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## RESEARCH ARTICLE

# Changes in sibling similarity in education among Finnish cohorts born in 1950–89: the contribution of paternal and maternal education

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Among the many social characteristics that run in the family, education is one of the most strongly persistent. The long-term changes in educational reproduction within families and across generations and the gender-specific drivers of these changes remain partially unclear. Using population data for all Finnish siblings and their parents, we assessed the level of and trends in the intergenerational persistence of education among cohorts born between 1950 and 1989. The variance in education shared among siblings was 37% and remained stable over time. Parental education steadily increased its explanatory power in the shared variance, from 30% among cohorts born in the 1950s to 40% among cohorts born in the 1970s and 1980s. The direct contribution of maternal education net of paternal education for sibling similarity more than doubled across cohorts (from 5% in 1950 to 13% in 1989). The direct contribution of paternal education (10–12%) remained stable. Same-gender siblings resembled each other in education more closely than their opposite-gender counterparts. The growing importance of maternal education over time, which surpasses the predictive power of paternal education, demonstrates an important qualitative change in the determinants of educational stratification. The growing importance of mothers' education can plausibly result from the strengthening meritocratic achievement of women in education and the associated increase of women in defining the social position of the family. Incorporating the education of both parents in future analyses of intergenerational reproduction of education will probably be increasingly salient.

**Key words** intergenerational social mobility • education • trends • siblings

### Key messages

- 37% of the variance in education is shared between Finnish siblings.
- This figure was remarkably stable across cohorts born between 1950 and 1989.
- 39% of the sibling similarity can be explained by parental education.
- The contribution of maternal education net of paternal education more than doubled across cohorts.

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## Introduction

The research literature on education reproduction between generations is well-developed in various social scientific fields. This attention is justified from multiple perspectives. Education is an important institution as such and is both a widely used indicator of an individual's socio-economic position (Galobardes et al, 2006; Erola et al, 2016). It is also a powerful variable in predicting many socially desirable outcomes, such as higher level of political participation (Gallego, 2007; Lahtinen et al, 2019b), lower mortality (Mackenbach et al, 2016), and lower delinquency (Aaltonen et al, 2011). It is also considered the main mechanism in both occupation-based social-class reproduction and mobility between generations (Hout and DiPrete, 2006; Bernardi and Ballarino, 2016; Bukodi and Goldthorpe, 2016). In addition, the persistence of education between generations tends to be relatively strong (Mazumder, 2008; Björklund and Jäntti, 2012).

There are two broad approaches to assessing how strongly family background is associated with an individual's education. A traditional and more widely available method is to examine the association of observed parental education with the education of their offspring. This has been done using various empirical strategies, including estimating the intergenerational correlation coefficients or regression coefficients for years of education (Hertz et al, 2008) or performing log-linear analyses of mobility tables (Pfeffer, 2008). An alternative strategy involves using the resemblance in education between siblings to infer the importance of an individual's background indirectly. The advantage of the sibling design is the more comprehensive approach it represents. As formally derived by Solon (1999, see also Björklund and Salvanes, 2011), sibling correlations capture both the intergenerational correlation of the outcome in question (in our case, education) and all other factors shared by siblings that contribute to it. Siblings share a virtually infinite number of factors that are potentially relevant to educational attainment. These factors include, among others, socio-economic and other resources, a common home environment, child-rearing practices (such as activities that foster cognitive skills), siblings' influence on each other, exposure to the same community norms, school experiences and partially shared genetic inheritance. Although no data set can accurately observe all these shared factors, the general measure of sibling similarity captures them. In this study, we aimed to provide a broader picture of the educational reproduction within

families and across generations by combining these two approaches, namely assessing both sibling correlation in education and the contribution that observed paternal and maternal education have in accounting for these similarities.

In contemporary research literature, sibling correlations in education have been under considerable scrutiny (for example, Björklund and Jäntti, 2020; Thaning and Hällsten, 2020; Grätz et al, 2021; Hällsten and Thaning, 2021). However, the evidence regarding cohort changes in these correlations is limited and cover mostly pre-1970s cohorts (for an exception, see Wiborg and Hansen, 2018). Moreover, findings related to trends in reproduction of education in different studies that adopt differing measurement strategies are inconsistent. In this study, we aimed to gain insight into these gaps in knowledge by examining the trends in intergenerational associations and sibling correlations simultaneously.

In addition, the research on social stratification and mobility has been criticised for having a male-centred research focus (Sørensen, 1994; Bihagen, 2008; Beller, 2009), and social stratification among women and gender-specific pathways that drive educational reproduction may still be insufficiently studied (for example, Karhula and Sirniö, 2019; Björklund and Jäntti, 2020). Potentially differing contributions of paternal and maternal education on offspring education may have been concealed by considering only one parent's education at a time. For example, it may be that same-gender siblings and parent-offspring pairs resemble each other more closely than opposite-gender family members. In addition, societal changes, such as strong educational expansion among women (DiPrete and Buchmann, 2013; Van Bavel et al, 2018), may have led to changes in gender-specific patterns of educational reproduction. In this study, we addressed the questions of possible gender-specific differences in sibling and intergenerational resemblance in education and their changes in detail.

Taken together, this study addressed three research aims. First, we assessed the level of sibling similarity in education in Finland. Second, we estimated the contribution of paternal and maternal education to these sibling similarities. Third, we assessed potential gender-specific associations by analysing same-gender sibling groups. For each of these questions, we both established the level of overall association during the study cohort and focused on changes over time in cohorts born between 1950 and 1989. The register-based data sample, which covers the entire population residing in Finland, allows representative and reliable estimation for these effects.

## **The effect of family background on education**

### *Intergenerational reproduction of education*

Among the many social characteristics that run in the family, education is among the most strongly persistent. Sibling correlations are generally stronger in education than in other socio-economic measures, such as occupation and income, and exceed correlations in height in the US context (Conley, 2008; Mazumder, 2008; Björklund and Jäntti, 2020). There are various theoretical reasons to suggest that family and community background matter in educational achievement, including factors related to family resources, the institutional context, and psychological and biological mechanisms.

First, there are numerous potential mechanisms stemming from the socio-economic characteristics of the childhood family (Breen and Jonsson, 2005; Broer et al, 2019).

More plentiful economic resources typical in highly educated families help cover the financial burden related to obtaining education, allow more stable and predictable living conditions, and ensure better access to information, other resources, and housing in neighbourhoods with typically better schools. Rearing practices and scholarly culture of better-educated families, such as having more books and more frequent reading to children, may provide skills facilitating educational success (Park, 2008; Evans et al, 2014). Educational institutions are also more familiar to children with highly educated parents, which helps in terms of making informed decisions, navigating through the school curriculum and communicating with teachers. Middle-class children may also be socialised into having higher educational aspirations (Sewell et al, 1969; Davis-Kean, 2005), for example due to having more role models with higher education and having larger direct pressure for success in schools. Moreover, investing in education is arguably a riskier proposition for people from a disadvantaged background relative to the expected gains (Breen and Goldthorpe, 1997; Goldthorpe 2016: 118–21).

Second, factors related to educational institutions affect the importance of family background in educational attainment. Perhaps the best-documented institutional feature is the tracking and dead ends of educational paths, which strengthen intergenerational persistence, particularly when adopted at early ages (Pfeffer, 2008; Pekkarinen et al, 2009; van de Werfhorst, 2018). Countries with more standardised curricula (for example, relying on country-wide test scores) also typically have lower stratification of educational attainment by socio-economic background (van de Werfhorst and Mijs, 2010).

Third, according to analyses in the field of behavioural genetics, intergenerational persistence is likely to reflect the influence of approximately equally strong genetic and shared environmental components (Branigan et al, 2013; Freese and Jao, 2017; Silventoinen et al, 2004; 2020). The biologically transmitted component is likely to work through psychological mechanisms (such as cognitive ability and personality characteristics) (Rimfeld et al, 2016; Bourne et al, 2018; Israel et al, 2019) and physiological mechanisms (such as anthropometric, metabolic and health-related traits) (Silventoinen et al, 2004; Bulik-Sullivan et al, 2015; Mikkonen et al, 2018).

