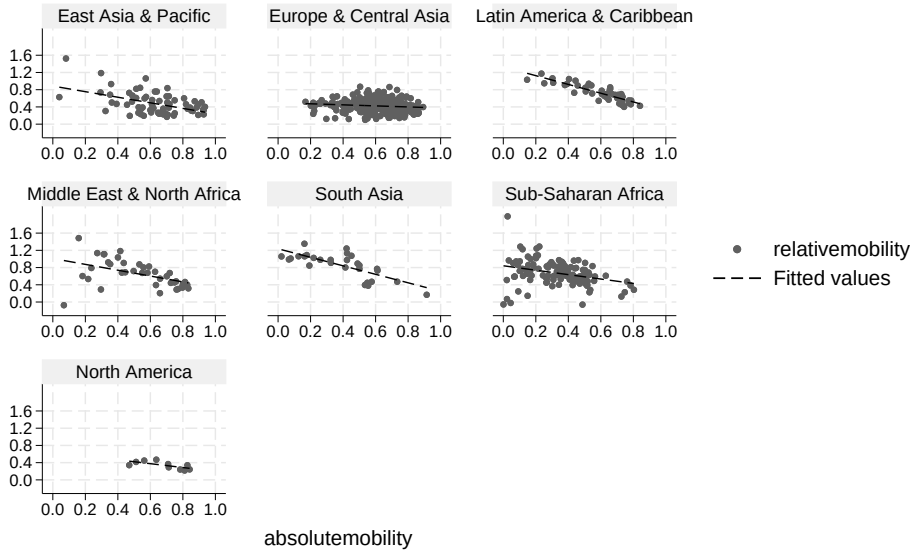
† *Notes*: This figure shows the scatterplot of relative and absolute intergenerational mobility. The horizontal axis denotes ABSCAT, a threshold variable that equals 1 if child's ordinal education category surpasses that of parents, and 0 otherwise. The vertical axis denotes IGP, the intergenerational regression coefficient, i.e., β. IGP is based on the average parental education, and the observations of both sons and daughters. Dots depict actual observations and the dashed line depicts the fitted linear line.

Figure AB.3: IGP (via Average Parental Education & Both Sons and Daughters) vs. ABSCAT by Birth Cohort



† *Notes*: This figure shows the scatterplot of relative and absolute intergenerational mobility by birth cohort. The horizontal axis denotes ABSCAT, a threshold variable that equals 1 if child's ordinal education category surpasses that of parents, and 0 otherwise. The vertical axis denotes IGP, the intergenerational regression coefficient, i.e., β. IGP is based on the average parental education, and the observations of both sons and daughters. Dots depict actual observations and the dashed line depicts the fitted linear line.

Figure AB.4: IGP (via Average Parental Education & Both Sons and Daughters) vs. ABSCAT by Region



† *Notes*: This figure shows the scatterplot of relative and absolute intergenerational mobility by region. The horizontal axis denotes ABSCAT, a threshold variable that equals 1 if child's ordinal education category surpasses that of parents, and 0 otherwise. The vertical axis denotes IGP, the intergenerational regression coefficient, i.e., β. IGP is based on the average parental education, and the observations of both sons and daughters. Dots depict actual observations and the dashed line depicts the fitted linear line.