These considerations make us expect substantial within-family persistence in both sibling similarity and intergenerational associations of education. However, in the specific context of Finland, previous evidence suggests that this resemblance may be lower than in most high-income countries (Pfeffer, 2008; Hertz et al, 2008; Grätz et al, 2021).

### *Trends*

The overall improvement in living standards during the 20th century, such as the declining prevalence of poor housing conditions and nutrition, and increasing leisure time, may contribute to educational mobility by removing barriers of education for less-advantaged socio-economic groups. Furthermore, improvements in welfare-state provision, including longer compulsory schooling, better school support and cost subvention, and decrease of tracking, may have reduced overall selectivity into higher education (Breen et al, 2009; 2010; Pekkarinen et al, 2009).<sup>1</sup> However, many of these macro trends have shown signs of saturation and slowed down or even changed direction in recent decades, coinciding with the stall of welfare-

state expansion (Swank, 2005). Recent decades have also experienced increases in economic inequality, which has an inverse relationship with intergenerational mobility, particularly regarding income and earnings (Corak, 2013; DiPrete, 2020).

Empirical evidence indicates diverging trends in educational reproduction depending on the measurement strategy used. Studies based on sibling correlations show rather ambiguous trends. They do not indicate strong changes over time in central European countries and Sweden among cohorts born before the 1970s (Toka and Dronkers, 1996; Sieben et al, 2001; Björklund et al, 2009). In Norway, one study revealed stability in sibling correlation of education for earlier cohorts in 1946–56 (Raum et al, 2006), whereas Wiborg and Hansen (2018) observed some decline in sibling correlations in more recent cohorts (1960–80).

In European countries, the association between an individual's occupational social-class background and achieved education generally points towards a decline in this relationship for cohorts born in the 20th century. However, there are signs of slowing or stalling of this decline in the latter half of the century (Breen and Jonsson, 2007; Breen et al, 2009; 2010; Erola, 2009; Bernardi and Ballarino, 2016; Barone and Ruggera, 2018; Breen, 2019). In Finland, a decline in the intergenerational odds of following one's parents in completing a higher-educational degree has been observed (Kivinen et al, 2001; 2007; Karhunen and Uusitalo, 2017; Thomsen et al, 2017). Hertz and colleagues (2008) and Pfeffer (2008) observed some decline of the intergenerational associations between the 1940s and turn of the 1980s, but this trend depended on the statistics used (clearer decline on regression coefficients than correlation or log multiplicative unidiff parameters). However, analysis of more recent (post-1960s) cohorts using full educational distribution and strong register-based data revealed evidence of strengthening of the intergenerational relationship (Karhunen and Uusitalo, 2017; Härkönen and Sirniö, 2020).

In summary, societal and institutional changes point towards the possibility of hypothesising various patterns in educational reproduction over time. For studies assessing sibling similarity, empirical evidence has been relatively limited. Perhaps the most comparable previous study regarding social context and cohorts studied is from Norway (Wiborg and Hansen, 2018), where some decline in sibling correlation in education was observed. However, most other studies revealed stability. For parent–offspring associations in education, evidence is more plentiful and even more mixed. Perhaps the most closely comparable studies to ours (Karhunen and Uusitalo, 2017; Härkönen and Sirniö, 2020) provide evidence to expect strengthening of intergenerational reproduction of education after the cohorts born in the 1960s.

### *Gender effects*

Another aim of this study concerned possible gender-specific pathways in within-family similarities, namely whether father–son, mother–daughter and same-gender sibships show stronger associations in education than opposite-gender pairs. According to the social-cognitive theory of gender development (Bussey and Bandura, 1999), individuals actively construct their behaviour considering the cues and sanctions received from surrounding society. Moreover, as gender-related norms tend to imply strong normative expectations, they may have particular significance in shaping behaviour, including educational decisions. Hence, same-gender family members are possibly more influential role models and peers than their opposite-gender

counterparts in the early life and adolescent socialisation process. Among other things, this may be reflected in the stronger same-gender parent–child and sibling–sibling associations. Many previous studies, but not all, confirm such expectation of stronger same-gender intergenerational association in education (Fessler and Schneebaum, 2012; Schneebaum et al, 2015 Minello and Blossfeld, 2017). In a similar vein, higher sibling correlations in same-gender than in opposite-gender sibships have typically been observed (Conley and Glauber, 2008; Mazumder, 2008; Björklund and Salvanes, 2011; Wiborg and Hansen, 2018).

We also assessed the relative strength of the same-gender associations for men and women. There are grounds to hypothesise either larger male, female, or similar associations. On one hand, it has been argued that the gender roles are more rigid among men than among women (Bussey and Bandura, 1999; DiPrete and Buchmann, 2013). There is also evidence of greater sensitivity of boys than girls to educational expectations and available school resources (Legewie and DiPrete, 2012). This could manifest in stronger same-gender associations between men than between women. On the other hand, educational expansion has been weaker among men than among women. The educational level of women was clearly lower among cohorts born in the early 20th century, but approached the level among men around the middle of the century and has since exceeded it in most developed countries (Hout and DiPrete, 2006; Pekkarinen, 2012; DiPrete and Buchmann, 2013; Van Bavel, 2018). There have been societal obstacles hindering the educational attainment of women, such as social norms and values preferring the education of male offspring or lower return of education due to more limited labour-market opportunities and more responsibilities in (unpaid) domestic work and childcare. Some of these obstacles have gradually and partially eroded. Taken together, these societal changes open up the chance to observe stronger similarity in education among female family members. Finally, a hypothesis of gender similarity is also plausible, as the overall mechanisms of intergenerational transmission are expected to be rather similar for both men and women (Breen et al, 2010); this is an expectation consistent with overall ambiguous results in gender-specific intergenerational transmission in the psychological literature (Russell and Saebel, 1997).

Consistent with theoretical expectations, empirical results regarding whether the associations differ between male and female offspring are also mixed. Kleinjans (2010) observed somewhat stronger effects of parental education on offspring's educational aspirations among mother–daughter than father–son pairs. Slightly higher sibling correlations have been observed among sisters than among brothers in some studies (Raum et al, 2006; Conley and Glauber, 2008; Björklund and Salvanes, 2011; Bredtmann and Smith, 2018; Wiborg and Hansen, 2018), although others estimate higher correlations among brothers than sisters (Björklund and Jäntti, 2012; Schnitzlein, 2014). In particular, little evidence is currently available to assess the gender-specific changes in sibling similarity. However, Wiborg and Hansen (2018) observed a more consistent decline in sister than in brother correlations.

To summarise, a stronger resemblance among same-gender than opposite-gender family members in both sibling and intergenerational associations in education may be expected. There are also reasons to expect an increase of the contribution of mothers in explaining this similarity across cohorts. However, given the inconsistent evidence, we do not put forward firm expectations on whether male or female family pairs show stronger resemblance.



### *The Finnish context*

Finland has an extensive social-democratic welfare state, with policies such as wide income redistribution and free-of-charge education on all levels (Kvist et al, 2012; Kalalahti and Varjo, 2020). Dead ends for progression have been avoided in educational institutions, and the level of tracking is comparatively low (Pfeffer, 2008; Bol and van de Werfhorst, 2013). The comprehensive school reform that was gradually implemented between cohorts born between 1961 and 1966 (see Kalalahti and Varjo, 2020; Pekkala Kerr et al, 2013) reduced tracking in the educational system. Socio-economic inequality is generally rated as modest in international comparisons. High social mobility prevails, specifically in terms of the low intergenerational persistence in education in international comparisons (Hertz et al, 2008; Pfeffer, 2008; Bol and van de Werfhorst, 2013).

A strong dual-earner ideal prevails in Finnish households, demonstrated in the small gender gap in labour-force participation and high full-time employment among women, for example (Jaumotte, 2003). Nevertheless, there is a high level of gender segregation both in labour markets and in educational fields (Bettio and Verashchagina, 2009; Charles and Bradley, 2009; Stoet and Geary, 2018). As in other high-income countries, there was considerable educational expansion in Finland during the 20th century, although this stalled among cohorts born in the 1970s onwards. The expansion was stronger among women than among men (Pekkarinen, 2012). Girls also tend to have better Programme for International Student Assessment (PISA) test scores than boys, and the gap in Finland is relatively large in cross-country comparison (OECD, 2019).

## **Materials and methods**

### *Sample and variables*

Data were derived from the population registers maintained by Statistics Finland, which covers the full sample of permanent residents in Finland. Measures were made at five-year intervals between 1970 and 1985 and yearly between 1987 and 2020. We included families with at least two children born between 1947 and 1992. We observed 2,283,120 individuals nested in 908,402 full sibships. The data and individuals were linked using the personal identification codes that have been in use since the 1960s, which allow highly reliable linkage. We analysed anonymised data via the secure remote access system FIONA, maintained by Statistics Finland.

We used a rolling measurement in the trend analysis, including an index person born in a given year (1950–89) and siblings born within three years either side. For example, the analysis of the 1960 cohort contains sibships with at least one individual born in 1960 and their siblings born between 1957 and 1963. The advantage of rolling measurement is that it allows for adequately sized and representative samples. The advantages of this strategy relative to fixed cohorts is to avoid arbitrary cut-off decisions, and that the likelihood of an individual to be included is not dependent on being born in the middle of the interval (as those individuals are more likely to have siblings born within it). A limitation of rolling measurement is that if there were any sharp discontinuities in the trends, their timing could not be precisely identified. The number of observations for each rolling cohort varied between 58,592 and 122,538 individuals nested in 26,700–47,851 sibships in the trend analyses, whereas in the



analyses of single-gender sibling groups the number of observations varied between 14,589 and 35,033 individuals and 6,908–15,389 sibships. Although the families of single-gender sibships families may also include opposite-gender siblings, these were not included in the single-gender sibship models.

Education in both generations was measured in years beyond basic education, based on the expected years of education needed for the highest level achieved during the individuals' lives. We transformed the categories following the guidelines of the International Standard Classification of Education 2011 (UNESCO, 2012) as follows:

- 1 No qualification beyond basic education (ISCED 0–2): 0 years.
- 2 Upper-secondary education (ISCED 3): 3 years.
- 3 Specialised vocational qualification (ISCED 4): 4 years.
- 4 Lowest tertiary degree (ISCED 5): 5 years.
- 5 Bachelor's or equivalent, including basic polytechnic degree (ISCED 6): 6 years.
- 6 Master's level or equivalent (ISCED 7): 8 years.
- 7 Doctoral or licentiate degree (ISCED8): 12 years.

Table A.1 in the Appendix presents the descriptive statistics related to the years of education achieved.

#### *The variance decomposition method*

The analysis is based on two-level linear random intercept models fitted utilising the maximum likelihood method. Level 1 contains individuals and level 2 sibships. In each phase of the analysis, we estimated four different models, specified as follows:

$$edu_{is} = \beta_1 + \zeta_s + \varepsilon_{is} \tag{1}$$

$$edu_{is} = \beta_1 + \beta_2 fatedu_s + \zeta_s + \varepsilon_{is} \tag{2}$$

$$edu_{is} = \beta_1 + \beta_2 motedu_s + \zeta_s + \varepsilon_{is} \tag{3}$$

$$edu_{is} = \beta_1 + \beta_2 fatedu_s + \beta_3 motedu_s + \beta_4 (fatedu * motedu)_s + \zeta_s + \varepsilon_{is} \tag{4}$$

Here,  $edu_{is}$  refers to years of education of individual  $i$  in sibship  $s$ ;  $fatedu_s$  refers to the father's and  $motedu_s$  to the mother's education;  $\zeta_s$  is the sibship-level residual term and  $\varepsilon_{is}$  is the individual-level residual term. We focused on the decomposition of the variances of the residual terms,  $\psi$  (sibship-level residual variance) and  $\theta$  (individual-level residual variance).

Our analysis proceeded in two phases. First, we specified the proportion of the total residual variance that is accounted for by the sibship-level residual variance. This statistic is the intra-class correlation, or sibling correlation in this specific design. Sibling correlations ( $\rho$ ) were estimated in the main analysis from the empty model (equation 1). The number within parentheses in the following equations refers to the model number as defined above:

$$\rho = \frac{\psi_{(1)}}{\psi_{(1)} + \theta_{(1)}} \tag{5}$$

We next calculated how large a share of the family-level variance observed in the empty model (equation 1) was attributable to the observed parental education variables in the model. We assessed both the total contribution of both parents' education and the direct contribution of the mother's and the father's education independently of the other parent's education. These statistics are sometimes called level-2 coefficients of determination,  $R_2^2$  (Raudenbush and Bryk, 2002: 74; Lahtinen et al, 2019a):

$$R_{2, \text{mother \& father}}^2 = \frac{\psi_{(1)} - \psi_{(4)}}{\psi_{(1)}} \tag{6}$$

$$R_{2, \text{direct\_father}}^2 = \frac{\psi_{(3)} - \psi_{(4)}}{\psi_{(1)}} \tag{7}$$

$$R_{2, \text{direct\_mother}}^2 = \frac{\psi_{(2)} - \psi_{(4)}}{\psi_{(1)}} \tag{8}$$

The numbers within parentheses in subscripts of  $\psi$  in equations 6–8 again refer to the model numbers 1–4. Total contribution ( $R_{2, \text{mother \& father}}^2$ ) is the proportion of the sibship-level residual variance in the empty model (1) that was accounted for in the full model (4). Thus,  $R_2^2$  is analogous to  $R^2$  in an ordinary least squares model, but on sibship-level variance instead of overall residual variance. The direct contribution of a parent, in turn, is based on the difference between the full model and the model *without* the information of the parent in question (relative to the level-2 variance of the empty model). Thus, the interpretation of this direct contribution is the proportion of the variance explained that is independent from the contribution of the other parent.

Using models (1)–(4) as a basis, we first calculated sibling correlations ( $\rho$ ) and level-2 coefficients of determination ( $R_2^2$ ) for overall sample. Second, we performed the same analysis on same-gender sibships separately. Third, we estimated the models for each cohort separately and fourth, by each cohort and same-gender sibships separately. The 95% confidence intervals (95% CIs) were obtained by means of nonparametric bootstrapping (percentile method with 2,000 replications).

## Results

The sibling correlations in Table 1 (for corresponding regression tables, see Table A.2 in the Appendix) show that across all the studied cohorts, 37.1% (95% CI 36.9, 37.3;  $\rho$ ) of the total variance in education was shared between siblings. Of this percentage, the education of both parents explained 38.7% (95% CI 38.4, 39.0;  $R_2^2$ , mother&father). The direct contribution of the father's education (0.099; 95% CI 0.097, 0.100;  $R_2^2$ , father) was slightly more pronounced than that of the mother's education (0.076; 95% CI 0.075, 0.078;  $R_2^2$ , mother). The total contribution of parental education was 2.2 times ( $= 0.387 \div (0.076 + 0.099)$ ) as large as the combined direct contributions of the mother and the father, which means that approximately half of the contribution of parental education was shared between the parents. This reflects the relatively strong association between educations of spouses in Finland (Mäenpää and Jalovaara, 2015).

For single-gender sibships, we observed that the sibling correlations were slightly higher than among all sibships (0.397 for groups of brothers, 95% CI 0.394, 0.400; 0.408 for sisters, 95% CI 0.405, 0.411). In turn, the contribution of parental

**Table 1:** Sibling correlations of education and the proportion of shared variance between siblings attributable to parental education for those born between 1950 and 1988

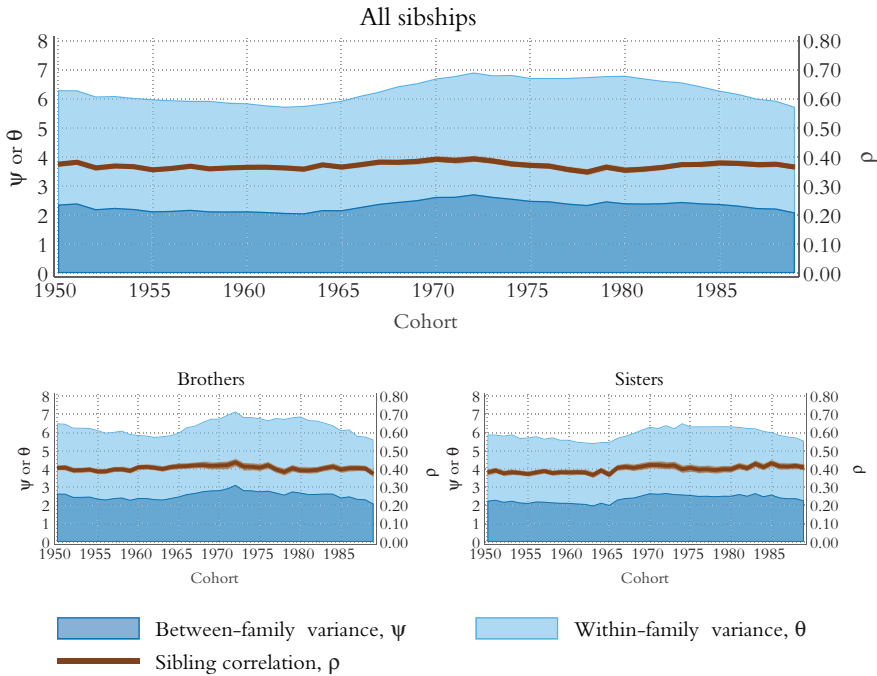
	All sibships			Brothers			Sisters		
	estimate	95% lci	95% uci	estimate	95% lci	95% uci	estimate	95% lci	95% uci
$\rho$	0.371	0.369	0.373	0.397	0.394	0.400	0.408	0.405	0.411
$R^2_2$ , direct father	0.099	0.097	0.100	0.110	0.107	0.113	0.074	0.072	0.077
$R^2_2$ , direct mother	0.076	0.075	0.078	0.056	0.054	0.057	0.091	0.089	0.094
$R^2_2$ , mother & father	0.387	0.384	0.390	0.374	0.370	0.378	0.373	0.369	0.377
N, individuals	2,283,120			810,077			750,723		
N, sibships	908,402			368,243			340,376		

education in explaining this similarity was marginally lower (0.374 for brothers, 95% CI 0.370, 0.378; 0.373 for sisters, 95% CI 0.369, 0.377) than in the case of all sibships. No difference was observed between sisters and brothers in either their overall similarity or in the contribution of the education of both parents. However, the direct contributions of the mother's and the father's education differed. Among groups of brothers, the direct contribution of paternal education (0.110; 95% CI 0.107, 0.113) was substantially higher relative to maternal education (0.056; 95% CI 0.054, 0.057). In contrast, among groups of sisters, the direct contribution of maternal education was slightly stronger relative to paternal education (0.091; 95% CI 0.089, 0.094 vs. 0.074; 95% CI 0.072, 0.077).

After establishing these patterns among the overall study population, we turned our attention to changes across cohorts. Figure 1 shows the variance components ( $\psi$  and  $\theta$ ) as well as the sibling correlations ( $\rho = \psi / (\psi + \theta)$ ) by each cohort between 1950 and 1989. Although the overall educational level of the population substantially increased across these cohorts and the overall variance in education fluctuated as well, sibling correlation was stable, varying between 0.348 (95% CI 0.338; 0.357, cohort 1978) and 0.393 (95% CI 0.383; 0.403, cohort 1972). Similar persistence, or trendless fluctuation, prevailed in the sibling correlations of single-gender sibships, which varied between 0.374 (95% CI 0.356; 0.391, cohort 1989) and 0.437 (95% CI 0.415; 0.456, cohort 1972) for brothers, and between 0.367 (95% CI 0.351; 0.382, cohort 1963) and 0.432 (95% CI 0.415; 0.448, cohort 1985) for sisters.

Contrary to the persistent sibling correlations, Figure 2 shows consistent trends in the contribution of parental education in explaining the shared variance of education between siblings. Among cohorts born in the 1950s, parental education explained approximately 30% of the variance shared by siblings, whereas among those born after the early 1970s this proportion was 40%. This change was largely attributable to the steady and almost monotonic increase in the direct contribution of maternal education. This more than doubled across the cohorts, from 5.0% (95% CI 4.4; 5.6) for those born in 1950 to 12.8% (95% CI 11.9; 13.8) for those born in 1989. The direct contribution of paternal education net of maternal education was rather stable

**Figure 1:** Trends in variance components and sibling correlations of education from empty models, cohorts 1950–89



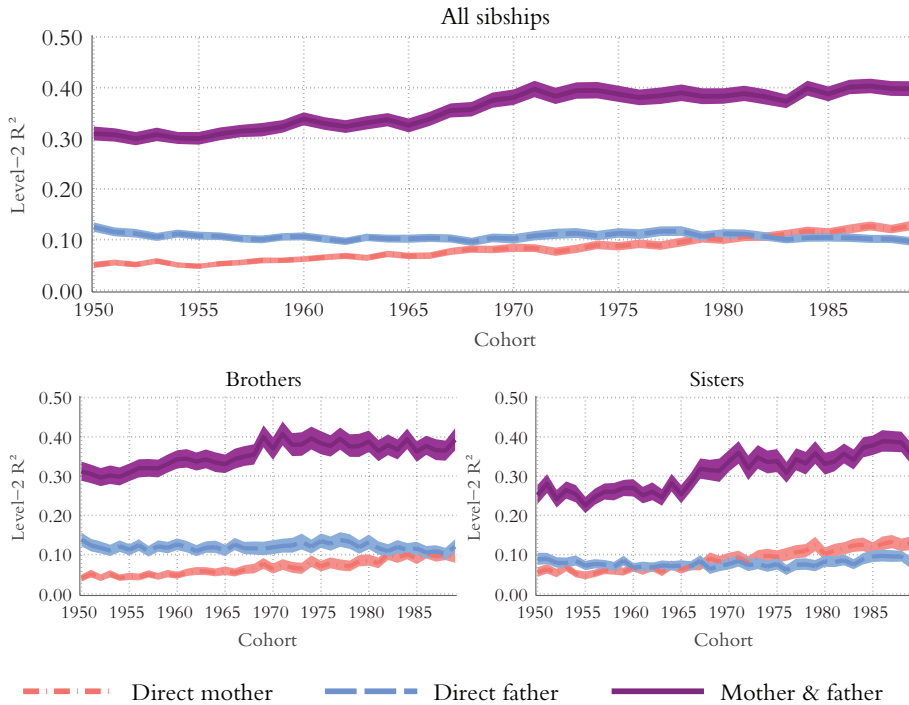
Shaded area around  $\rho$ : 95% confidence interval

at approximately 10% across all cohorts, fluctuating between 9.5% (95% CI 8.7, 10.4, cohort 1968) and 12.5% (95% CI 11.5, 13.4 cohort 1950).

The explanatory power of maternal education equalled that of paternal education for the cohorts born at the turn of the 1980s among all sibships, and then surpassed it. Among sisters, this turn happened for cohorts born in the late 1960s. Paternal education was at least as important as maternal education among brothers until the end of our follow-up but the differences converged among cohorts born in the 1980s as well.

To provide robustness checks and further insights into the intergenerational transmission of education, three sets of additional analyses are presented in the Appendix. First, [Table A.3](#) and [Figure A.1](#) show analyses based on regression models that include controls related to family background that siblings do not necessarily share. These include sibship size, age difference relative to the oldest sibling, birth order and gender, and correspond to those used by [Wiborg and Hansen \(2018\)](#). Sibling correlations were marginally higher in unadjusted models among the early cohorts but were higher in the adjusted models among the latter cohorts. Nevertheless, the results are robust to the inclusion of these variables, as the differences between adjusted and unadjusted sibling correlations were very small ( $\pm 0.016$  to each other in each cohort).

In the second set of additional analyses ([Table A.4](#) and [Figure A.2](#)), we measured parental education by dominance approach, namely using the education of the parent with the higher education level. [Figure A.2](#) shows that among the early cohorts, the dominance method explained sibling similarity in education nearly as well as using

**Figure 2:** Trends in the contribution of parental education in explaining the shared variance between siblings, cohorts 1950–89

Shaded areas are 95% confidence intervals

the education of both parents. However, the advantage of using both maternal and paternal education strengthened steadily over time. The advantage of using both the father's and mother's education relative to the dominance method in explaining the between-family variance in education was only 0.012 (1950 cohort) compared with 0.028 (1960 cohort), 0.037 (1970 cohort) and 0.055 (1989 cohort).

In the third set of additional analyses (Table A.5 and Figure A.3 in the Appendix), we relaxed the linearity assumption and measured parental and offspring's education in seven categories using random intercept ordered logit models. Although this analysis produced slightly higher sibling correlation than linear analysis (0.403, see Table A.5), the cohort trends were very similar to the main analysis regarding all parameters of interest (Figure A.3). However, the methodology behind the estimates is less well established in non-linear than linear models. In particular, intra-class correlations in non-linear models have known limitations in their validity and comparability (Breen and Ermisch, 2021). Therefore, caution should be exercised when interpreting these estimates.

## Discussion

### Main results

In this study, we estimated the level and changes between cohorts born in 1950 and 1989 in educational reproduction from various perspectives in Finland. We addressed

three research questions. The first concerned the overall level and changes of sibling similarity. The proportion of variance of years of education in the population that was shared between siblings across all studied cohorts was 0.37. Although indicating substantial family-related influence, this estimate is low compared to many other contexts, such as 0.6–0.7 in the US (Conley and Glauber, 2008; Mazumder, 2008), 0.5–0.7 in Germany (Schnitzlein, 2014; Grätz, 2018; Grätz et al, 2021), 0.54 in France (Boutchenik et al, 2015) or 0.4–0.5 in Norway (Wiborg and Hansen, 2018; Grätz et al, 2021). In contrast, approximately similar sibling correlations have been observed in Denmark (0.38) (Bredtmann and Smith, 2018; Grätz et al, 2021) and in Sweden (0.36–0.44) (Björklund and Jäntti, 2012; Hällsten and Thaning, 2021). The recent Finnish estimate of 0.36 (Grätz et al, 2021) was almost identical to ours. This observation triangulates evidence from previous comparative analyses of parent–child associations that place Finland among the countries with the most educational mobility (Hertz et al, 2008; Pfeffer, 2008).

Moreover, we observed minimal change in sibling similarity among cohorts born between 1950 and 1989. The stability is remarkable considering that Finnish society has experienced many changes, including substantial welfare-state expansion (Esping-Andersen and Korpi, 1986), which has been reflected in changes in the educational system and its secular expansion (Thomsen et al, 2017; Kalalahti and Varjo, 2020). There has been a decline in sibling correlations with education over corresponding cohorts in Norway (Wiborg and Hansen, 2018). We conducted additional analyses with models that used similar controls, which did not explain the discrepancy with the Norwegian findings. A task for future studies would be to disentangle the reasons for the diverging trends in these rather similar countries both that both share a Nordic welfare and cultural context. However, previous studies covering sibling correlations among earlier cohorts in Sweden, Germany and the Netherlands did not reveal strong evidence on trends (Sieben et al, 2001; Björklund et al, 2009).

The second research question concerned the extent to which parental education explained such sibling similarity in education. This was 39%, which places this result in the middle of previous Swedish estimates of 30–46% of intra-class correlation explained (Thaning and Hällsten, 2020; Hällsten and Thaning, 2021).<sup>2</sup> In contrast to the stability in the sibling correlations, there was also a steady increase in the extent to which parental education accounted for this sibling similarity. This proportion increased from approximately 30% among cohorts born in the 1950s to 40% among those born after the early 1970s. This observation is consistent with the observations of recent studies analysing the intergenerational associations of education by Härkönen and Sirniö (2020) and Karhunen and Uusitalo (2017). In our contribution, we found that such an increase may be attributed primarily to the growing direct contribution of maternal education, which more than doubled across studied cohorts, and secondarily to the shared parental contribution. In contrast, the direct contribution of the father's education was stable over time. A corresponding increase of the maternal and stability in paternal contribution was also observed by Beller (2009) in her analysis of social-class mobility in the US context.

The observation of divergent findings according to the measurement strategy used, namely stable sibling correlations but increasing intergenerational association, is indeed somewhat troubling. Even with powerful and reliable data, there is no straightforward answer to this superficially simple descriptive question. The differing results in the previous literature probably reflect in part the statistical approaches

used. Odds ratios and regression coefficients are sensitive to educational expansion, making those with higher degrees less exceptional, whereas correlations, contingency tables and sequential models are less sensitive to changes in distributions (Hertz et al, 2008; Pfeffer, 2008; Thomsen et al, 2017; Härkönen and Sirniö, 2020). However, diverging findings in our analysis do not reduce into pure technicalities, as both intergenerational and sibling associations relate closely to correlation.<sup>3</sup> The different approaches and measures reveal the picture of educational reproduction from different angles. For example, consider a couple with both spouses born in the early 20th century into working-class families who both showed talent in school. However, because of material hardship in the family and a lack of peers and adults with substantial schooling or influential social networks, they did not realise their educational skills and aspirations. The offspring of this couple did not experience such constraints, given the overall improvement in living standards and the expansion of education, and obtained high educational qualifications. Such a scenario shows high sibling resemblance, and low intergenerational resemblance. It is open to interpretation whether this is a case of weak or strong family-background effects in education.

Something analogous to that example may be behind the increasing contribution of maternal education. There have been societal obstacles to education among women as well as low returns, discouraging participation of women in education. These obstacles may have gradually diminished over time (for example, DiPrete and Buchmann, 2013), resulting in an increasingly meritocratic society in which women have better chances to fulfil their educational ambitions and talent. Although many of these societal obstacles for women have not disappeared, such as lower average earnings (Boschini and Gunnarsson, 2018) or higher responsibility for domestic work and childcare (Fahlén, 2016; Van Bavel et al, 2018; García-Mainar et al, 2011), the mother-offspring association in education has surpassed the corresponding father-offspring association. Some of the persistent factors hindering labour-market returns do not necessarily strongly hinder educational attainment (the average age of childbearing has risen beyond typical years of completing education, for example) but may continue to contribute to offspring education.

The third research question concerned the gender-specific pathways in this transmission. The sibling correlations were about one tenth larger among same-gender relative to all sibships; 40–41% of the variance in years of education was shared within same-gender sibships relative to 37% in all sibships. However, there were no marked differences between groups of sisters and brothers, in contrast to some previous results (for higher sister correlations, see Raaum et al, 2006; Conley and Glauber, 2008; Björklund and Salvanes, 2011; Bredtmann and Smith, 2018; Wiborg and Hansen, 2018; for higher brother correlations, see Björklund and Jäntti, 2012; Schnitzlein, 2014). We observed pronounced same-gender transmission in the intergenerational associations, which was strongest between fathers and sons. We also observed a slightly more pronounced mother's contribution among daughters. Maternal education mattered more among groups of sisters from the late 1960s onwards and equalled paternal education among groups of brothers in the most recent cohorts of our study. The observation that same-gender intergenerational transmission was stronger among sons than daughters is consistent with findings on the gendered effects of the father's presence on education (Kalil et al, 2016) but contradicts previous findings on the gendered transmission of educational expectations (Kleinjans, 2010).



Several potential pathways could explain the stronger same-gender transmission in education. First, there may be more intense social interactions between same-gender family members. Second, there may be family-specific gender norms, such as whether education is considered a masculine or feminine pursuit in the local culture of a particular family. Third, gender norms in the educational system may reward skills and personality characteristics that are transmitted within a family in a gender-differentiated way. For example, consider a hypothetical case that girls are expected to show high levels of trait A, whereas trait B is the key to success among boys. Even if a mother with high levels of trait A has an equal chance of passing on this trait to her son as to her daughter, it has a similar effect on her daughter's success as on her own, thus leading to a higher mother–daughter than mother–son association. On the other hand, a boy who has inherited a high level of trait B from his mother is likely to be successful in school, but this would not have increased his mother's education, leading to a relatively low mother–son association. Future studies could empirically test the extent to which different mechanisms contribute to gendered transmission.

### *Methodological considerations*

Although our register-based data have various strengths, including statistical power, representativeness and reliable measurement, limitations must also be acknowledged. First, the achieved level of education may further improve after the end of the follow-up in 2020 for some individuals. However, this is likely to have only a minor impact on our results, given that even those in the youngest cohort were 28–34 years old then. Second, we were unable to distinguish between different basic level (ISCED 0–2) educational qualifications in our data. The parents of the earliest study cohorts may have had less than nine years of education, but this measurement bias is likely to be modest, as the mandatory duration of education has been at least six years from 1921 onwards (Statistics Finland, 2007). Third, there were individuals for whom we were unable to identify both parents, particularly among the earlier cohorts, although this missingness was again relatively modest (Figure A.4 in the Appendix).

We also point out that our estimates do not reveal the causal mechanisms involved. Factors not directly accounted for may include family-transmitted health profiles or parenting styles and aspirations (Erola and Moisio, 2007; Liu, 2018; Mikkonen et al, 2018). In addition, although genetic inheritance may also explain persistence in the overall level of intergenerational transmission (Branigan et al, 2013; Liu, 2018), it cannot offer a direct explanation of trends or differences between paternal and maternal contributions. One further task for future studies is to assess the relative contribution and pathways that different socio-economic measures (such as parental occupational social class and income) play in intergenerational transmission of education.

Despite their broader interpretation, sibling correlations have limitations. Furthermore, sibling correlations are conservative estimates of the importance of an individual's family and community background, as the background also includes factors that siblings do not share. These non-shared factors are not captured in the sibling correlation. They may include changing family situations between the various births, differing peer groups or sibling dynamics. Siblings may also be treated differently by parents. For example, if one sibling shows substantial interest in reading and other academic pursuits, parents may actively

foster this behaviour, whereas another, more practically oriented sibling, would be encouraged to follow their vocational ambitions. One way of correcting for this lower-bound estimate would be to include some observed factors that siblings do not share but that relate to their family and community in the analysis. Inclusion of sibship size, relative age, birth order and gender produced virtually identical results. Overall, these variable choices are necessarily rather arbitrary, given that it is equally impossible to observe all the background-related factors that siblings do not share as to observe all the factors they do share. Thus, we resorted to more descriptive and transparent models in our main analysis (see Björklund and Jäntti, 2020, for further argumentation).

## Conclusion

In this study, we observed that sibling correlations in education have been stable across cohorts born between 1950 and 1989, whereas the contribution of parental education was increasing, particularly the contribution of maternal education. The increasing maternal contribution in educational reproduction may stem from increasing possibilities of socio-economic achievement among women, which contrasts with at least implicit assumptions often made that higher mobility implies fairer life chances. This calls for a nuanced analysis of the essentially philosophical question of the optimal amount and type of social mobility (some insightful discussions on the topic exist; see, for example, Jencks and Tach, 2006; Conley, 2008). It also implies that the effort scholars studying social mobility have invested into debating their methodological choices (Erikson and Goldthorpe, 1992; 2002; 2010; Hertz et al, 2008; Björklund and Jäntti, 2020) is warranted.

The increasing explanatory power of maternal achievement also resonates with the old debate on the role of women in the measurement of social mobility. The literature has traditionally relied more on paternal than maternal status (for example, Sørensen, 1994; Beller, 2009; Björklund and Jäntti, 2020). Our results demonstrate that, although such a strategy may have once been justified on empirical grounds, this is no longer the case. Conventional approaches would have concealed an important dimension of social fluidity, namely the strengthening mother–offspring association (compare Beller, 2009). A corresponding observation also holds when education is measured via the dominance method. Among cohorts born in the 1950s, it was almost as effective to use information on the parent with the higher qualification as opposed to information on both parents to explain their offspring's education. However, it gradually became more advantageous to include both parents. Thus, for analysis including recent cohorts, we found support from the thesis by Thaning and Hällsten (2020) that advocated the benefits of using the information from both parents over the dominance method in Sweden. Given their long tradition of dual-earner households and the extensive employment of women, these Nordic contexts also provide a possible scenario for future development among countries in which widespread female participation in the labour force has been more recent.

## Notes

<sup>1</sup> However, there is no straightforward support that tuition costs or educational expansion modify the effect of socio-economic background on education (for example, Denny, 2014; Pfeffer and Hertel, 2015).

<sup>2</sup> However, [Thaning and Hällsten \(2020\)](#) and [Hällsten and Thaning \(2021\)](#) measured the attenuation of sibling correlation, whereas we estimated the level-2 variance explained. These measures are not equivalent, although the difference is likely to be small as the variation in parental education occurs between sibships.

<sup>3</sup> In case of two-sibling families, Pearson correlation of years of education between siblings corresponds to intra-class correlation of our empty models ([McGraw and Wong, 1996](#); [Liu et al, 2016](#)). Changes in  $R_2^2$  statistics in our analyses are also closely related to the changes in intergenerational correlation. Assuming that sibling correlation of the empty model is stable over time and adding parental education into model changes only between-sibship residual variance, an increase of  $R_2^2$  over time implies a proportional increase in overall model  $R^2$  (compare [Rabe-Hesketh and Skrondal, 2012](#): 136), and given that coefficient of determination is (multiple) correlation squared, also an increase in the absolute value of intergenerational correlation.

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### Data availability statement

Authors take responsibility for the integrity of the data and the accuracy of the analysis. The dataset is confidential and permission to use it can be obtained from Statistics Finland (this includes costs).

### Conflict of interest

The authors declare that there is no conflict of interest.

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## Appendix

**Table A.1:** Descriptive statistics: mean years of education beyond basic level and corresponding standard deviations (SD) by cohorts 1950–89

	Men		Women		Fathers		Mothers	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD
1950	2.9	2.6	3.1	2.5	0.8	1.9	0.5	1.4
1951	3.0	2.6	3.3	2.5	0.8	2.0	0.6	1.5
1952	3.1	2.6	3.3	2.4	0.9	2.0	0.6	1.5
1953	3.1	2.6	3.5	2.5	1.0	2.1	0.7	1.6
1954	3.2	2.6	3.5	2.4	1.0	2.1	0.7	1.6
1955	3.3	2.5	3.6	2.4	1.0	2.2	0.8	1.7
1956	3.3	2.5	3.7	2.4	1.1	2.2	0.8	1.7
1957	3.3	2.5	3.7	2.4	1.2	2.2	0.9	1.8
1958	3.4	2.5	3.9	2.4	1.2	2.3	0.9	1.8
1959	3.4	2.5	4.0	2.4	1.3	2.3	1.0	1.9
1960	3.5	2.4	4.1	2.4	1.4	2.4	1.1	1.9
1961	3.6	2.4	4.2	2.3	1.4	2.4	1.1	1.9
1962	3.6	2.4	4.2	2.3	1.5	2.4	1.2	2.0
1963	3.7	2.5	4.3	2.3	1.6	2.5	1.3	2.0
1964	3.7	2.5	4.4	2.4	1.6	2.5	1.4	2.1
1965	3.7	2.5	4.4	2.4	1.7	2.5	1.5	2.1
1966	3.7	2.5	4.5	2.4	1.8	2.6	1.6	2.2
1967	3.8	2.5	4.5	2.4	1.9	2.6	1.7	2.2
1968	3.8	2.6	4.6	2.4	2.0	2.6	1.8	2.3
1969	3.9	2.6	4.6	2.5	2.1	2.7	2.0	2.3
1970	3.9	2.6	4.7	2.5	2.3	2.7	2.1	2.3
1971	4.0	2.6	4.8	2.5	2.4	2.8	2.3	2.4
1972	4.0	2.6	4.9	2.5	2.6	2.8	2.5	2.4
1973	4.1	2.6	5.0	2.5	2.7	2.8	2.6	2.4
1974	4.1	2.6	5.0	2.5	2.8	2.8	2.7	2.4
1975	4.2	2.5	5.1	2.5	2.8	2.8	2.8	2.4
1976	4.2	2.5	5.2	2.4	2.9	2.8	2.9	2.4
1977	4.3	2.6	5.2	2.5	3.0	2.8	3.0	2.4
1978	4.3	2.6	5.2	2.5	3.1	2.7	3.1	2.4
1979	4.2	2.6	5.2	2.5	3.1	2.7	3.2	2.4
1980	4.2	2.6	5.2	2.5	3.2	2.7	3.3	2.4
1981	4.2	2.6	5.1	2.5	3.2	2.7	3.4	2.4
1982	4.2	2.6	5.1	2.5	3.3	2.7	3.5	2.4
1983	4.1	2.6	5.0	2.5	3.3	2.7	3.6	2.3
1984	4.1	2.5	5.0	2.4	3.4	2.7	3.6	2.3
1985	4.1	2.5	5.0	2.4	3.4	2.6	3.7	2.3
1986	4.1	2.5	4.9	2.4	3.5	2.6	3.8	2.3
1987	4.1	2.4	4.9	2.4	3.5	2.6	3.9	2.3
1988	4.0	2.4	4.8	2.4	3.6	2.6	4.0	2.3
1989	3.9	2.3	4.6	2.3	3.6	2.6	4.0	2.3
Overall	3.8	2.6	4.5	2.5	2.2	2.7	2.1	2.4

**Table A.2:** Years of education beyond basic level on parental education: two-level random-intercept linear regression models, cohorts 1950–89

<b>All sibships (N = 2,283,120 individuals; N = 908,402 sibships)</b>								
	<b>Model 1</b>		<b>Model 2</b>		<b>Model 3</b>		<b>Model 4</b>	
	<b>b</b>	<b>std. err.</b>	<b>b</b>	<b>std. err.</b>	<b>b</b>	<b>std. err.</b>	<b>b</b>	<b>std. err.</b>
Father's education				0.318	0.001			0.227
Mother's education						0.338	0.001	0.219
Interaction								-0.004
Constant	4.169	0.002	3.445	0.003	3.425	0.003	3.204	0.003
psi	2.434	0.007	1.676	0.005	1.731	0.006	1.491	0.005
theta	4.126	0.005	4.137	0.005	4.138	0.005	4.140	0.005
rho	0.371	0.001	0.288	0.001	0.295	0.001	0.265	0.001
<b>Brothers (N = 810,077 individuals; n = 368,243 sibships)</b>								
	<b>Model 1</b>		<b>Model 2</b>		<b>Model 3</b>		<b>Model 4</b>	
	<b>b</b>	<b>std. err.</b>	<b>b</b>	<b>std. err.</b>	<b>b</b>	<b>std. err.</b>	<b>b</b>	<b>std. err.</b>
Father's education				0.337	0.001			0.243
Mother's education						0.340	0.001	0.191
Interaction								-0.001
Constant	3.715	0.003	3.007	0.004	3.028	0.004	2.825	0.005
psi	2.585	0.011	1.761	0.009	1.902	0.010	1.617	0.009
theta	3.932	0.008	3.939	0.008	3.939	0.008	3.942	0.008
rho	0.397	0.001	0.309	0.001	0.326	0.001	0.291	0.001
<b>Sisters (N = 750,723 individuals; n = 340,376 sibships)</b>								
	<b>Model 1</b>		<b>Model 2</b>		<b>Model 3</b>		<b>Model 4</b>	
	<b>b</b>	<b>std. err.</b>	<b>b</b>	<b>std. err.</b>	<b>b</b>	<b>std. err.</b>	<b>b</b>	<b>std. err.</b>
Father's education				0.317	0.001			0.223
Mother's education						0.363	0.001	0.269
Interaction								-0.009
Constant	4.408	0.004	3.737	0.004	3.672	0.004	3.463	0.005
psi	2.567	0.011	1.843	0.010	1.799	0.009	1.608	0.009
theta	3.727	0.008	3.728	0.008	3.729	0.008	3.730	0.008
rho	0.408	0.001	0.331	0.001	0.325	0.001	0.301	0.001

**Table A.3:** Years of education beyond basic level on parental education: two-level random-intercept linear regression models controlled by sibship size, age difference relative to oldest sibling, birth order and gender, cohorts 1950–89

All siblings (N: 2,283,120 individuals; N: 908,402 sibships)								
	Model 1		Model 2		Model 3		Model 4	
	b	std. err.	b	std. err.	b	std. err.	b	std. err.
Father's education			0.310	0.001			0.217	0.001
Mother's education					0.329	0.001	0.209	0.001
Interaction							-0.002	0.0003
Sibship size	-0.214	0.002	-0.114	0.002	-0.102	0.002	-0.076	0.002
Age difference to oldest sibling	0.031	0.001	0.042	0.001	0.046	0.001	0.048	0.001
Birth order	-0.107	0.002	-0.126	0.002	-0.133	0.002	-0.137	0.002
Female	0.697	0.003	0.696	0.003	0.696	0.003	0.696	0.003
Constant	3.641	0.006	2.762	0.006	2.721	0.006	2.461	0.006
psi	2.316	0.006	1.626	0.005	1.683	0.005	1.455	0.005
theta	4.013	0.005	4.018	0.005	4.018	0.005	4.019	0.005
rho	0.366	0.001	0.288	0.001	0.295	0.001	0.266	0.001
$R^2_2$ , direct father	0.099							
$R^2_2$ , direct mother	0.074							
$R^2_2$ , mother & father	0.371							

**Table A.4:** Years of education beyond basic level on parental education measured via dominance method (education of parent with higher educational level). Two-level random intercept linear model, cohorts 1950–89

	b	std. err.
Parental education	0.339	0.001
Constant	3.148	0.003
$\psi$	1.548	0.005
$\theta$	4.139	0.005
$\rho$	0.272	0.001
$R^2_2$	0.363	

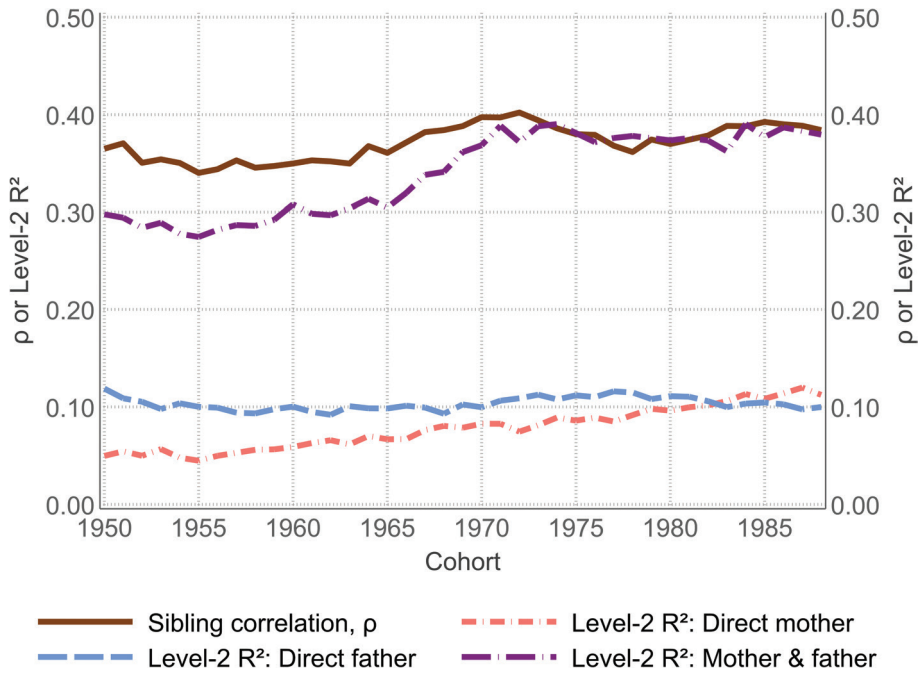
Note: N = 2,283,120 individuals; N = 908,402 sibships.

**Table A.5:** Parameter estimates from two-level random intercept ordered logit models where parental and offspring education has been measured as seven-class categorical variables, cohorts 1950–89

	Estimate
$\rho$	0.403
$R^2_2$ , direct father	0.102
$R^2_2$ , direct mother	0.081
$R^2_2$ , mother&father	0.404

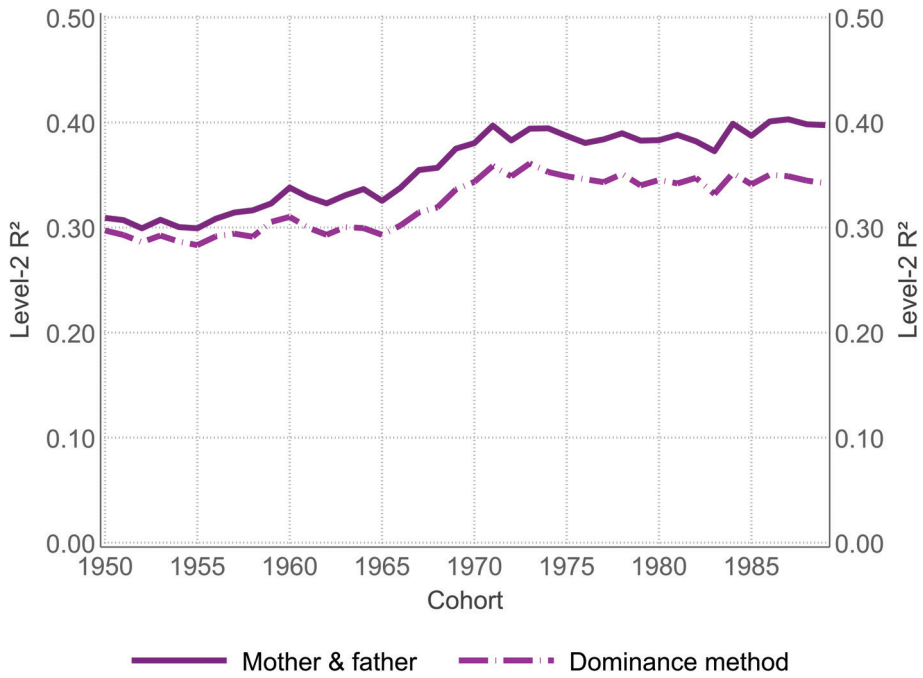
N = 2,283,120 individuals; N = 908,402 sibships.

**Figure A.1:** Trends in the sibling correlation and contribution of parental education in explaining the shared variance between siblings, cohorts 1950–89; models controlled by sibship size, age difference relative to oldest sibling, birth order and gender.



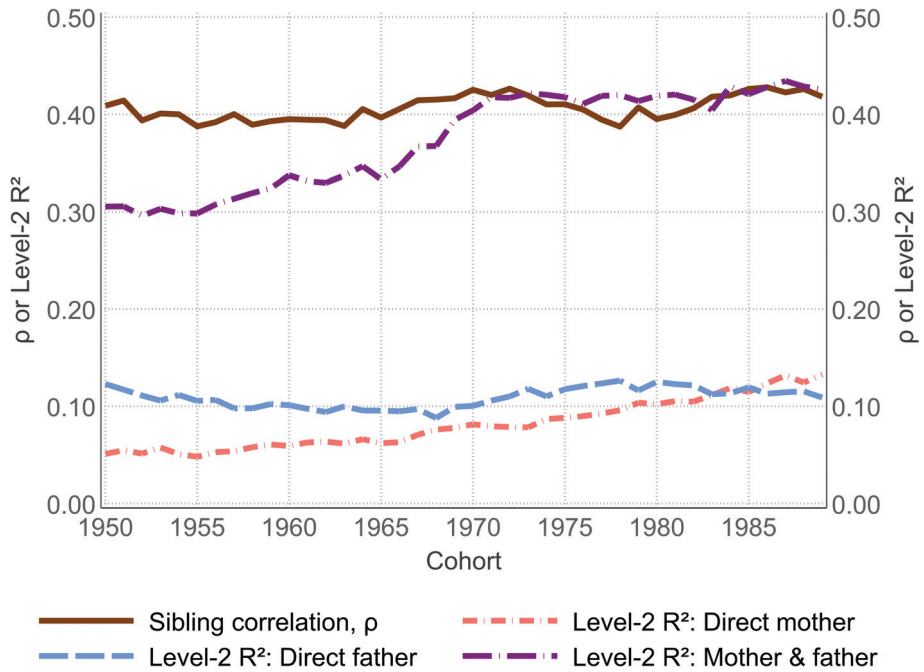


**Figure A.2:** Trends contribution of parental education in explaining the shared variance between siblings; parental education measured via dominance method and education of both parents, cohorts 1950–89.



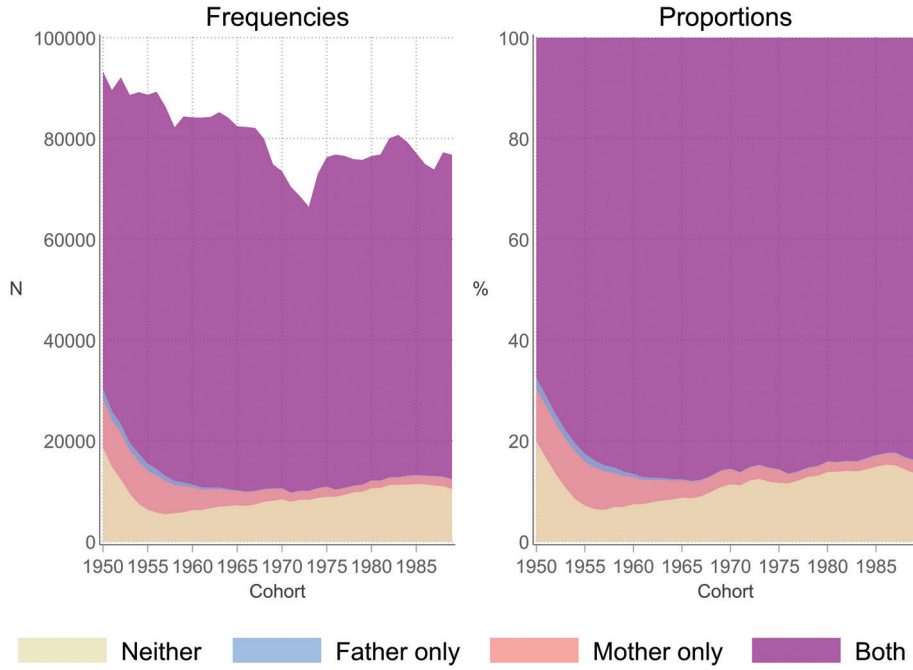
N = 2,283,120 individuals; N = 908,402 sibships.

**Figure A.3:** Trends in the sibling correlation and contribution of parental education in explaining the shared variance between siblings, cohorts 1950–89: Ordered two-level random intercept logit models where parental and offspring education has been measured as seven-class categorical variables.



Note: N = 2,283,120 individuals; N = 908,402 sibships.

**Figure A.4:** Share of the individuals with data on parents available by cohort, cohorts 1950–89.



Note: Unlike in other analyses in this study, restriction of population into those individuals who have siblings was not applied in the Figure A.4, as it is necessary to identify both parents in the data to identify full sibships.